Trade liberalization and the inter-industry wage premia: the missing role of productivity

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Abstract

The literature concerning the effect of tariffs on the inter-industry wage premium has not addressed the role of total factor productivity (TFP) in determining both the wage premium and tariffs. This omission invalidates the use of the pre-reform tariff level as an instrument for the change in tariffs. Based on an analysis of Colombian data, I find that including TFP in the estimated model of the effects of tariffs on the wage premium leads to a 36% decrease in the effect of tariffs on the inter-industry wage premium relative to the model that omits TFP. More specifically, a ten percentage point decrease in tariffs reduces the wage premium by 1.1%. This finding suggests the importance of using policies that boost productivity to offset the effect of tariffs on the wage premium.

JEL codes: F1, F16, J3, O15
Keywords: inter-industry wage premium; trade liberalization; productivity; Colombia

1. Introduction

Recent experience with trade liberalization in developing countries has raised concerns about the impact of trade liberalization on wage inequality. In particular, there have been concerns about changes in labor market outcomes such as returns to skills and the inter-industry wage premium (hereafter called the wage premium), which is the part of the wage that is explained by the characteristics of the worker's industry of affiliation (after controlling for the worker's characteristics). The empirical evidence provided by Gonzaga et al. (2005) for Brazil, Cragg and Epelbaum (1996) for Mexico, Galiani and Sanguinetti (2003) for Argentina, and Attanasio et al. (2004) for Colombia suggests that the observed changes in the skill premium (another way of expressing returns to skills) do not entirely explain the increase in wage inequality.[1] As a result, some researchers started to explore the idea that trade policy may affect wage inequality through other channels such as the wage premium. For instance, changes in the wage premium can result in increased inequality if, as a result of trade policy changes, the wage premium decreases less in skill-intensive industries than in industries that rely primarily on unskilled workers

The wage premium has drawn a lot of attention in this context because the trade liberalization conducted in countries like Colombia, Mexico, and India not only decreased overall protection levels, but also altered protection levels differently across industries. Interestingly, there is mixed evidence concerning the effect of tariffs on the wage premium: no effect for Mexico (Feliciano, 2001), a positive effect for Colombia (Goldberg and Pavcnik, 2005), and a negative effect for India (Mishra and Kumar, 2008).

A weakness of this literature is that it has ignored the role of total factor productivity (TFP) in determining both the wage premium and tariffs. Bartel and Sicherman (1999) found

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[1] Goldberg and Pavcnik (2007) provides a good survey of this literature.

that for the U.S., the higher the industry TFP, the higher the wage premium enjoyed by its workers. In addition, Karacaovali (2011) presented evidence for Colombia that industry TFP and its anticipated future changes are correlated with contemporaneous import tariffs. This means that the instrument used for the change in tariff level (the pre-reform tariff level) is correlated with the current TFP level, which is included in the error term. Hence, the instrument is invalid and the current methodology provides inconsistent estimates of the effects of tariff on the wage premium.

To address this weakness in the current literature, I use the Colombian data set in Goldberg and Pavcnik (2005) to examine whether trade liberalization affects the inter-industry wage premium when the impact of TFP on trade policy and on the wage premium has been taken into account. I also estimate the magnitude of the bias in the Colombian case that is due to their failure to consider this source of endogeneity.

I find that TFP is an important determinant of the wage premium in the estimated model, with a 10% increase in TFP implying a 1.6% increase in the wage premium. The estimated effect of a 10 percentage point decrease in tariffs is a 1.1% reduction in the wage premium, which is about 36% less than the result obtained when TFP is omitted. My estimates also indicate that as an included instrument, TFP has a positive correlation with the change in tariffs, indicating that it is an important time-varying factor in the import tariff setting.

These findings are significant because, according to Goldberg and Pavcnik (2005), industries that have the largest share of unskilled workers suffered the largest tariff cuts, which means that in principle, trade liberalization increased the skilled-unskilled wage inequality through changes in the wage premium. However, my results indicate that there could be a different outcome if the role of productivity were considered. Suppose these same industries presented increases in TFP that were sufficiently large so as to counterbalance the effect of the tariffs, then inequality would be reduced. The implication for policy makers is that they should adopt policies that boost productivity as a tool to reduce the increase in inequality that occurs through the wage premium.

The remainder of the paper is organized as follows. In section 2, I present a stylized model of wage premium setting and discuss the methodological problems of previous studies caused by their omission of TFP. The data used in the empirical analysis are described in section 3. Section 4 presents and discusses the empirical results. Section 5 summarizes the findings and presents conclusions.

2. Methodology

The literature (e.g. Krueger and Summers, 1988) defines the industry wage premium as the share of the worker's wage that is explained by the industry's characteristics after controlling for the worker's observable characteristics. Although there is no consensus, the two most commonly cited reasons for the existence of an industry wage premium are different industry-level productivity levels due to industry-specific inputs, and industry rents that arise from imperfect competition in the output market. These industry rents may be affected in the short and medium term by trade liberalization, as discussed in Goldberg and Pavcnik (2005), and they induce wages to be higher than the market average due to rent sharing, as shown by Blanchflower et al. (1996) and Hildreth and Oswald (1997), who used firm-level data for U.S. labor markets.

2.1 A Simplified Model of the Wage Premium

To assess the effects of trade liberalization on the wage premium, I present a simplified representative firm model in which the wage premium is due to industry rent sharing in a small open economy. I assume that the capital stock is fixed in the short and medium run, thus representing a fixed cost (f_j). This assumption is equivalent to assuming that capital is sector specific for the time horizon considered here, and allows for the existence of positive profits in equilibrium.

Let p_{jt} be the price of the good produced in industry j at time t. It is a function of the import tariff (τ_{jt})[2], in the sense that as the import tariff decreases, the output price faced by the domestic representative firm decreases ($\partial p_{jt}/\partial \tau_{jt} > 0$). The profit maximization problem faced by industry j representative firm in the short-run is

$$\max_{L_{jt}} \Pi_{jt} = \max_{L_{jt}} p_{jt} Y\left(A_{jt}, L_{jt}\right) - \left(w_t + \beta_{jt} \frac{\Pi_{jt}}{L_{jt}}\right) L_{jt} - f_j \tag{1}$$

where $Y(A_{jt}, L_{jt})$ is the production function that depends upon A_{jt}, which is the total factor productivity (TFP), and the number of workers (L_{jt}), which is adjustable over time; w_t is the average wage paid in the economy, which can be seen as the worker outside option; and β_{jt} is the workers' share of industry j rents at time t, which can also be thought of as the workers' bargaining power. So, the industry wage premium received by a worker at industry j at time t is given by $wagepremium_{jt} \equiv \beta_{jt}\Pi_{jt}/L_{jt}$. Let $Y(A_{jt}, L_{jt})=A_{jt}L^{\rho}_{jt}$, and in order to have an interior solution $0<\rho<1$ is needed. Solving the profit maximization problem, and plugging the demand for labor back into equation (1), we obtain the profit function as described by equation (2):

$$\Pi_{jt} = [1 + \beta_{jt}]^{-1} [p_{jt} A_{jt} w_t^{\rho} \rho^{\rho}]^{\frac{1}{1-\rho}} (1 - \rho) - f_j \tag{2}$$

Then, we can see from equation (3) that the wage premium depends on the price of the good produced and on the TFP (A_{jt}):

$$wagepremium_{jt} = \frac{\beta_{jt} w_t^{\frac{1}{1-\rho}}}{1+\beta_{jt}} \left(\rho^{-1} w_t^{\rho} (1 - \rho) - [p_{jt} A_{jt} \rho]^{\frac{-1}{1-\rho}} f_j\right) \tag{3}$$

[2] In a monopolistic competition framework, one can also model the increase in competitiveness induced by trade liberalization as a shock to the price-elasticity of demand for output, i.e. more competitiveness results in a more elastic demand, and as a result, a lower markup.

The level of competitiveness is affected by trade policy through a decrease in import tariffs, which forces a decrease in domestic prices[3], and implies a decrease in industry rents that is shared among the factors of production[4], as shown in equation (4).

$$\frac{\partial wagepremium_{jt}}{\partial p_{jt}}\frac{\partial p_{jt}}{\partial \tau_{jt}} = \left(\frac{1}{1-\rho}\right)\frac{\beta_{jt}}{1+\beta_{jt}} w_t^{\frac{1}{1-\rho}}\left[A_{jt}\rho\right]^{\frac{-1}{1-\rho}} p_{jt}^{\frac{\rho-2}{1-\rho}} f_j \frac{\partial p_{jt}}{\partial \tau_{jt}} > 0 \qquad (4)$$

The result from equation (4) relies on the assumption that firms cannot influence the tariff level, in other words, they cannot lobby government to choose a tariff level that maximize their profits. In the next subsection, I relax this assumption and explore its consequences on the current methodology used to estimate the effect of tariffs on the wage premium.

2.2 Current methodology and its shortcomings

Empirically, this relationship between trade liberalization and the wage premium has been assessed based on the fact that in several countries trade liberalization led to a decrease in tariffs that was not uniform across industries. This changed the protection pattern across industries, which can be used to identify the relationship between the wage premium and import tariffs. Goldberg and Pavcnik (2005) applied this methodology to estimate the effect of tariffs on the wage premium using data from the Colombian trade liberalization episode that started in the 1980s,[5] as shown in equation (5):

$$wagepremium_{jt} = \alpha + \phi\tau_{jt} + \delta_j + \theta_t + error_{jt} \qquad (5)$$

[3] Evidence that a decline in tariffs implies lower markups was found by Harrison (1994) for Cote d'Ivoire, Levinsohn (1993) for Turkey, and Currie and Harrison (1997) for Morocco.
[4] In addition, more competition might change the bargaining power of unions and/or shift their focus from wages to job security, hence decreasing their share of the industry rents and therefore reducing the wage premium. But this channel is subject to the specific wage negotiation framework of each industry and country.
[5] Trade liberalization in Colombia started when President Virgilio Vargas took office in 1986 and was completed by his successor President Cesar Gaviria by 1992. This reform consisted of a large decrease in all tariffs across all industries (generating a large decrease in average tariff) and a removal of almost all non-tariff barriers.

where τ_{jt} is the Colombian import tariff for industry j in year t. The use of industry (δ_j) and year (θ_t) fixed effects makes the within-industry variation the source of identification of ϕ, which suggests the use of clustered standard errors at the industry level.

Changes in tariffs across industries have received a great deal of attention in the literature on the political economy of trade protection. This research provides both a theoretical basis (Grossman and Helpman, 1994, and Karacaovali, 2011) and empirical support (Trefler, 1993, Goldberg and Maggi, 1999, and Karacaovali, 2011) for the idea that some characteristics of the industry (e.g. market concentration, share of unskilled workers) are taken into account when setting tariffs. This means that if some of these characteristics also influence the wage premium, the current tariff level becomes an endogenous variable. For instance, a higher wage premium in a given industry might induce its workers to lobby against trade liberalization. A possible time-variant political economy factor is union activity. Fortunately, this does not appear to be the case, based on anecdotal evidence in Edwards (2001) suggesting almost no union activity in Colombia. Thus, in my theoretical model, I assume $\beta_{jt} = \beta_j$. Time-invariant industry characteristics that affect tariff setting can be controlled by adding industry fixed effects (δ_j) to the estimated model.

As in many developing countries, Colombia's trade reforms were accompanied by other reforms.[6] There was a labor market reform in 1990 that reduced dismissal costs and introduced a new severance payment system. Banking system reforms in the early 1990s were aimed at liberalizing deposit rates, extinguishing credit subsidies, and modernizing capital markets. Moreover, foreign direct investment restrictions were greatly reduced in 1991, when multinational companies were basically given the same treatment as Colombian firms. Finally, in 1993, there was a major change in the social security system. These reforms, which affected all industries equally, underline the importance of including year

[6] See Kugler (1999) and Eslava et al. (2004) for details about Colombia's economic reforms.

effects (θ_t) in the empirical specification, which would also account for business cycle effects on both tariff and wage premium setting. It is possible that the reforms affected industries differently, in which case the industry effects will capture this variation.

Nevertheless year and industry effects do not account for industry-level variables that change over time and that affect tariffs and the wage premium simultaneously. Goldberg and Pavcnik (2005) recognized this shortcoming and suggested addressing it by estimating the first difference of equation (5) with two-stage least squares using instruments for the change in tariffs.

The instruments they chose are the 1983 (pre-reform) tariff level and its interactions with the exchange rate and the coffee price. The rationale for this choice is that the 1983 tariff level happens to be proportional to the tariff cut at the industry level, i.e. the higher the 1983 tariff in a given industry, the larger the magnitude of the decrease in the tariff for this industry under trade liberalization, this negative correlation is clearly depicted in Figure 1. The use of the interactions of tariffs with the real exchange rate and coffee prices (the most important Colombian export product) is motivated by the fact that in designing Colombia's trade reform, trade balance levels were taken into account to prevent a Balance of Payments crisis. More specifically, the idea is that exchange rate devaluations, which induce smaller imports and larger exports, would lead to a larger decrease in tariffs, and increases in coffee prices, which generate larger export revenues, would also lead to a larger decrease in tariffs.

But there is still a problem with the strategy adopted by Goldberg and Pavcnik (2005): it ignores the role of productivity in the determination of both the wage premium and the tariff. This is a problem because an increase in industry productivity (TFP) leads to an increase in the wage premium, as shown in equation (6):[7]

[7] For example, Bartel and Sicherman (1999) found empirical evidence of a positive correlation between TFP and the wage premium in the U.S. In the Colombian manufacturing industries sample used in this paper, the simple

$$\frac{\partial wagepremium_{jt}}{\partial A_{jt}} = \left(\frac{1}{1-\rho}\right)\frac{\beta_j}{1+\beta_j} w_t^{\frac{1}{1-\rho}} [p_{jt}\rho]^{\frac{-1}{1-\rho}} A_{jt}^{\frac{\rho-2}{1-\rho}} f_j > 0 \qquad (6)$$

This fact alone would not pose an econometric problem if the tariff and TFP were not correlated. But this is clearly not the case, since there is a growing literature concerning the effect of trade liberalization on industry and firms' productivity levels. In particular, a decrease in import tariffs was found to increase TFP in Chile (Pavcnik, 2002), Brazil (Ferreira and Rossi, 2003), India (Krishna and Mitra, 1998, and Topalova, 2004), and Colombia (Karacaovali, 2011).

Overall, while on the one hand trade liberalization decreases the rents to be shared and then decreases the wage premium (direct effect), on the other hand trade liberalization may boost productivity, which increases the wage premium (indirect effect). This means that in the current methodology the total effect of trade liberalization on the wage premium is ambiguous. And this could be the reason behind the difference in the sign of the tariff coefficient, for instance a positive effect for Colombia (Goldberg and Pavcnik, 2005) and a negative effect for India (Mishra and Kumar, 2008).

Karacaovali (2011) presents a theoretical model in which industry TFP is an important political economy factor that affects tariff setting. This is because under the plausible assumption of a small price elasticity of supply, the higher the industry TFP, the greater the benefits from protection that accrue to firms in that industry, and therefore the greater the incentives to lobby the government for more protection. Karacaovali (2011) further finds that in the context of a trade liberalization that is similar to the one adopted by Colombia, the extent of industry-level liberalization (the change in tariffs) depends positively upon the change in its productivity, which means that future tariffs may be influenced by anticipated changes in productivity.

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correlation between the natural logarithm of TFP and the wage premium is 0.373, and the correlation between the first-difference of the same variables is 0.45.

On the empirical side, Karacaovali's (2011) estimates using Colombian data support his theoretical predictions. He examines the correlation between tariffs and TFP over time, and finds a statistically significant correlation between future TFP and present tariffs, and that there is autocorrelation in TFP over time.[8]

From the discussion above, it is clear that the bias of the tariff coefficient resulting from the omission of TFP should be positive, due to the positive correlation of the tariff with TFP. Once TFP is included in the estimated model, the estimated coefficient of the tariff should decrease in magnitude. Second, the exclusion restriction for using the 1983 tariff level as an instrument for tariff changes is not met. This is because the pre-reform (1983) tariff level takes into account the industry-level TFP at that time, which in turn is correlated with the current TFP level (which is part of the error term due to the omission of TFP in equation (5)). This means that the instrument (1983 tariff level) and the error term are correlated, making the instrument invalid.

2.3 Incorporating Productivity into the Estimated Model

To address the problems described earlier and obtain consistent estimates of the impact of tariffs on the wage premium, TFP has to be incorporated into the estimated model[9]. This eliminates the problem of correlation between the instrument (1983 tariff level) and the error term. The estimated model then becomes equation (7), where Δ indicates time difference:

$$\Delta wagepremium_{jt} = \phi \Delta \tau_{jt} + \gamma \Delta log(TFP_{jt}) + \Delta \theta_t + \Delta error_{jt} \tag{7}$$

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[8] Using Colombian data, Fernandes (2007) also found that TFP has autocorrelation, with one reason for this behavior being the econometric technique used to estimate TFP.

[9] The introduction of the productivity variable on the right-hand side can generate reverse causation if firms pay efficiency wages (Krueger and Summers, 1988) to avoid shirking, which means they pay wages higher than the worker outside option. This causes the workers' productivity to increase, as a result TFP becomes an endogenous regressor. However, good instruments for TFP are not available to investigate this hypothesis.

3. Data

I use two data sets for the empirical analysis. The first is the Colombian data set in Goldberg and Pavcnik (2005), from which I use the estimated wage premium and its respective estimated standard error, import tariff, real exchange rate, and coffee price variables. The second data set comes from Eslava et al. (2004), and consists of industry-level TFP.

3.1 The Goldberg and Pavcnik (2005) Data Set

To estimate the wage premium, Goldberg and Pavcnik (2005) used worker-level data from the Colombian National Household Survey (*Encuesta Nacional de Hogares*) conducted by the Colombian Statistical Agency (*DANE*) in 1984, 1986, 1988, 1990, 1992, 1994, 1996, and 1998.[10] This survey, which is a repeated cross-section covering urban areas of Colombia, contains questions about demographic characteristics (age, gender, marital status, education, literacy, etc.), job type, industry of employment at the 2-digit ISIC level (total of 33 industries, with nine in the manufacturing sector), and region of residence. Following Krueger and Summers (1988), Goldberg and Pavcnik (2005) estimated the wage premium separately for every year of the sample using the following specification:

$$\log(w_{ijt}) = \mathbf{H}_{ijt}\alpha_H + \sum_j I_{ijt} * wagepremium_{jt} + \varepsilon_{ijt} \qquad (8)$$

where $\log(w_{ijt})$ is the natural logarithm of the hourly wage received by the worker i affiliated with industry j at time t. $\mathbf{H}_{ijt}$ is a vector that contains worker's age, age squared, a male indicator, a married indicator, a head of household indicator, completed elementary, secondary and college education indicators (the omitted category is incomplete elementary

[10] Because the data are for every other year, in this paper the first difference of a variable should be interpreted as the difference between the variable at year t and the variable at year t-2.

education), a literacy indicator, a residence in Bogotá indicator, occupational indicators, and job type indicators. I_{ijt} is "1" if person i worked in industry j in year t and its estimated coefficient captures the wage premium.

Goldberg and Pavcnik (2005) estimated the wage premium using data from all industries, with industry ISIC 62 (retail sales) being the omitted category. Notice that this choice does not affect the results because the estimated wage premium is normalized using the Haisken-DeNew and Schmidt (1997) procedure, which transforms the estimated wage premium into deviations in percentage terms from the employment-weighted average wage premium. Hence this procedure provides estimates for the omitted industry and for the wage premium standard errors. The inverse of the standard errors is used as a weight in the estimation of equations (5) and (7).

One important concern is that worker self-selection could be the reason behind the wage premium, and the specification used to estimate the wage premium does not control for it. The literature offers no clear solution to this problem. Gibbons and Katz (1992) investigated this possibility in detail and could not reject the hypothesis that the wage premium is not related to this endogenous worker-industry matching. On the other hand, Abowd et al. (1999) using matched employer-employee data found that individual specific effects are more important than firm effects in the wage determination. Nevertheless, if individuals do not move across industries, this will not pose a problem for the methodology used here. And this is indeed the case for Colombia, since Goldberg and Pavcnik (2003) found little evidence of inter-industry switching by workers. Thus, if there is self-selection to a specific industry, it appears to be stable over time. This suggests that the time-differenced version of equation (5) and equation (7) will control for this self-selection issue.

The import tariff data at the 2-digit ISIC level are from the Colombian Planning Department. These data are available for only 19 industries, with manufacturing industries among them, and they consist of an average of more disaggregated tariffs weighted by the number of product lines. The coffee price and exchange rate data are from the International Monetary Fund's International Financial Statistics (2001). The exchange rate is the nominal effective rate, which is a currency bundle encompassing the currencies of Colombia's most important trade partners.

3.2 Eslava et al. (2004) Data Set

In Eslava et al. (2004), the TFP is estimated with the Capital-Labor-Energy-Material methodology (known as KLEM, and described in detail in their paper) using firm-level data from the Colombian Annual Manufacturers Survey (AMS) conducted by DANE. The AMS contains firm-level information about the prices paid for inputs, thus providing less biased firm-level estimates because there is no need to rely on non-parametric methods and industry-level price deflators as proxies for missing plant-level price data. The industry-level TFP series was built by aggregating firm-level TFP using production shares as weights. The productivity data are available only for the manufacturing sector at the ISIC 2-digit industry level, as shown in Table 1. The sample used in this paper was ultimately determined by the TFP data availability.

3.3 Descriptive Statistics

Table 2 presents descriptive statistics of the variables used. Here we can see that the average import tariff decreased from 36.6% in 1986 to 13.1% in 1998. The average wage premium declined from 5.7% in 1986 to 2.9% in 1998, whereas the average natural logarithm of TFP increased from 0.332 to 0.647. These statistics suggest that the three variables of interest changed during the trade liberalization period.

4. Empirical Results

In this section, I examine empirical evidence concerning the relationship among trade liberalization, the wage premium, and TFP, in order to estimate the bias caused by the omission of TFP from previous analyses in the literature. Then, I estimate additional regressions to assess the robustness of my results.

To obtain an empirical estimate of the bias generated by the omission of productivity from previous models, the first step is to estimate the first difference of equation (5), as done by Goldberg and Pavcnik (2005) for manufacturing and non-manufacturing industries (see their Table 6b), but here using only manufacturing industries due to our sample restrictions. The results are presented in Table 3.

The technique used in column (1) is ordinary least squares (OLS). The other columns present the 2SLS regressions, using as instruments the 1983 tariff level and its interactions with the exchange rate and coffee prices. The results in column (1) of Table 3 are exactly the same as those reported in Goldberg and Pavcnik (2005) (see their Table 3b column (2)), since they are based on identical data. The estimated tariff coefficient here implies that a ten percentage point decrease in tariffs generates a 1.44% decline in the wage premium. In columns (2) to (6) the estimated tariff coefficients are positive and statistically significant at the 5% level in columns (4), (5), and (6), with the coefficients ranging from 0.096 to 0.199, and statistically significant at the 10% level in column (3), with an estimated coefficient of 0.051. These coefficients are at least two times larger than the estimates in Goldberg and Pavcnik (2005) (see their Table 6b), which included all industries.

The excluded instruments are statistically significant in all first stage regressions, exhibiting an *F*-statistic of at least 136, which is enough to reject the null hypothesis that the 2SLS t-test at the 5% significance level has an actual size larger than 10%. This means that the *t*-tests conducted do not suffer from size problems due to weak instruments[11]. For all specifications, *the* null hypothesis of underidentification (the included and excluded instrument matrix does not have full rank) is *rejected* at the 5% level[12]. Although Goldberg and Pavcnik (2005) did not conduct over-identification tests and Hausman endogeneity tests, these tests are conducted here. The endogeneity tests for the 2SLS estimates do not reject the null hypothesis of exogeneity of tariffs for all specifications. The over-identification test for the specification in column (4) has a *p*-value of 14%, and the null hypothesis that the model is correctly specified is not rejected at the 10% level. In column (6), the null hypothesis is rejected at the 10% level of confidence. This result should make the researcher concerned about the existence of specification problems, which could be caused by the omission of the TFP.

Next I incorporate the natural logarithm of TFP into the econometric model to address the consistency problem presented in section 2. These results are reported in Table 4. The technique used in column (1) is OLS and the technique employed in the remaining columns is 2SLS using the Goldberg and Pavcnik (2005) instruments.

The TFP coefficient is positive as expected and statistically significant in all models, with values ranging from 0.155 to 0.176. The tariff coefficient is positive in all models, but it is statistically significant at the 5% level only in columns (1), (4), (5), and (6). Notice that the lack of statistical significance for the tariff coefficient in columns (2) and (3) is also present in Table 3 columns (2) and (3), suggesting that this is caused by the choice of instruments rather than by the inclusion of TFP.

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[11] Since these specifications contain only one endogenous regressor and less than three instruments, the Stock-Yogo reference values for the bias of 2SLS as a fraction of OLS bias greater than 10% cannot be calculated.
[12] The underidentification test used here is based on the Kleibergen-Paap LM statistic (Baum et al., 2007, p. 21).

The instruments are statistically significant in the first stage regressions, as was also the case in Table 3. The estimated TFP coefficient is positive as expected in all specifications, but statistically significant at the 10% level only in columns (2), (3), and (5).

The null hypothesis of underidentification is rejected for all specifications at the 5% level of significance. The first stage *F*-statistics are above 40, which is greater than the Stock-Yogo reference values for the actual size of 2SLS *t*-tests at 5% significance. These two econometric tests indicate that, assuming that the instruments are valid, instrument irrelevance does not seem to be a problem here. The null hypothesis of the over-identification test is not rejected for the overidentified specifications in columns (4) and (6). Lastly, for all 2SLS specifications, the null hypothesis of exogeneity is not rejected using the Hausman test.

The results here support the argument that the absence of TFP generates biased estimates of the effect of tariffs on the wage premium. When TFP is included in the estimated equation, the tariff coefficient decreases by 36% (column (1) of Table 4) or by 44% (column (6) of Table 4) relative to the specifications that exclude TFP (Table 3).

To illustrate why this bias is important, let's calculate the predicted change in the wage premium using the estimates when TFP is ignored (Table 3 column (6)) and the estimates when TFP is incorporated (Table 4 column (6)) evaluated at the average changes in the tariff (-23.5%) and TFP (31.5%) that occurred between 1986 and 1998. By ignoring TFP, this tariff change leads to a 3.38% decrease in the wage premium. When TFP is included, the tariff effect alone implies a decrease of 2.6% in the wage premium, and the TFP effect is an increase of 4.55% in the wage premium, resulting in a net 2.88% increase in the wage premium.

According to Goldberg and Pavcnik (2005), in Colombia, the unskilled intensive industries suffered the largest tariff cuts. I find that the change in TFP seems to be large when the unskilled share is small (correlation of -0.35). The skilled-unskilled average wage premium gap increases by 0.76% when TFP is ignored. When changes in TFP are incorporated, we can see that the gap actually increases by 0.88%, which is 15% larger than before. So, the current methodology underestimated the effect of trade liberalization on this wage inequality measure.

4.1 Robustness checks

The correlation between TFP and import tariffs may raise concerns that there is multicollinearity in the regressions estimated so far. If this were the case, then in my theoretical framework, the majority of the tariff effect on the wage premium would occur through changes in TFP. Greene (2002, p. 56-57) identifies three common symptoms of multicollinearity. First, the estimated coefficients will have high standard errors and low significance levels, which is not the case with the estimates reported in Tables 3 and 4. Second, the estimated coefficients will have the wrong signs or implausible magnitudes, which again are not the case with the estimates reported in Tables 3 and 4. Third, small changes in the data will lead to large swings in the estimated coefficients. I re-estimated the models by excluding one year at a time and the estimated coefficients did not exhibit any large change, either in magnitude or in the standard errors.[13]

The number of clusters (nine manufacturing industries) in the data set is small and, according to simulations performed by Cameron et al. (2008), this can produce estimates of the coefficients' standard errors that are biased downward. Unfortunately, the number of clusters cannot be increased because nine manufacturing industries is the finest level of disaggregation that can be obtained from the Colombian household surveys. To address this concern I re-estimated all specifications using a pairs bootstrap with the bootstrap-se procedure (Cameron et al. 2008, p. 416) that randomly resample at the cluster level. It means

[13] These results are available from the author upon request.

that if the j-th cluster is selected, then all observations in that cluster will be in the resample. An important aspect is how to incorporate the sample weights in the resampling procedure, given these weights vary over time within cluster. Unfortunately, the literature does not provide guidance on how to procedure in this special case. I chose the naïve strategy of ignoring weights in the bootstrap exercise. The results are displayed in Table 5.

First, I estimated the first difference of equation (5) by OLS, column (1), and then by 2SLS, columns (2) and (3). The tariff coefficients are positive, but statistically significant only in column (3) and at the 5% level. The null hypotheses of underidentification and of exogeneity of the tariff cannot be rejected at the 10% level for both columns (2) and (3). In the first stage of the 2SLS estimates, only in column (2) the excluded instruments are statistically significant. The first stage F-statistic is larger than the Stock-Yogo reference values only in column (2).

Next I re-estimated equation (7). The results for the OLS estimate are reported in Table 5 column (4). The tariff coefficient is positive and not statistically significant at the 10% level. Notice that it is smaller than its Table 5 column (1) counterpart. The TFP coefficient is positive and statistically significant at the 5% level. The test of joint significance of the tariff and TFP rejects the null hypothesis that both coefficients are equal to zero at the 5% level.

The 2SLS estimates of equation (7) in columns (5) and (6) also presented a positive and statistically significant coefficient for TFP, and a positive and not statistically significant coefficient for tariff. Again, the estimated tariff coefficients are smaller than their counterparts in columns (2) and (3) as expected. The null hypothesis that both the TFP and tariff coefficients are zero is rejected at the 5% level for columns (5) and (6).

In the first stage regression, the excluded instruments are statistically significant only in column (5). The TFP coefficient is positive and statistically significant at the 5% level in column (5) and at 10% level in column (6). The null hypothesis of underidentification is rejected at the 5% level in both cases. The first-stage F-statistic is larger than the Stock-Yogo reference value only for column (5). The null hypothesis of exogeneity of tariffs cannot be rejected at the 5% level of confidence for columns (5) and (6).

The results in Table 5 provide support to the Table 4 findings. Adding TFP to the estimated model resulted in statistically significant TFP coefficients in both the first and second stage of the 2SLS regressions, and in smaller estimated tariff coefficients, as predicted. At the end of the day, the inclusion of TFP appears to make an important difference on the estimated effect of tariffs on wage premium.

5. Conclusions

This paper has assessed a major weakness of the current methodology used to estimate the effect of import tariffs on the wage premium: the omission of total factor productivity from the estimated model. This omission invalidates the instrument commonly used in the literature, the pre-trade reform import tariff level, because TFP is positively correlated with both the wage premium and the import tariff setting. This implies a correlation between the error term and the excluded instrument, and thus leads to inconsistent estimates of the tariff's effect on the wage premium.

When the change in TFP is incorporated into the estimated specification, the results reveal that TFP does indeed positively affect the wage premium directly, and, as an included instrument, also has a positive correlation with the change in tariffs, the endogenous regressor in the estimated model. I find that a 10 % increase in TFP leads to a 1.6% increase in the wage premium. The estimated effect of the import tariff is positive, and a ten percentage point decrease in import tariffs decreases the wage premium by 1.1%, which is about 36%

smaller than the estimates obtained in previous studies, which omitted productivity from the methodology.

One consequence of the use of biased estimates from previous studies is that in the Colombian case, the omission of TFP resulted in a smaller estimated increase in the gap between the average wage premium for skilled and unskilled workers, which leads to an underestimation of the increase in inequality due to changes in the wage premium.

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Appendix

Figure 1 – Relationship between industry-level 1983 tariff and the 1998-1984 change in tariff for Colombian manufacturing industries (ISIC 31-39).

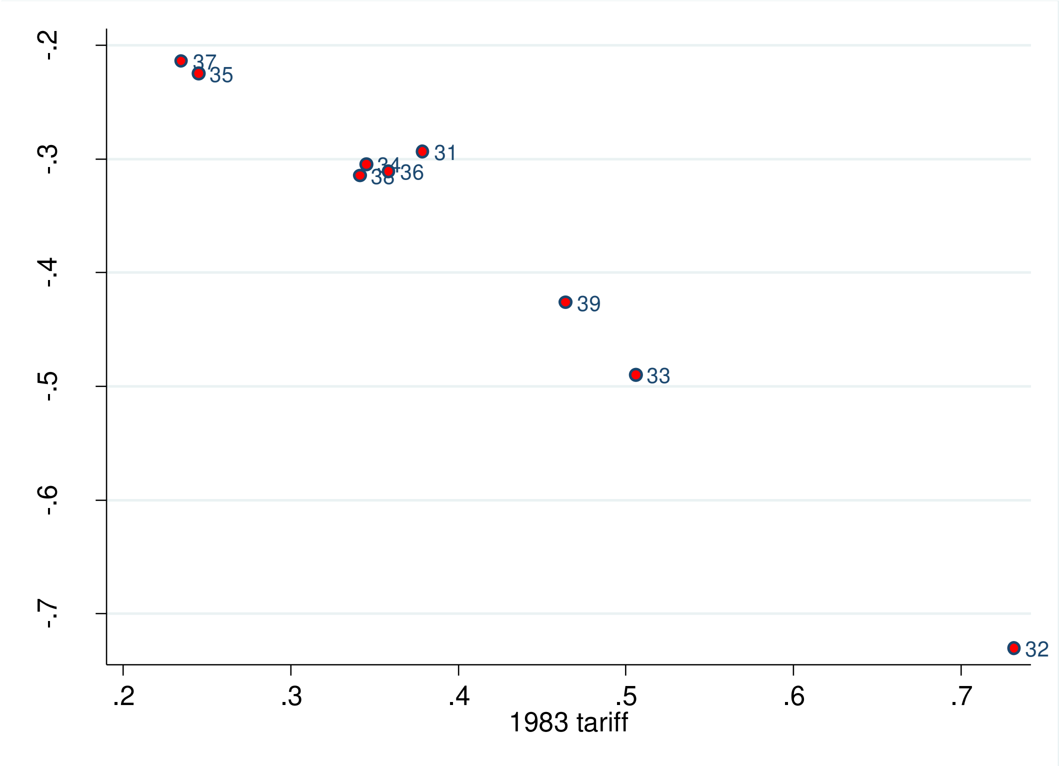

Table 1 – List of Colombian manufacturing industries according to ISIC Rev. 2 classification.

ISIC Rev. 2 code	Industry description
31	Manufacture of Food, Beverages and Tobacco
32	Textile, Wearing Apparel and Leather Industries
33	Manufacture of Wood and Wood Products, Including Furniture
34	Manufacture of Paper and Paper Products, Printing and Publishing
35	Manufacture of Chemicals and Chemical, Petroleum, Coal, Rubber and Plastic Products
36	Manufacture of Non-Metallic Mineral Products, except Products of Petroleum and Coal
37	Basic Metal Industries
38	Manufacture of Fabricated Metal Products, Machinery and Equipment
39	Other Manufacturing Industries

Table 2 - Descriptive statistics of variables for manufacturing industries (ISIC 31-39)

Year	1986-1998			
Variable	Mean	Std. Dev.	Minimum	Maximum
Wage premium	0.040	0.064	-0.096	0.175
Tariff	0.216	0.122	0.075	0.501
1983 tariff	0.401	0.145	0.235	0.732
Coffee Prices	84.314	25.954	42.6	129
Coffee Prices * 1983 tariffs	33.775	16.449	10.007	94.377
Exchange Rate	106.714	22.175	75.5	151.3
Exchange Rate*1983 tariff	42.748	18.086	17.736	110.692
Log(TFP)	0.511	0.257	-0.039	1.138
ΔWage premium	-0.002	0.053	-0.111	0.168
ΔTariff	-0.053	0.079	-0.408	0.007
Δlog(TFP)	0.056	0.083	-0.102	0.277
Number of observations	63	63	63	63

Where Δwage premium = wage premium$_t$ – wage premium$_{t-2}$, with analogous formulas for the other variables.

Table 3 – First difference of equation (5) estimated using OLS and 2SLS estimators with TFP omitted

Technique	OLS	2SLS	2SLS	2SLS	2SLS	2SLS
2nd Stage	(1)	(2)	(3)	(4)	(5)	(6)
ΔTariff	0.144**	-0.007	0.051*	0.168**	0.096**	0.199**
	(0.038)	(0.033)	(0.028)	(0.032)	(0.032)	(0.044)
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Observations	63	63	63	63	63	63
1st Stage						
1983 Tariff		-0.170**		0.951**		0.299**
		(0.007)		(0.144)		(0.115)
Exchange rate* 1983Tariff			-0.002**	-0.010**		
			(0.000)	(0.001)		
Coffee Price* 1983Tariff					-0.002**	-0.006**
					(0.000)	(0.001)
Underidentification test		711.61**	338.48**	2641**	172.47**	2065**
		[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
1st stage F-statistic		561.12	266.90	1022	136	799.5
Stock-Yogo reference values 10% max. IV size		16.38	16.38	19.93	16.38	19.93
Over-identification Test				2.177		2.815*
				[0.140]		[0.093]
Endogeneity test		1.757	1.472	0.133	0.631	0.557
		[0.185]	[0.225]	[0.715]	[0.427]	[0.456]

Dependent variable is Δwage premium, where Δwage premium = wage premium$_t$ – wage premium$_{t-2}$, with analogous formulas for the other variables. Robust standard errors are clustered at the industry level, and are reported in parentheses. p-values are reported in brackets. Sample weights are the inverse of the estimated wage premium standard error.

**, * indicates statistical significance at the 5% and 10% levels, respectively.

Table 4 – First difference of equation (7) estimated using OLS and 2SLS estimators with TFP as an exogenous regressor

Technique	OLS	2SLS	2SLS	2SLS	2SLS	2SLS
2nd Stage	(1)	(2)	(3)	(4)	(5)	(6)
ΔLog TFP	0.160**	0.176**	0.170**	0.156**	0.164**	0.155**
	(0.043)	(0.040)	(0.039)	(0.038)	(0.040)	(0.041)
ΔTariff	0.092**	0.037	0.058	0.106**	0.079**	0.111**
	(0.036)	(0.058)	(0.039)	(0.024)	(0.030)	(0.029)
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Observations	63	63	63	63	63	63
1st Stage						
ΔLog TFP		0.333*	0.293*	0.097	0.250*	0.172
		(0.167)	(0.154)	(0.097)	(0.131)	(0.114)
1983 Tariff		-0.183**		0.867**		0.182*
		(0.029)		(0.104)		(0.081)
Exchange rate*1983Tariff			-0.002**	-0.010**		
			(0.000)	(0.001)		
Coffee Price*1983Tariff					-0.002**	-0.004**
					(0.000)	(0.001)
Underidentification		52.71**	78.44**	201.41**	76.96**	187.12**
test		[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
1st stage *F*-statistic		40.81	60.73	76.52	59.58	71.09
Stock-Yogo reference						
values 10% max. IV size		16.38	16.38	19.93	16.38	19.93
Over-identification				0.956		1.603
test				[0.328]		[0.205]
Endogeneity test		0.804	0.623	0.001	0.108	0.028
		[0.370]	[0.430]	[0.969]	[0.742]	[0.868]

Dependent variable is Δwage premium, where Δwage premium = wage premium$_t$ – wage premium$_{t-2}$, with analogous formulas for the other variables. Robust standard errors are clustered at the industry level, and are reported in parentheses. *p*-values are reported in brackets. Sample weights are the inverse of the estimated wage premium standard error.

**, * indicates statistical significance at the 5% and 10% levels, respectively.

Table 5 – Equations (5) and (7) estimated by OLS and 2SLS with bootstrapped standard errors.

Technique	OLS	2SLS	2LS	OLS	2SLS	2SLS
2nd Stage	(1)	(2)	(3)	(4)	(5)	(6)
ΔTariff	0.188	0.129	0.283**	0.092	0.024	0.192
	(0.120)	(0.120)	(0.121)	(0.114)	(0.135)	(0.252)
ΔLog TFP				0.172**	0.185**	0.151**
				(0.048)	(0.059)	(0.046)
Test of joint significance				17.46**	15.96**	14.98**
for ΔTariff and ΔLogTFP				[0.000]	[0.000]	[0.000]
1st Stage						
ΔLog TFP					0.112**	0.168*
					(0.051)	(0.091)
1983 Tariff		0.702**	0.128		0.609*	0.037
		(0.305)	(0.337)		(0.341)	(0.259)
Exchange rate*1983Tariff		-0.008**			-0.007**	
		(0.003)			(0.003)	
Coffee Price*1983Tariff			-0.003			-0.002
			(0.004)			(0.003)
Underidentification		29.79**	16.63**		27.71**	15.69**
test		[0.000]	[0.000]		[0.000]	[0.000]
1st stage *F*-statistic		24.22	9.682		20.81	8.786
Stock-Yogo reference values						
10% max. IV size		19.93	19.93		19.93	19.93
Over-identification		0.058	0.713		0.641	0.103
Test		[0.810]	[0.398]		[0.423]	[0.748]
Endogeneity test		0.245	0.242		0.266	0.248
		[0.621]	[0.623]		[0.606]	[0.618]

Dependent variable is Δwage premium, where Δwage premium = wage premium_t – wage premium_{t-2},, with analogous formulas for the other variables. Year dummies included in all specifications. Standard errors obtained from pair bootstrap of 50 repetitions and are reported in parenthesis. *p*-values are reported in brackets. The underidentification test is based on the Kleibergen-Paap rk LM statistic. **, * indicates statistical significance at the 5% and 10% levels, respectively. Number of observations is 63 for all specifications.