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Author: Branko Milanovic, Lyn Squire

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Does Tariff Liberalization Increase Wage Inequality? Some Empirical Evidence

Branko Milanovic and Lyn Squire

4.1 Introduction

The evidence reported and reviewed elsewhere in this volume suggests that increasing openness to trade is associated with higher growth and that growth can in turn explain much of the observed reduction in poverty (see in particular Harrison's introduction to this volume). A secondary question is whether the poor benefit as much as, more than, or less than other members of society as a result of trade liberalization. The relationship between trade liberalization and the distribution of income remains a hotly debated issue even though standard theory in the shape of the two-factor, two-country Heckscher-Ohlin model provides an unambiguous prediction: trade liberalization will increase the relative price of the abundant factor, which in the case of developing countries is usually taken to be unskilled labor. This in turn should reduce inequality.

As argued elsewhere in this volume, however, the Heckscher-Ohlin specification is a drastic simplification of a complex phenomenon, and relatively minor steps toward greater realism or a shift in focus toward different aspects of trade liberalization complicate matters (Davis and Mishra, chap. 2 in this volume). To take just one example, Feenstra and Hanson (1997) focus on a different form of "trade": the transfer of production from

Branko Milanovic is a senior associate in the Global Policy Program at the Carnegie Endowment for International Peace and an economist in the research department of the World Bank. Lyn Squire is president of Global Development Network.

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developed to developing countries. In their model, the wage gap between skilled and unskilled workers in developing countries increases, pointing toward increased inequality. Thus, plausible models can lead to quite different predictions.

Whenever theory leads to different predictions, empirical evidence is required to help us choose among alternatives. The available empirical literature, however, does not lead easily to robust conclusions. The combination of a complex phenomenon and data inadequacies renders empirical work both hazardous and partial. Different authors focus on different aspects of the phenomenon ranging from wage inequality to income inequality; they employ different specifications, sometimes relating levels of openness to levels of inequality and sometimes relating changes in openness to changes in inequality; and they use various alternative definitions of key variables, including the measure of openness, with some authors using quantities (trade volumes) and others using policies (tariff levels). The end result is that a careful interpretation of the existing literature requires attention to all these possible points of difference in the various studies.

The purpose of this paper is to present the results of a new empirical investigation of the relationship between trade liberalization and inequality, one that we hope addresses some of the concerns raised above. To this end, the paper draws on a review of the existing empirical literature to identify preferred ways of specifying the empirical model. One outcome of our review is that it leads to the use of two large databases on the distribution of wage income in various forms, sources that have not previously been tapped for this purpose.

The paper begins in section 4.2 with a review of existing empirical work in two critical dimensions: domain and specification. *Domain* refers to the measures of trade liberalization (volumes or policies) and of inequality (incomes or wages) under examination. It also refers to the focus of the study: whether it is a single-country or a multicountry study. *Specification* deals with the issue of whether variables should be measured in levels or in first differences. It also encompasses the important issue of interaction between variables. In section 4.3 we discuss the variables that we use in the empirical analysis. The estimation is presented in section 4.4 for interoccupational wage inequality and in section 4.5 for interindustrial wage inequality.

4.2 Review of the Empirical Literature

As noted in the introduction, the literature contains a diverse collection of empirical efforts to identify the relationship between trade liberalization and inequality. This diversity plagues the interpretation of results and comparisons across studies, but at the same time it provides a valuable source of material to guide the empirical specification estimated in this paper.

Our review covers fifteen papers completed within the last ten years. Of these, six point to a positive relationship between the chosen measure of openness and the chosen measure of inequality. Three indicate that openness increases inequality in low-income countries. Five studies find no impact on inequality. Only one paper points to declining inequality among the "globalizing countries" including the Organization for Economic Cooperation and Development (OECD). In addition, two other papers (Freeman 1995 and Richardson 1995) provide reviews of the then existing empirical literature and conclude that trade liberalization has a positive (increasing) albeit modest impact on inequality. What is surprising about this quick summary is that none of the studies indicate declining inequality in low-income countries, the one region where standard theory predicts such an outcome. The choice, then, seems to be between no impact and increased inequality.

Two qualifications are in order, however. First, the results are often quite fragile: small changes in specification or definition of variables can undermine statistical significance. And second, each of the fifteen studies focuses by necessity on only one aspect of the relationship between trade liberalization and equity. In principle, then, these apparently contradictory results could in fact be perfectly consistent. To explore this further, we examine the studies in each of two dimensions: domain, or the focus of the investigation, and specification, especially whether estimates are levels on levels, or changes on changes. At the end of our discussion of each dimension, we select our preferred option(s) for our subsequent empirical analysis.

4.2.1 Domain

Openness in the majority of papers is defined in terms of trade volumes. Only three papers use some indicator of policy to measure openness. And with respect to inequality, more papers analyze income inequality than wage inequality, with the latter typically being explored in the context of single-country studies exclusively in Latin America.

Income inequality in a cross-country sample is the subject of several papers. An early example is that by Edwards (1997). He regresses the change in the Gini index between the 1970s and the 1980s on a dummy indicating whether a country had engaged in trade liberalization as measured by the average black market premium or the average collected tariff ratio. He finds that trade reform did not significantly affect inequality. Other authors arriving at similar results—albeit using different specifications, time periods, and data—include Londono (2002) and Dollar and Kraay (2001). Barro (2000), however, finds that openness, as measured by trade volumes, is associated with higher levels of inequality in a panel of countries. He concludes: "Basically, the data reveal a long-term positive association between the levels of openness and inequality" (p. 5). Other authors, again

using different methods and variable definitions, concur. Spilimbergo, Londono, and Szekely (1999) and Lundberg and Squire (2003) also detect a link between openness and increased inequality.

Reconciling these results is difficult because they cover different countries and time periods (and could therefore be reflecting different relationships) and because they use different specifications and variable definitions. One possibility that emerges from other work is that country categorization may be important. Several authors (Ravallion 2002; Milanovic 2005; Savvides 1998) find that their preferred measure of openness increases inequality in low-income countries. Barro (2000) also finds the relationship more pronounced in poorer countries. In Spilimbergo, Londono, and Szekely (1999, p. 88) openness affects countries differently depending on their endowments: in capital-rich countries, openness reduces inequality, while in countries with abundant skilled labor, openness increases inequality. The authors argue that the former effect is driven by reduction of capital rents; the latter effect, however, is consistent with Heckscher-Ohlin.

The mix of countries in aggregate studies may therefore be the crucial factor leading to different results. Either way, this is a significant result, for two reasons. First, it runs counter to the prediction of conventional trade theory and raises obvious policy concerns. And second, it suggests that empirical work would benefit from some attempt to interact policy changes and initial conditions to capture the possibility of different effects at different levels of development, a point to which we return below.

Wage inequality is addressed by several authors in the context of specific Latin American countries. For example, Harrison and Hanson (1999) examine the extent to which the increase in wage inequality in Mexico was associated with the 1985 trade reform. They find that the reform did play a part but that other factors, including foreign direct investment, export orientation, and technological change, were also important. Regarding Mexico, Robertson (2000) argues that trade liberalization and "labor flexibilization" led to an erosion of rents in protected industries (which in the case of Mexico were less skilled) while foreign investments increased demand for highly skilled labor. The two effects resulted in widening wage distribution.

Beyer, Rojas, and Vergara (1999) find a similar effect of trade reform on wage inequality in Chile because skill-intensive, resource-based industries expanded following liberalization. Arbache, Dickerson, and Green (2003) find that following the extensive trade liberalization in Brazil in the 1990s, average wage in the traded sector fell compared to the nontraded sector (even after adjusting for education, experience, etc.) and that the only category that was spared a decline was the highly educated, because the returns to education went up. They argue that these results are consistent with the erosion of rents in the traded sector in the wake of liberalization,

and complementarity between skilled labor and new technology brought in by openness.

Behrman, Birdsall, and Szekely (2003) look at the impact of various policies (trade, financial liberalization, privatization, and tax reform) jointly or independently on wage differentials in Latin America during the last twenty years. This study's use of policy indicators (developed by the Inter-American Development Bank) rather than outcomes is very similar to the approach we shall adopt here. Behrman and coauthors conclude that more liberal trade regimes did not have an impact on wage differentials between different education categories. Financial liberalization and high-technology imports in the context of a liberal trade regime, however, contributed to the rising inequality. They conclude, "it is not increases in trade but changes in technology that are associated with growing wage gaps" (p. 30).

These studies suggest two overall conclusions for future empirical work. First, it is important to allow for each country's initial conditions, especially with respect to level of income and the prereform structure of protection, and the reduction in protection by sector in order to understand the impact of trade reform. And second, since trade reforms are seldom undertaken in isolation, allowance has to be made for other reforms. Most often, trade reforms come together, in a package with labor reforms. Disentangling the two effects—in addition to accounting for the effects of technological progress that may be nonneutral—is difficult.

Turning to the choice of variables, we select wage inequality rather than income inequality for both theoretical and empirical reasons. The link between policy reforms and wage inequality is likely to be much stronger than the link between policy reforms and inequality in total income. What happens to total income and its inequality is mediated by a number of other factors, including the role of social transfers (pension spending or family benefits), demographics of the population, family formation and mating, labor force participation, and so on. Since wage inequality is relatively immune to such factors, the link between policy and the distribution of wages should be much stronger than that between policy and the distribution of total income, and it should therefore be easier to detect empirically.

Moreover, labor is the main asset owned by the poor whether they are engaged as unskilled labor or informal workers in the urban areas or as landless laborers or small farmers in rural areas. The return to labor at low skill levels is therefore a critical determinant of poverty. Provided that there is some degree of informal-formal and urban-rural labor mobility, average wages in occupations or industries employing mainly low-skill or unskilled labor will reveal what is happening to the returns to the labor of the poor in general. Any worsening in the distribution of wages is therefore a strong indicator that the poor, both those in wage employment and those in various forms of self-employment, are not benefiting from trade liberalization

to the same extent as everyone else as far as the returns to labor are concerned. In effect, a worsening in the distribution of wages will be magnified when it is translated into the distribution of labor returns because of the large number of low-skill and unskilled workers not receiving a wage for their labor.

There is also an empirical reason for our choice. Inequality measures of total income are not available annually; we have inequality statistics for most countries only for a few years in a decade. The Deininger-Squire database, for example, gives on average an inequality statistic for one out of every five possible country/year combinations. In contrast, the two databases on wages that we use—Freeman and Oostendorp (2000) and the University of Texas Inequality Project database—have annual data for a large number of countries and years. This should increase the power of our empirical estimation and tests.

Trade liberalization can also be measured in many different ways. The primary choice is between policies (tariff reductions, elimination of nontariff barriers, etc.) and outcomes such as trade volumes that are a consequence of trade policies. Both approaches have been used in the literature. Most of the studies reviewed here used trade shares as their measure of globalization. Lundberg and Squire (2003) use the Sachs-Warner index, which, although linked to policies, has been criticized on the grounds that it captures more than trade policy. Edwards (1997) uses a variety of policy measures: average tariff, average quantitative restrictions (QR) coverage, and average black market premium. Savvides (1998) uses a specially created measure of protection covering both tariff and nontariff barriers compiled from United Nations Conference on Trade and Development (UNCTAD) data at the four-digit level of the Customs Cooperation Council Nomenclature. The measure is only available, however, for 1988. Finally, those studying wage inequality within a country are often able to make use of industry-specific tariff rates and quotas.

All of the various ways of specifying variables representing trade liberalization are useful and answer interesting questions. If trade volumes are chosen, then the study says something about the impact of trade volumes on inequality. And for some purposes that may be an interesting question. But, in our view, it does not say much about the impact of policy on inequality, primarily because trade volumes are not determined exclusively by policy. A wide range of factors will influence a country's trade volume: the country's geography, technology, demand conditions in importing countries, competitors' supply conditions, weather, and so on. Even attempts to control for these other factors will inevitably leave a residual that captures more than trade policies. We suspect that the widespread reliance on trade volumes in the empirical literature reflects the relative ease of obtaining data compared with the difficulty of achieving the same for trade policies. Since we are interested primarily in how pro-openness reforms

affect inequality, we prefer to focus on policies and thus place ourselves squarely in the policymakers' corner. We attempt to answer the question that many policymakers naturally formulate when they envisage trade reforms: "What will be the effect of liberalization reforms such as tariff reduction on wage differences between various occupations and industries?"

4.2.2 Specification

Turning to the econometric specification of the relationship to be estimated, we take two points from our review: first, although most researchers have regressed levels on levels, we believe that the work undertaken to date points to the importance of focusing on changes in both the dependent and the explanatory variables; and second, several studies suggest that the impact of policy change depends on the level of development and that therefore interactive relationships need to be incorporated.

The specification in most studies is a relationship between levels of inequality and levels of globalization. These studies generally have more success in finding statistically significant results. Thus, the studies that find a negative impact of globalization on inequality rely on regressions run in levels. For example, Barro (2000) regresses the Gini index on the share of trade in gross domestic product (GDP). Lundberg and Squire (2003) regress the Gini index on the Sachs-Warner measure of openness. On the other hand, the studies that regress changes in inequality on changes in globalization have a much more difficult time finding significant results. For example, Edwards (1997) uses the change in inequality between the 1970s and the 1980s as his dependent variable and a dummy indicating whether a country undertook trade reform as his explanatory variable. Dollar and Kraay (2001) use the growth in the income of the bottom 20 percent and changes in trade volume. Both sets of authors conclude that trade reform and/or changes in openness have no impact on inequality.

Interestingly, two papers undertake both levels-on-levels and changes-on-changes analyses. Milanovic (2005) finds that openness hurts poorer deciles in low-income countries when the analysis relates levels to levels, but he finds no measurable effect when he switches to changes on changes. Similarly, Harrison and Hanson (1999) find that high industry tariffs are associated with greater wage inequality when they conduct the analysis in levels but not in changes. This suggests that either there is no relationship between changes in openness and changes in inequality, or the data are not sufficiently fine to capture such a relationship.

This is an important observation because in our view changes-on-changes regression is the preferred specification. Trade liberalization is presumably a dynamic concept and a continuing one. Regressions of levels on levels, however, typically attempt to compare stable points of equilibrium. Consider this argument. Define liberalization for present purposes as trade openness measured by trade policies. Now imagine two countries,

one of which liberalized trade policy ten years ago and the other of which has literally just implemented its trade liberalization. One would imagine that resource reallocation, changes in factor prices, and other adjustments would have played out in the ten years following the reform in the first country, and the distribution of income would have arrived at a new stationary state. The relationship between policy and inequality could therefore be interpreted as an equilibrium. In the other country, however, trade policy will have changed but the economy, including inequality, will not have had a chance to adjust. If these two countries appear as two observations in a cross-country regression of levels on levels, it is very difficult to interpret the meaning of any results whether statistically significant or not. On the other hand, if the change in policy is related to the change in inequality after some common period of time in both countries, then the results, whatever they may be, are more easily interpreted. With this argument in mind, we focus our empirical work on variables measured in first differences. That is, we focus on *changes* in countries' policy stances and *changes* in inequality outcomes.

A second point that emerges clearly from the review as noted above is that the impact of liberalization may differ depending on the initial conditions of the liberalizing country. However, in the empirical work this approach is not always implemented. The implicit assumption is often that the effects of reforms are the same regardless of the initial level of policy openness or income. In other words, opening up an entirely closed economy by one reform point yields the same results as further opening an already open economy. We shall try to avoid this type of simplification by controlling for the initial level of openness and income and, of course, for other initial differences between economies. Similarly, reforms that are each represented by one policy variable are seen, for econometric convenience, to affect outcomes additively. This is a strong simplification: reforms might often act multiplicatively in that the absence of one type of reform negates the effects of another. We shall try to allow for this by including interaction terms.

4.3 Data Description

4.3.1 Inequality Measures

The first of the two large databases we use to derive inequality measures is that of Occupational Wages around the World (OWW).[1] The data cover the period from 1983 to 1999 and more than 150 countries. The coverage in all its dimensions, however, is problematic and fragmentary. Although

1. The OWW database is available at http://www.nber.org/oww/.

there are 156 countries in total, each country does not provide data (occupational wages) for every year. The yearly country coverage varies between 48 and 76. Occupations included also vary from country to country. Moreover, for a given country, even when the occupational coverage does provide the annual data, it is not necessarily uniform for each year.[2]

Furthermore, it should be noticed that each observation is an observation on "habitual" countrywide wages for a given occupation. Thus, some averaging is already built into the data. That, however, need not be a problem since, for example, the differences in earnings by skill levels are also based on averaging. There is, however, a difference in that the latter are obtained through a statistical analysis that covers a well-defined spectrum of wage earners (labor force survey) and controls for other relevant factors (gender, experience), while the International Labour Organization (ILO) data represent a mish-mash of average "habitual" wages for different underlying populations: some countries—for some years—report monthly wage rates, others report collectively bargained wages, yet others report hourly wages. At times men and women are combined, and at times only wages for men are reported. Freeman and Oostendorp (2000) overcome the problems of data comparability by "calibration," which is essentially a process of finding the adjustment coefficients (based on a regression analysis) for the data given in a "nonstandard" form, where the standard form is defined as the most common form being used in the data set—that is, monthly wages for male workers.[3]

The great advantage of the database (which incidentally also makes the calibration possible) is its size: in the Freedman-Oostendorp "summary" (compendium) of the ILO sources, there are more than 72,000 observations of average occupational wages.[4] For each of the three indexes of interoccupational wage inequality that we calculate (Gini coefficient, standard deviation, and absolute mean deviation from the median), inequality indexes are calculated only for the country/years that contain more than fifteen occupational wages (of the "calibrated" type). After this "filter" and a few others (dropping data for a number of small island economies and dependencies), we are left with 680 observations (country/years) covering the 1983–99 period and 118 countries. The average Gini is about 23.8, the median 21.7, with the standard deviation of about 10. A summary of the data is given in appendix table 4A.1. These inequality statistics can be regarded,

2. For example, the United States gives the data on 11 occupations in 1983 and 150 occupations in 1999.

3. They do several such calibrations and show (in an appendix) that the results (inequality statistics) do not depend on a particular calibration. For our calculations, we have used their suggested base-wage calibration, denoted xlwu in the OWW database.

4. The Freeman-Oostendorp database is indeed a "summary" of ILO data since the data on occupational wages have been collected by the ILO since 1924, while the Freeman-Oostendorp data begin with 1983.

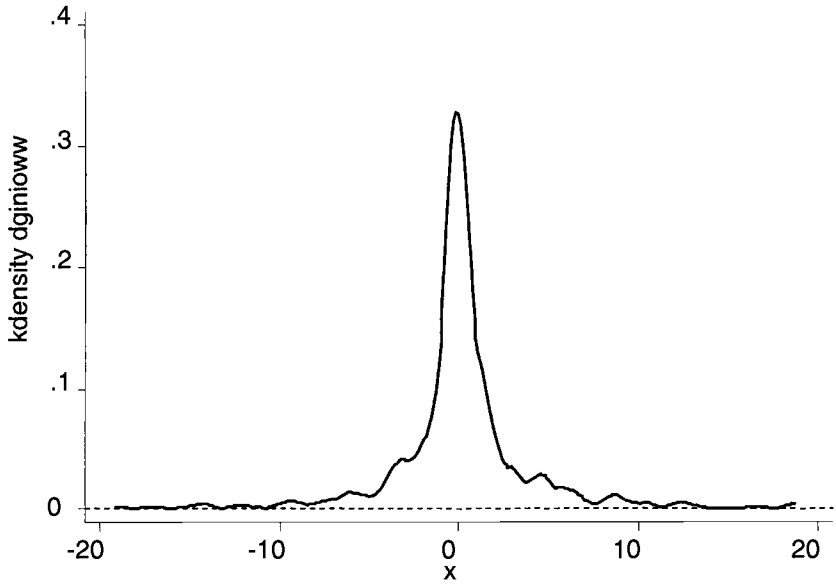

Fig. 4.1 Distribution of changes in occupational wage inequality (Dginioww; in percentage points, 1983–99)

Notes: There are 532 Dginioww observations. Changes are expressed in Gini points.

according to Freeman and Oostendorp, as indicators of both occupational wage inequality and skill premium.[5]

Figure 4.1 shows the distribution of annual changes in the calculated Gini coefficients (Dginioww) over the 1984–99 period. As we observe, the distribution is close to being symmetrical and normal, with the mean being slightly positive (0.17 Gini point) and a zero median.

The second large database of interindustrial wage differences was created by James Galbraith and associates and is known as the University of Texas Inequality Project (UTIP) database (see Galbraith and Kum 2003).[6] The original data come from United Nations Industrial Development Organization (UNIDO) statistics. The UNIDO statistics provide average manufacturing pay by industry. The number of industries (which provide their mean wages) varies between countries and years. On average, there are twenty-four industries per country/year (with the standard deviation of about seven). From these average industrial wages for a given country/year, Galbraith and his associates calculate the Theil index of inequality (variable Theil). The UTIP database covers on average about 90 countries an-

5. Implicitly, the greater the dispersion of interoccupational wages, the greater the return to skills.

6. The data are available at http://utip.gov.utexas.edu/.

Table 4.1 **Simple correlations between various inequality measures and inequality concepts**

	OWW average wage by occupation		
	Gini coefficient	Standard deviation	Absolute mean deviation from median
UTIP (average wage by industry) Theil	0.45*** (513)	0.48*** (513)	0.41*** (518)
OWW			
Gini coefficient		**0.96*** (723)	**0.81*** (723)
Standard deviation			**0.85*** (723)

Notes: Number of observations given in parentheses. Each country/year represents one data point; that is, for each country/year, there is one inequality statistic. Null hypothesis: correlation = 0. Boldface indicates correlation coefficients calculated between various inequality measures from the same database.

***Significant at the 1 percent level.

nually over the period 1975–99.[7] In total, we use 1,651 Theil indexes from 141 countries (see appendix table 4A.2 for details). The average Theil is 5.5, the median 3.8, and the standard deviation 6.4. In about 10 percent of observations intersectoral wage differences are minimal with Theils less than 1. Many of these cases include developed countries (Nordic countries, the Netherlands) but also Algeria, Cuba, Iran, and (until the mid-1980s) China.[8]

Table 4.1 shows simple correlations between different inequality measures from the two databases. We have three inequality statistics from the OWW database (Gini coefficient, standard deviation, and absolute mean deviation from the median) and only one from UTIP (Theil coefficient). Different inequality statistics from the OWW database are obviously strongly correlated (see the figures shown in boldface). The correlation between Theil index from UTIP and Gini from the interoccupation inequality is much less—around 0.4 (see also figure 4.2). Still, it shows that higher skill premium is associated with greater intersectoral inequality. The cor-

7. The data are available at http://utip.gov.utexas.edu/. More recently, the database has expanded to the years prior to 1975. As of January 2004, the UTIP database has almost 3,200 country/year Theils and covers more than 150 countries.

8. It will be noticed that we do not use Gini coefficient here (although we would have liked to do so for a more direct comparison with the OWW data set). The reason is that the UTIP database does not provide individual mean industrial wages, which would allow us to calculate different inequality measures. The authors provide only the "finished" statistic—that is, the Theil index—and not the underlying data. This is not the case with the OWW database, where individual occupational wages by country/year are available and one can thus calculate various inequality indexes.

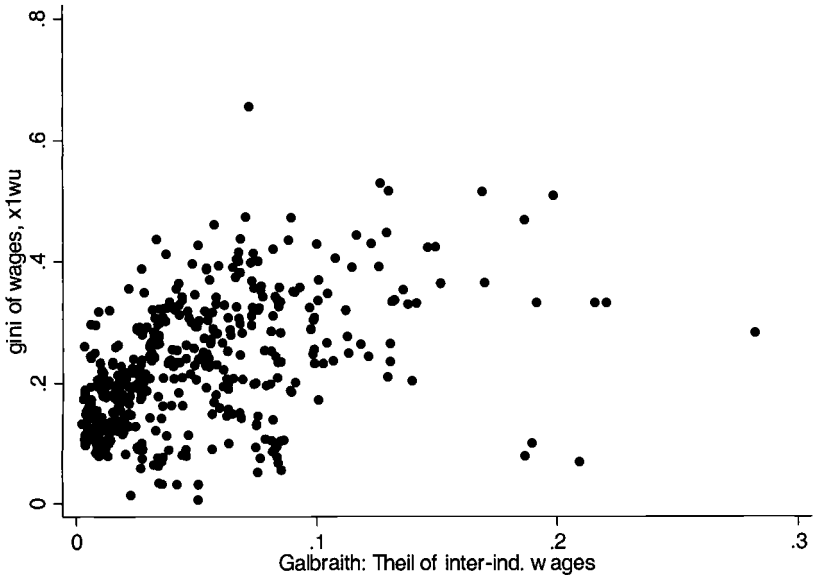

Fig. 4.2 Interoccupational (Gini) and interindustry (Theil) inequality
Notes: Calculated from 513 observations from 79 countries. Total number of observations is 723 (from 103 countries) for interoccupational inequality and 2,160 (from 141 countries) for interindustry inequality.

relation between the changes in the two measures (which we shall be using in our regressions) is virtually zero, however.

4.3.2 Import Liberalization Measures

For import liberalization, we use the World Bank measure of unweighted average tariff (variable Tarf) rate that covers the period from 1980 to 2000, includes 144 countries, and provides 1,255 observations (country/years) in total. The list of countries and number of country/years are shown in appendix table 4A.3. Over this period, the average tariff rate (calculated across the available countries) has been reduced from 28 percent to about 10 percent. Figure 4.3 shows how the distribution of average tariff rates by countries has shifted leftward, with the median, mean, and the standard deviation all significantly less today than in the early and mid-1980s.

The reduction has affected both rich and poor countries. The average tariff rate in poor countries (defined as those with GDP per capita less than $9,000 at international prices) was reduced from 33 percent to 13 percent; for the rich countries, the reduction was from 16 to 7 percent. The pattern of reduction for both poor and rich countries has been very similar to the one shown in figure 4.3: not only are average tariff rates less in 2000 than some twenty years ago, but the differences between the countries is much

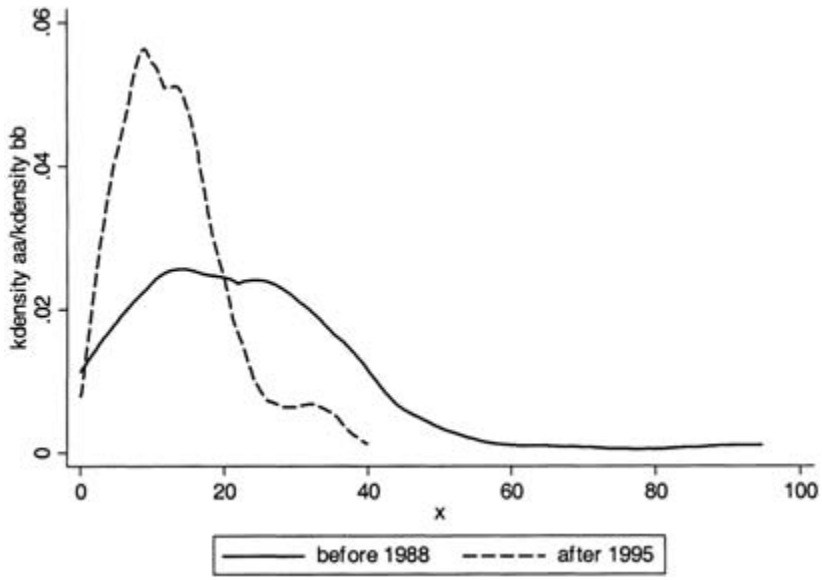

Fig. 4.3 Distribution of countries' average tariff rates in the periods 1980–88 and 1995–2000

Notes: Average tariff rate for a country over a period (1980–88 or 1995–2000) represents one observation. Number of countries is 106 for the first period and 132 for the second.

smaller too (in other words, the distribution of average tariff rates across countries is much more compressed now than in 1980).

One problem when trying to link tariff liberalization reforms to domestic outcomes such as wage distribution is that they are seldom undertaken in isolation. Most frequently, pro-openness trade reforms are accompanied by other "globalization" policies that may well affect labor market outcomes: for example, easier direct or portfolio investment by foreign residents or more liberal regulation of international labor flows. And just as frequently, trade reforms are accompanied by domestic reforms that impact directly on labor markets: "flexibilization" of the labor market, changes in the minimum wage legislation, more (or less) liberal severance pay, reform in the pension regimes, and so on. These accompanying domestic reforms often concern labor—whether they are "anti" or "pro" labor. Sometimes anti-labor legislation accompanies openness reforms because it is felt that liberalization in the foreign arena can be emptied of content (or cannot produce the desired results) if there is no improvement in the domestic legislation—that is, if the latter is deemed too restrictive. Mexico provides one such example (Robertson 2000; Hanson and Harrison 1999).

Alternatively, labor policies, at least for a segment of the labor force, can become more generous if that is the short-term cost the government needs

to pay in order to convince trade unions not to wreck the reforms. In that case, more generous severance pay, low-interest loans to start businesses, and early retirement schemes can all be used to reduce the resistance to reforms and to buy off potential losers. In addition to labor reforms, there may also be accompanying financial reforms: liberalization of interest rates, increased competition in the banking sector, and so on. All of this complicates any attempt to isolate the impact of trade reform on wage inequality. We shall therefore try to control for some of these other policies (labor markets, social transfers).

To measure labor market conditions, we use the Labor Market Data Base constructed by Martin Rama and Rachel Artecona (see Rama and Artecona 2002).[9] Their database has, at five-year intervals (year 1975, 1980, etc.), a number of labor-related measures such as social security contributions (in percentage of gross salary), unemployment rate, replacement rate in case of unemployment, and the like. For our purposes, we focus on two variables—share of labor force covered by collective agreements, and share of the unionized labor force—that allow us to capture the power of trade unions and organized labor.

4.4 Trade Liberalization and Occupational Wage Inequality

We look first at the level relationship between occupational wage inequality and mean tariff rate. Figure 4.4 shows that occupational wage inequality (or returns to education) tends to decrease with average income level of the country (panel A). This is of course what we expect since rich countries have a greater proportion of skilled labor. Likewise, the average tariff rate tends to be lower in richer countries (see panel B). Finally, returns to education increase in level of protection (panel C). This last point would seem to imply that protection is calibrated in such a way as to boost incomes of more skilled workers.[10] However, this relationship may be only apparent and due to the tendency of poorer countries to have, as we have just seen, higher average tariff rates. In fact, once we control for the difference in the returns to education that is due to income levels, the correlation between returns to education and protection vanishes (panel D). It is no longer statistically significant. We can conclude that in a cross-sectional setting, average level of protection and occupational wage inequality do not display any obvious relationship—once we adjust for the fact that poorer countries tend to have both higher returns to education and higher levels of protection.

But this does not necessarily imply that there is no relationship between the *changes* in mean tariff rate and *changes* in returns to education. The

9. The data have been kindly supplied by Martin Rama.

10. Which, by the way, would contradict the general finding of higher protection for less-skilled industries (see discussion above).

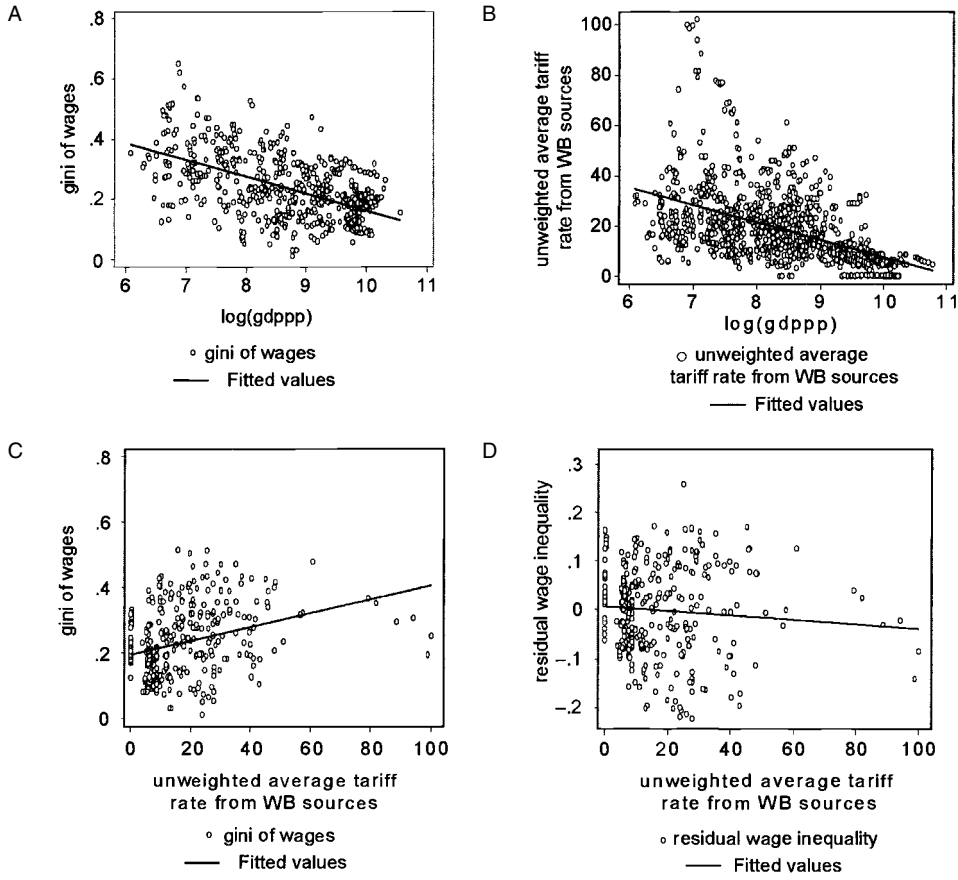

Fig. 4.4 Occupational wage inequality, average level of protection, and mean income: *A,* occupational wage inequality and level of income; *B,* average tariff rate and level of income; *C,* occupational wage inequality and average tariff rate; *D,* occupational wage inequality (controlled for income) and average tariff rate

correlation coefficient is –0.10 (see figure 4.5) and is significant at the 10 percent level. It suggests that there may be a weak negative (and uncontrolled for other variables) relationship such that a decrease in domestic protection (i.e., liberalization) is associated with an increase in returns to education.[11]

11. The two variables are run here and further below contemporaneously. However, since the data on mean tariff rates are often not available for all consecutive years, the Dtarf variable is defined in such a way as to include annual changes wherever available—that is, not only Tarf (t) – Tarf (t – 1) but also Tarf(t) – Tarf(t – 2) when Tarf(t – 1) is not available. Thus, Dtarf is partly lagged (about 20 percent of observations refer to changes between years t and t – 2).

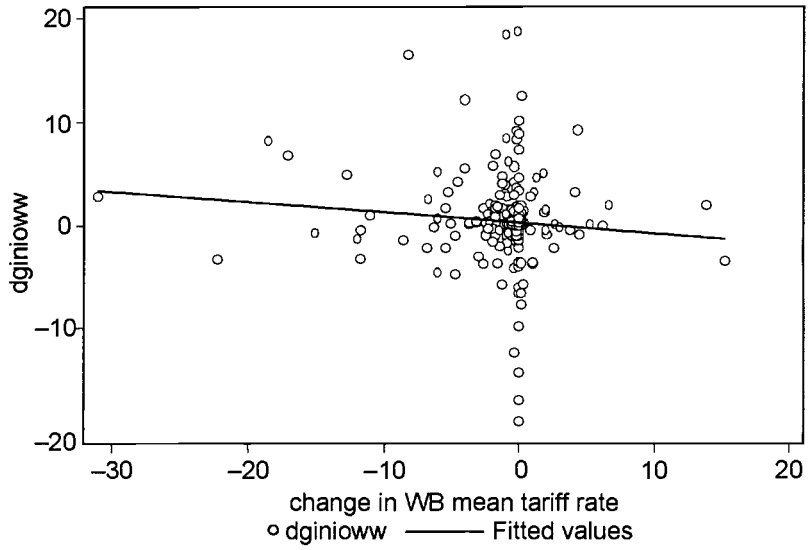

Fig. 4.5 Relationship between change in mean tariff (Dtarf) and change in occupation wage inequality (Dginioww)
Note: The regression coefficient remains negative and significant if outliers—that is, observations such that Dtarf ≤ –25—are eliminated.

Figure 4.6 shows the distribution of changes in occupational Ginis for country/years for which we have corresponding data on changes in protection (that is, figure 4.6 shows the distribution of Dginioww for our sample, not for all the observations of Dginioww that we have[12]). There is, on average, a tendency for occupational inequality to increase (the mean Gini change is +0.36, median +0.05), matching the tendency of tariff rates to go down over the last twenty years (in our sample, the average tariff change is –1.05 percentage points, the median –0.2). Thus, there is some prima facie evidence that decreases in protection and increases in occupational wage inequality may be related.

We look further at this relationship by breaking down changes in returns to education (Dginioww) across average protection changes (table 4.2). There is some evidence that deeper cuts in protection are associated with greater increases in occupational inequality. For example, when tariff protection goes down by more than 10 percentage points, occupational Gini increases on average by 1.45 points. When the reduction in protection is less (between 0 and 5 percentage points), the increase in wage inequality is

12. The shape of the two distributions, though, is almost exactly the same. The number of cases, however, is quite different. Our sample contains only 268 observations, while there is a total of 532 observations of changes in occupational inequality.

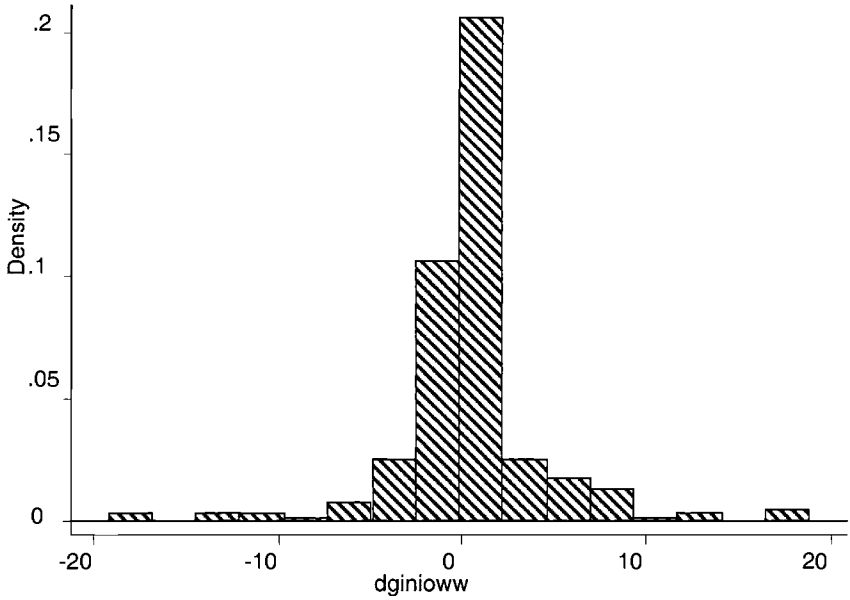

Fig. 4.6 Distribution of changes in occupation inequality (when data on both occupational inequality and tariff changes are available)

Table 4.2 Relationship between occupational wage inequality and protection (average tariff rate)

Change in average tariff rate	Mean change in Dginioww (Gini points)	Standard deviation (Gini points)	No. of observations
Greater than –10 points (in absolute amounts)	+1.45	4.06	10
Between –5 and –10 points	+1.77	5.65	11
Between 0 and –5 points	+0.75	3.55	137
Zero	–0.79	4.50	70
Between 0 and +5 points	+0.43	3.91	34
Between +5 and +10 points	+0.68	1.15	3
Greater than +10 points	–0.73	3.78	2
Total	+0.36	4.00	268

also smaller (+0.75 Gini points). This relationship is not very strong and uniform, though. The change in Gini is, on average, positive even when average tariff rate goes up (by less than 10 percentage points). This in turn suggests that other factors must be at play too. Furthermore, in a number of cases where there was no change in mean tariff rate, average (and me-

Table 4.3 **Relationship between interoccupational wage inequality and level of protection (average tariff rate) in poor and rich countries**

	Poor countries			Rich countries		
Change in average tariff rate	Mean change in Dginioww (Gini points)	Standard deviation (Gini points)	No. of observations	Mean change in Dginioww (Gini points)	Standard deviation (Gini points)	No. of observations
Decrease	+1.31	4.63	77	+0.44	2.56	82
No change	–2.71	6.64	18	–0.13	3.30	52
Increase	+0.29	4.21	30	+0.71	1.01	9
Total	+0.49	5.02	125	+0.25	2.80	143

Note: Poor countries are defined as those with GDP per capita less than $9,000 at international 1995 prices; rich countries are those above that threshold.

dian) wage inequality tended to go down. On balance, we conclude that, while there is some evidence that import liberalization is associated with increasing occupational wage inequality, this is unlikely to be the only factor that matters.

We next split the sample into rich and poor countries (table 4.3). We take $9,000 (in purchasing power parity, or PPP, at 1995 prices) as the cutoff point. This means that in 1980 about three-quarters of all countries in the world are regarded as poor (the proportion is about 70 percent in 2000). Since the data for the rich countries are, on average, more frequently available than for the poor, the cutoff point neatly splits our sample into about two halves.

The table illustrates that the same regularity applies to both poor and rich countries: decreases in protection are associated with higher wage inequality, but so are increases in protection (although the magnitudes are substantially lower). It is mostly when there is no change in mean tariff rate that we find shrinking occupational wage distribution. In effect, out of 122 cases when occupational inequality goes down, about one-third (39) involve situations with no change in mean tariff rate. Poor countries display in all cases (decrease, no change, or increase in protection) greater variability in outcomes. This is illustrated in figure 4.7, where we look at changes in occupational Gini when protection is reduced. The strongly spiked density function for the rich countries (dashed line) shows that reduced protection is accompanied by relatively small and very similar changes in rich countries' Ginis; in contrast, in poor countries, Gini changes (solid line) are much more spread out. The hypothesis of equality of the two distributions is soundly rejected (the Kolmogorov-Smirnov test is significant at less than 0.1 percent). This suggests that while average Dginioww for poor countries may, in response to liberalization, increase more than in rich countries (see table 4.3), the variability of outcomes will also be much greater and thus other variables (and possible measurement error) may play a more important part in explaining changes in wage inequality.

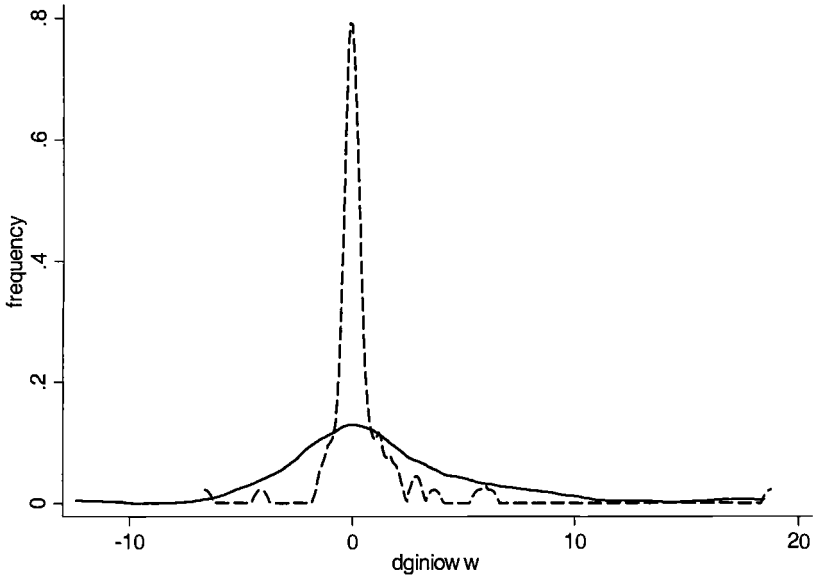

Fig. 4.7 Distribution of Dginioww in poor and rich countries when tariff protection goes down

Notes: Number of observations: 77 for poor countries, 82 for rich countries. Definition of poor and rich countries given in table 4.3 notes. Poor countries are shown by the solid line, rich countries by a dashed line.

In figure 4.8 we therefore focus on poor countries. We look at the change in their occupational wage Gini when tariff protection goes up or down. There are some notable differences: the "down" (solid) line both is thicker in the range Dginioww > 0 and has a much longer right-end tail. Thus, not only is the average Gini change greater when protection is lowered than when it is increased (as we know from table 4.3), but the distribution of Gini changes looks different.[13] There are many more instances of large increases in occupational wage inequality when protection is reduced than when protection is raised.

We now want to investigate how this simple relationship will hold when subjected to a more rigorous analysis. To do this, we estimate the following equation for the change in interoccupation (ΔIneq_O):

$$\Delta \text{Ineq}_O = \text{fct}(\Delta\text{average tariff, labor market conditions, income level})$$

or

$$\Delta \text{Ineq}_O = \text{fct}(\Delta t, s, y)$$

13. However, the Kolmogorov-Smirnov test cannot reject the hypothesis that the two distributions are the same (it is significant at p level is 0.22). The equality of means is rejected at the 10 percent level.

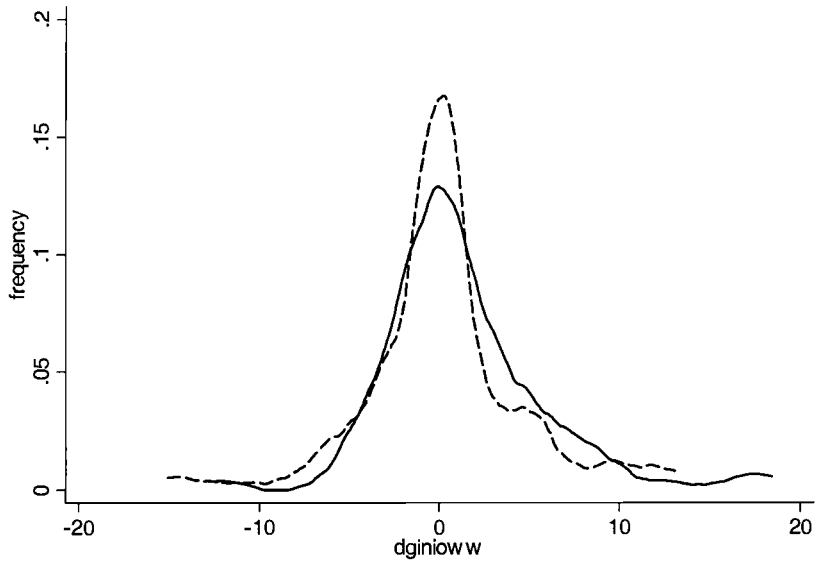

Fig. 4.8 Distribution of changes in interoccupational Gini in poor countries when protection goes up or down

Notes: "Down," denoted by the solid line, indicates the situation when mean tariff rate is reduced; "up," denoted by the dashed line, showed the situation when the mean tariff rate is increased.

A word about the estimation procedure. One might wish to allow changes in average protection level to affect inequality not only contemporaneously but through several time periods (introducing this as a lagged protection on the right-hand side). However, in that case our number of observations—whose low number is already an obstacle to better estimation—drops precipitously and the quality of results deteriorates. We thus assume that one or two years (to the extent that Dtarf also includes some two-year lagged observations) are a sufficient period of time for changes in protection to work their way through wage distribution. Endogeneity is unlikely in levels, and particularly so in a first-difference formulation as here, since change in interoccupational inequality is not likely to have much to do with change in protection. We therefore do not use instruments.[14] Furthermore, the use of first differences implies that idiosyncratic country effects are included.[15]

14. It is also difficult to find reasonable and workable instruments. We tried initial tariff level, on the assumption that reduction in tariffs bears some proportion to their initial levels, but the results were disappointing.

15. Behrman, Birdsall, and Szakely (2003) have the same formulation as here but present also the first-difference formulation of policy changes, or in other words the difference-of-differences formulation (with distributed lags over seven periods on the right-hand side). Their first-difference in levels formulation (table 2) is the same as ours.

Table 4.4 **Explaining interoccupational inequality, 1984–99 (%; dependent variable: annual change in Gini)**

	Regression 1	Regression 2	Regression 3
Δtariff	–0.118	–1.490	–5.707
	(0.097)	(0.033)**	(0.009)**
Ln (GDP per capita)	–0.060	0.448	0.456
	(0.816)	(0.236)	(0.320)
Δtariff · ln (GDP per capita)		0.168	0.688
		(0.057)	(0.008)**
Trade union members as % of labor force (TUMMBR)		0.002	
		(0.920)	
Percentage of workers covered by collective bargaining (TUCVGE)			–0.002
			(0.855)
ΔTarf · TUMMBR		0.001	
		(0.915)	
ΔTarf · TUCVGE			–0.012
			(0.197)
Constant	0.651	–4.221	–4.132
	0.780	(0.205)	(0.331)
Adjusted R^2	0.005	0.02	0.06
F value (p)	1.6	1.7	2
	(0.19)	(0.15)	(0.1)
No. of observations	233	176	79

Note: Levels of significance, p values, given in parentheses.
**Significant at less than the 5 percent level.

Table 4.4 gives the results of the regressions for interoccupational wage inequality. We begin with a very parsimonious formulation where change in interoccupational inequality (Dginioww) is explained by change in average tariff rate (Dtarf) and income. None of the variables is found significant at the 5 percent level; however, Dtarf is negative and significant at the 10 percent level. The situation changes when we introduce the interaction term between the change in average tariff rate and level of income, and trade union membership or percentage of workers covered by collective bargaining agreements. Now, decrease in protection is strongly pro-inequality, with a 1 point decrease in average tariff rate associated with 5.7 percent annual increase in interoccupational inequality.

This pro-inequality effect, however, is reduced the richer the country (because of the positively signed interaction effect; see regression 3), and even for the very poor countries is less than it appears at first sight. Thus, in a very poor country with an income of PPP$1,000, a 1 point decrease in the average tariff rate will be associated with a Gini increase of only 1 percent. Around PPP$5,000 (using regression 3) the effect reverses and trade liberalization begins to be associated with a *decrease* in interoccupational inequality. For example, at the year 2000 mean value of lnGDP per capita

(8.4), the effect of the interaction term is stronger than the effect of change in tariff rate alone; in consequence, pro-openness reforms will be associated with a *decline* in measured interoccupational inequality in richer economies. Finally, note that the fact that labor market conditions are not statistically significant suggests that labor market conditions do not affect the change in the skill premium, while the fact that income is not significant in any formulation is consistent with industry-based (rather than skill-based) bargaining.

The results seem to provide some weak evidence that reduction in average tariff rate contributes to interoccupational wage inequality in poor countries, although the statistical properties of the regressions (most notably R^2) are not strong and the number of observations that we ultimately have to make the regressions is small (79 versus more than 500 observations on changes in interoccupational inequality and more than 1,000 observations on changes in average tariff rates). Therefore we have to take these results with a strong dose of caution.

4.5 Trade Liberalization and Interindustrial Wage Inequality

In figure 4.9 we inspect the relationship between interindustry wage inequality and several relevant variables (all in levels). Panel A shows that when a greater percentage of the labor force participates in collective bargaining, interindustrial wage differences are less. Panel B shows that interindustry wage differences increase as average tariff rate goes up. Now, low tariff rates are found—as we have seen before—more frequently in rich than in poor countries. So are high levels of unionization (collective bargaining). Thus, the two seem to be associated (panel C). This finding implies that some of the positive relationship between the average tariff rate and interindustry inequality from panel B may be due to the presence of high unionization. In other words, the upward slope detected in panel B may be due not to the existence of a real relationship between tariff rates and interindustry inequality but to the fact that countries with low tariffs also display high unionization—with the latter driving interindustry wage inequality down.

When we check for it, however, we find that this is not the case. As panel D shows, once we control for collective bargaining, the relationship between interindustry wage inequality and average level of tariff rates remains positive—in fact, it even becomes sharper. Protection thus indeed seems to drive interindustry wage differences up. We do a further check to make sure that the relationship is not due, in part, to a change in the sample.[16] This is

16. This happens because we have data on tariff rates and interindustry inequality for many more countries than is the case with collective bargaining. Thus, once we control for collective bargaining, the sample shrinks from 757 observations, as in panel B, to 286 observations in panel D.

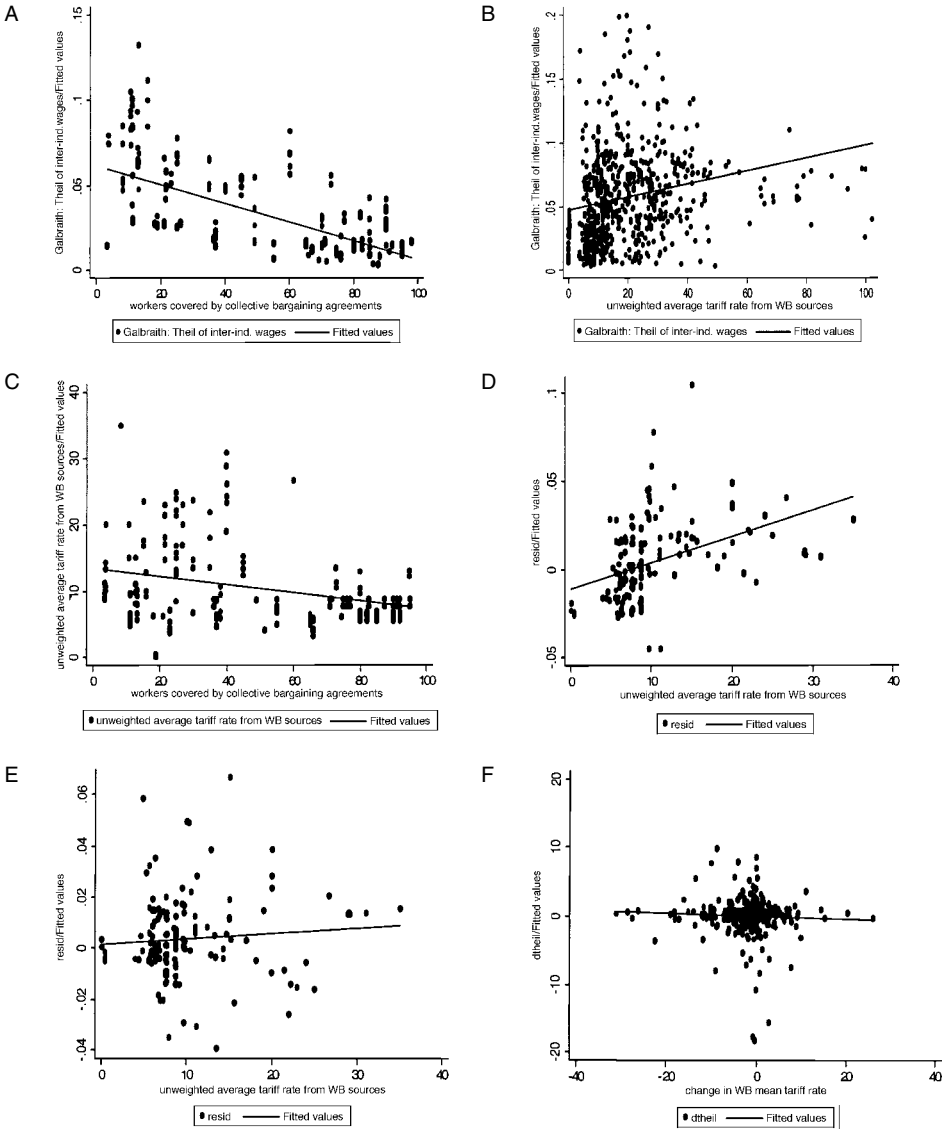

Fig. 4.9 Interindustry wage inequality, average level of protection, and unionism: *A,* interindustry wage inequality and collective bargaining; *B,* average tariff rate and interindustry wage inequality; *C,* collective bargaining and average tariff rate; *D,* interindustry wage inequality (controlled for collective bargaining) and average tariff rate; *E,* interindustry wage inequality (controlled for collective bargaining and level of income) and average tariff rate; *F,* change in average tariff rate and change in interindustry wage inequality

Table 4.5 **Relationship between interindustry wage inequality and level of protection (average tariff rate) in poor and rich countries**

	Poor countries			Rich countries		
Change in average tariff rate	Mean change in Theil (Theil points)	Standard deviation (Theil points)	No. of observations	Mean change in Theil (Theil points)	Standard deviation (Theil points)	No. of observations
Decrease	+0.02	2.56	219	+0.15	0.85	137
No change	+0.13	2.49	44	+0.05	0.60	72
Increase	–0.08	2.00	113	–0.32	3.41	23
Total	–0.01	2.39	376	+0.07	1.29	232

Note: Poor countries are defined as those with GDP per capita less than $9,000 at international 1995 prices; rich countries are those above that threshold.

not the case. When we run the relationship between the average tariff rate and interindustry wage differences (as in panel B) across the sample of country/years in panel D, the results do not change (graph not displayed here). Moreover, even after we control for *both* collective bargaining and income level,[17] the positive relationship between average tariff rates and interindustry wage differences remains (figure 4.9, panel E).

But the relationship between levels may not necessarily be indicative of the relationship between changes. And in effect, inspection of figure 4.9 (panel F) does show that there is a mild negative relationship between changes in average tariffs and changes in the Theil index of interindustry inequality. In table 4.5 we look at whether this relationship holds for poor and rich countries. We easily notice that for rich countries a decrease in protection is associated with an increase in interindustry wage inequality; and the reverse is true for the increase in protection. This in turn indicates that the protected sectors tended to be sectors with lower average wage (that is, less skilled). An increase in protection is associated with lower interindustry wage differences, implying again that higher tariffs will tend to protect sectors with lower average wage (presumably less skilled too). The same pattern, on average, holds for poor countries, although there the average changes are much less clear and the standard deviation much greater. Yet the fact that the same pattern is observable in poor countries as in rich countries (decreased protection associated with increased interindustry wage differences) would also tend to support the view that in poor countries too low-wage or lower-skill sectors tend to benefit from protection.

Figure 4.10 shows the change in interindustry Theil when protection is reduced. In rich countries, the effect does not vary much between the coun-

17. Since income level and interindustry inequality are negatively correlated.

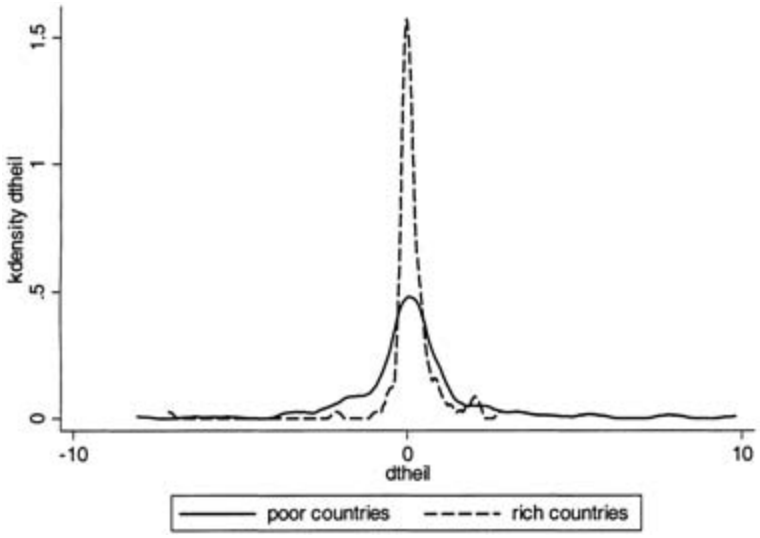

Fig. 4.10 Charge in interindustry Theil when average protection level goes down
Note: Definition of poor and rich countries given in table 4.3 notes.

tries and is bunched around zero with a longer right-end tail (which explains the positive sign of the average). For the poor countries, both right- and left-end tails are approximately equally long and the distribution is flatter.

The equation that we estimate for the change in interindustry inequality ($\Delta Ineq_I$) can be written as

$$\Delta Ineq_I = \beta_0 + \beta_1 (\text{change in average tariff}) + \beta_2 (\text{labor market conditions}) + \beta_3 (\text{change in labor market conditions}) + \beta_4 (\text{income level}).$$

Table 4.6 presents the results for interindustry wage inequality. The first, minimal, formulation shows that none of the variables is significant. In the second formulation, where we introduce the same two interaction terms as before (trade reform and income, and trade reform and union membership), the effect of change in protection on interindustry wage inequality becomes significant and negative. In other words, reduction in protection is associated with greater interindustry inequality: each percentage point of reduction in protection is associated with a 1.7 Theil point increase in interindustry inequality. This implies that as liberalization dissipates the rents from protection, the impact on the distribution of wages works, in relative terms, against those industries that engage more low-skill and unskilled workers. Since such workers will be drawn disproportionately from the ranks of the poor, the implication is that the poor who are engaged in

Table 4.6 **Explaining interindustry inequality, 1976–99 (%; dependent variable: annual change in Theil)**

	Regression 1	Regression 2	Regression 3
Δtariff	0.001	–1.731	–2.207
	(0.947)	(0)**	(0)**
Social expenditures as % of GDP	1.891	1.097	–2.487
	(0.08)	(0.558)	(0.21)
Ln (GDP per capita)		0.148	0.409
		(0.288)	(0.004)**
ΔTarf · ln (GDP per capita)		0.211	0.247
		(0)**	(0)**
Trade union members as % of labor force (TUMMBR)		–0.006	
		(0.268)	
Δtariff · TUMMBR		–0.005	
		(0.003)**	
Number of ILO conventions signed			0.007
			(0.006)**
Δtariff · number of ILO conventions signed			–0.002
			(0.044)**
Constant	–0.119	–1.224	–3.879
	(0.296)	(0.307)	(0.002)**
Adjusted R^2	0.0047	0.1232	0.127
F value (p)	1.56	4.98	5.91
	(0.2118)	(0.0001)**	(0)**
No. of observations	241	171	205

Note: Levels of significance, p values, given in parentheses.
**Significant at less than the 5 percent level.

wage employment benefit less from liberalization than their richer counterparts. Furthermore, provided there is at least some urban-rural and formal-informal labor mobility, the conclusion extends to the poor engaged in nonwage activities.

This effect, however, is less, or is overturned, at higher income levels (as the interaction term between income per capita and average tariff rate has a positive sign). At the median level of (ln) GDP per capita of the countries included in the sample (9.75), the interaction effect is greater than the direct effect of reform. We would thus expect to observe, at the median level of income and above, a decline in observed interindustry inequality even if proliberalization reforms alone tend to increase inequality between the industries. More exactly, the turning point would occur around the world median income, where (in the year 2000) we find countries such as Morocco, Ecuador, and Indonesia. For countries poorer than these we would observe trade reforms increasing interindustry inequality; for richer countries, we would observe a decrease in interindus-

try wage inequality. Similarly to what we found for interoccupational inequality, the effects are stronger and less ambiguous for poor than rich countries.

Reduction of the average tariff rate will tend to contribute to interindustry inequality more in countries with higher trade union density (see the interaction variable in regression 2). This suggests that union power is able to either limit tariff reduction for heavily unionized industries or introduce other, offsetting measures that protect their wages in some other way. It also suggests that union power tends to be concentrated in the higher-skill industries, thereby exacerbating the impact on wage inequality. The same result is observed in regression 3, where we replace trade union membership with the number of ILO conventions as the measure of union power.

4.6 Conclusions

The empirical results provide weak support for the hypothesis that a reduction of tariffs tends to be associated with an increase in interoccupational wage inequality (i.e., education premium) and somewhat stronger support that reduction in tariffs is associated with an increase in wage inequality between industries. The latter effect will be particularly strong in countries with a high density of trade unions. The implication is that the poor benefit less than the rich from liberalization but that their relative position could be improved by simultaneously taking measures to limit trade union power. Average country income plays an important role, though. Through its interaction with change in average tariffs, it offsets the effects of tariff reduction alone so that at income levels above the world median (that is, GDP per capita higher than PPP$4,000 in 1995 international prices) the net effect reverses both for interoccupational and interindustry inequality.

Our results are obtained from the data covering approximately a twenty-year period from 1980 to 2000. The data come from three large and relatively recent databases of occupational inequality (OWW), interindustrial inequality (UTIP) and tariff rates (World Bank data). Although all three databases are rich in terms of the number of observations and do represent a major improvement in data availability, a user cannot escape the impression that there is still a nonnegligible noise in the data, perhaps not so much because the data supplied by different countries and in different periods are wrong but because the coverage of sectors and occupations and the definitions of wages are uneven and vary not only between countries but within countries as well. Thus, the data issues still represent an important obstacle to our ability to draw stronger conclusions regarding the effect of import liberalization on wage inequality in a cross-sectional setting.

Appendix

Table 4A.1 Summary of data from Occupational Wages around the World (OWW)

	Gini of interoccupational wages		
Country	Mean	Standard deviation	No. of observations
Algeria	0.1492	0.0305	8
Angola	0.3787	0.1196	3
Argentina	0.3545	0.1718	3
Australia	0.1543	0.0315	14
Austria	0.1852	0.0212	17
Azerbaijan	0.5310	0.0292	4
Bangladesh	0.2757	0.0537	9
Barbados	0.2283	0.0205	12
Belarus	0.1232	0.0058	5
Belgium	0.0900	0.0092	16
Belize	0.3173	0.0226	12
Benin	0.3863	0.0327	5
Bolivia	0.3843	0.0378	11
Botswana	0.2297	0.0032	2
Brazil	0.2348	0.0000	1
Bulgaria	0.1611	0.0000	1
Burkina Faso	0.3305	0.1400	8
Burundi	0.4175	0.0325	8
Cambodia	0.3751	0.1494	7
Cameroon	0.3866	0.0908	7
Canada	0.1341	0.0099	3
Cape Verde	0.2430	0.0001	2
Chad	0.5411	0.0548	4
Chile	0.3496	0.0053	3
China	0.1509	0.0371	10
Colombia	0.3649	0.0626	2
Zaire or Congo, Democratic Republic	0.4401	0.0000	1
Costa Rica	0.1315	0.0856	3
Côte d'Ivoire	0.3648	0.0854	4
Croatia	0.1930	0.0000	1
Cuba	0.1621	0.0121	6
Cyprus	0.2550	0.0143	16
Czech Republic	0.1339	0.0227	7
Denmark	0.1217	0.0199	10
Djibouti	0.3321	0.0000	1
Estonia	0.2191	0.0145	4
Ethiopia	0.3533	0.0000	1
Fiji	0.3099	0.0198	4
Finland	0.1343	0.0167	14
Gabon	0.3768	0.0562	5
Germany	0.2110	0.0101	17
Ghana	0.3607	0.0000	1
Honduras	0.3637	0.0316	9
Hong Kong	0.2078	0.0403	16
Hungary	0.2217	0.0378	6
Iceland	0.0972	0.0115	2
India	0.3247	0.1436	13
Iran, Islamic Republic of	0.1434	0.0000	1
Ireland	0.1913	0.0014	2
Italy	0.1498	0.0228	12
Japan	0.1995	0.0107	15

Table 4A.1 (continued)

Country	Gini of interoccupational wages		
	Mean	Standard deviation	No. of observations
Korea, Republic of	0.1979	0.0798	10
Kyrgyz Republic	0.3011	0.0153	4
Latvia	0.2558	0.0175	3
Lithuania	0.2328	0.0000	1
Luxembourg	0.1557	0.0000	1
Madagascar	0.1643	0.0536	2
Malawi	0.4522	0.0501	6
Mali	0.3167	0.0000	1
Mauritius	0.3060	0.0172	16
Mexico	0.0616	0.0602	8
Moldova	0.2055	0.0282	5
Mozambique	0.3055	0.0000	1
Netherlands, The	0.1164	0.0080	7
New Zealand	0.2060	0.0145	7
Nicaragua	0.3685	0.0263	6
Niger	0.3754	0.0000	1
Nigeria	0.3616	0.0570	6
Norway	0.1049	0.0242	16
Papua New Guinea	0.3164	0.0048	2
Peru	0.3525	0.0574	10
Philippines, The	0.0974	0.0357	9
Poland	0.1731	0.0446	2
Portugal	0.1398	0.0884	13
Puerto Rico	0.2071	0.0447	13
Romania	0.2139	0.0646	12
Russian Federation	0.2968	0.1173	8
Senegal	0.2644	0.0000	1
Seychelles	0.2593	0.0557	6
Sierra Leone	0.3099	0.0325	8
Singapore	0.3086	0.0199	15
Slovak Republic	0.1490	0.0149	5
Slovenia	0.2078	0.0160	4
South Africa	0.0982	0.0000	1
Sri Lanka	0.2299	0.0426	12
Sudan	0.2917	0.1540	6
Suriname	0.2336	0.0160	4
Swaziland	0.2911	0.0398	2
Sweden	0.1250	0.0349	9
Thailand	0.3057	0.0416	5
Togo	0.3372	0.0678	5
Trinidad	0.2502	0.0235	7
Tunisia	0.2143	0.1523	6
Turkey	0.1805	0.0489	4
Uganda	0.4810	0.0000	1
Ukraine	0.3049	0.0247	3
United Kingdom	0.1660	0.0170	14
United States	0.2097	0.0306	14
Uruguay	0.2578	0.0279	7
Venezuela	0.2622	0.0233	6
Yugoslavia	0.1760	0.0233	10
Zambia	0.3263	0.0569	7
Total	0.2370	0.1082	680

Note: Variable is xlwu from OWW.

Table 4A.2 Summary of data from University of Texas Inequality Project (UTIP)

	Theil index of interindustrial wage differences		
Country	Mean	Standard deviation	No. of observations
Albania	0.0736	0.1213	8
Algeria	0.0144	0.0156	15
Angola	0.3115	0.1041	2
Argentina	0.0512	0.0102	11
Armenia	0.2128	0.1351	5
Australia	0.0110	0.0036	23
Austria	0.0189	0.0065	25
Azerbaijan	0.0385	0.0238	5
Bahamas	0.0987	0.0191	3
Bahrain	0.4035	0.0000	1
Bangladesh	0.0349	0.0196	18
Barbados	0.0584	0.0172	23
Belgium	0.0167	0.0009	18
Belice	0.1059	0.0097	2
Benin	0.0744	0.0141	7
Bolivia	0.0711	0.0317	25
Bosnia and Herzegovina	0.0305	0.0124	2
Botswana	0.0585	0.0153	15
Brazil	0.0776	0.0097	5
Bulgaria	0.0250	0.0300	24
Burkina Faso	0.0328	0.0123	9
Burundi	0.0744	0.0297	13
Cameroon	0.1508	0.0907	20
Canada	0.0199	0.0039	25
Cape Verde	0.0052	0.0038	2
Central African Republic	0.0652	0.0279	17
Chile	0.0657	0.0193	25
China	0.0029	0.0010	7
Colombia	0.0393	0.0055	25
Congo, Republic	0.1144	0.0231	8
Costa Rica	0.0398	0.0188	15
Côte d'Ivoire	0.0737	0.0092	13
Croatia	0.0210	0.0103	11
Cuba	0.0046	0.0009	13
Cyprus	0.0363	0.0086	25
Czech Republic	0.0078	0.0049	9
Denmark	0.0066	0.0010	24
Dominican Republic	0.0792	0.0137	11
Ecuador	0.0495	0.0255	25
Egypt	0.0387	0.0228	25
El Salvador	0.0496	0.0349	17
Equatoria	0.0892	0.0178	2
Equatorial Guinea	0.0301	0.0084	9
Fiji	0.0512	0.0311	21
Finland	0.0107	0.0013	25
France	0.0160	0.0015	17
Gabon	0.1191	0.0410	7
Gambia, The	0.0374	0.0112	8
Germany	0.0108	0.0003	18

Table 4A.2 (continued)

Country	Theil index of interindustrial wage differences		
	Mean	Standard deviation	No. of observations
Ghana	0.1277	0.0363	16
Greece	0.0383	0.0125	25
Guatemala	0.1058	0.0826	21
Haiti	0.0458	0.0084	14
Honduras	0.0712	0.0321	16
Hong Kong	0.0112	0.0065	25
Hungary	0.0188	0.0186	25
Iceland	0.0435	0.0324	22
India	0.0838	0.0100	20
Indonesia	0.0751	0.0205	19
Iran, Islamic Republic of	0.0211	0.0205	18
Iraq	0.0244	0.0118	15
Ireland	0.0311	0.0185	24
Israel	0.0579	0.0144	22
Italy	0.0164	0.0049	24
Jamaica	0.1816	0.1185	15
Japan	0.0355	0.0172	25
Jordan	0.0779	0.0226	23
Kenya	0.0748	0.0143	24
Korea, Republic of	0.0151	0.0059	25
Kuwait	0.2466	0.1247	23
Kyrgyz Republic	0.0851	0.0236	6
Latvia	0.0087	0.0093	6
Lesotho	0.1055	0.0621	7
Libya	0.0324	0.0373	6
Lithuania	0.0713	0.0522	5
Luxembourg	0.0140	0.0034	20
Macedonia	0.0432	0.0225	10
Madagascar	0.0310	0.0182	14
Malawi	0.1128	0.0499	21
Malaysia	0.0313	0.0073	25
Malta	0.0110	0.0035	22
Mauritania	0.1845	0.0583	2
Mauritius	0.0750	0.0245	25
Mexico	0.0290	0.0099	25
Moldova	0.0318	0.0364	9
Mongolia	0.4423	0.4006	6
Morocco	0.0810	0.0145	24
Mozambique	0.1752	0.1233	7
Namibia	0.0314	0.0000	1
Nepal	0.0681	0.0284	9
Netherlands, The	0.0094	0.0025	25
New Zealand	0.0213	0.0150	22
Nicaragua	0.0205	0.0059	11
Nigeria	0.0390	0.0186	14
Norway	0.0095	0.0011	24
Oman	0.1121	0.0118	6
Pakistan	0.0544	0.0124	18

(*continued*)

Table 4A.2 (continued)

Country	Theil index of interindustrial wage differences		
	Mean	Standard deviation	No. of observations
Panama	0.0669	0.0222	23
Papua New Guinea	0.0990	0.0309	15
Paraguay	0.0133	0.0000	1
Peru	0.0830	0.0351	12
Philippines, The	0.0655	0.0155	23
Poland	0.0158	0.0201	25
Portugal	0.0320	0.0064	15
Puerto Rico	0.0818	0.0398	15
Qatar	0.4041	0.0914	8
Romania	0.0103	0.0048	5
Russian Federation	0.0581	0.0090	6
Rwanda	0.0393	0.0092	6
Saudi Arabia	0.1847	0.0000	1
Senegal	0.0433	0.0299	23
Seychelles	0.0075	0.0036	11
Sierra Leone	0.1876	0.1344	2
Singapore	0.0434	0.0130	25
Slovak Republic	0.0163	0.0056	6
Slovenia	0.0165	0.0067	12
Somalia	0.0569	0.0258	6
South Africa	0.0616	0.0071	25
Spain	0.0287	0.0074	25
Sri Lanka	0.0526	0.0130	16
Suriname	0.0570	0.0221	19
Swaziland	0.0993	0.0456	20
Sweden	0.0077	0.0097	25
Syrian Arab Republic	0.0548	0.0566	24
Taiwan, China	0.0155	0.0031	23
Tanzania	0.0630	0.0263	13
Thailand	0.0945	0.0350	13
Togo	0.1050	0.0534	10
Trinidad	0.1579	0.0884	19
Tunisia	0.0896	0.0524	13
Turkey	0.0471	0.0189	24
Uganda	0.1739	0.1034	6
Ukraine	0.0347	0.0261	9
United Kingdom	0.0162	0.0022	25
United States	0.0312	0.0128	25
Uruguay	0.0481	0.0147	23
Venezuela	0.0484	0.0261	22
Yemen, Republic of	0.0670	0.0902	12
Yugoslavia	0.0847	0.0290	5
Zambia	0.0772	0.0147	6
Zimbabwe	0.0544	0.0298	24
Total	0.0548	0.0645	2,160

Table 4A.3 Summary of unweighted average tariff rates from World Bank data

	Average unweighted tariff rate		
Country	Mean	Standard deviation	No. of observations
Albania	17.00	0.00	1
Algeria	25.72	6.73	10
Argentina	18.33	8.05	16
Australia	8.17	3.37	11
Austria	7.05	1.34	11
Bahamas	31.37	1.37	3
Bahrain	5.20	2.63	6
Bangladesh	52.84	33.40	14
Barbados	16.02	4.11	6
Belarus	12.63	0.35	3
Belgium	7.05	1.34	11
Belize	14.66	4.86	5
Benin	33.75	14.30	11
Bolivia	12.58	4.20	16
Botswana	20.55	13.36	2
Brazil	31.89	16.33	20
Bulgaria	16.08	1.88	5
Burkina Faso	32.39	13.28	7
Burundi	29.80	14.94	4
Cambodia	35.00	0.00	1
Cameroon	21.77	5.83	7
Canada	6.74	2.08	9
Cape Verde	22.05	2.90	2
Central African Republic	21.80	6.81	4
Chad	15.75	0.07	2
Chile	14.75	6.57	16
China	33.48	11.59	12
Colombia	20.83	13.42	16
Zaire or Congo, Democratic Republic	23.66	4.76	8
Congo, Republic	19.72	7.44	5
Costa Rica	12.63	5.12	11
Côte d'Ivoire	24.85	3.54	18
Cuba	14.72	7.39	6
Cyprus	11.60	2.50	9
Czech Republic	6.14	1.03	11
Denmark	7.05	1.34	11
Dominican Republic	12.90	4.39	7
Ecuador	17.08	10.70	12
Egypt, Arab Republic of	34.79	8.81	10
El Salvador	11.86	5.83	11
Estonia	0.55	1.25	6
Ethiopia	30.30	1.62	5
Fiji	12.40	0.00	1
Finland	7.05	1.34	11
France	7.05	1.34	11
Gabon	20.16	0.77	5

(*continued*)

Table 4A.3 (continued)

Country	Average unweighted tariff rate		
	Mean	Standard deviation	No. of observations
Gambia, The	13.55	0.07	2
Germany	7.05	1.34	11
Ghana	20.59	8.71	16
Greece	7.05	1.34	11
Guatemala	11.80	4.92	9
Guinea	21.14	24.54	7
Guyana	17.44	4.50	5
Haiti	16.43	9.79	3
Honduras	8.88	1.01	4
Hong Kong	0.00	0.00	21
Hungary	14.42	4.77	13
Iceland	5.97	2.83	10
India	56.49	25.21	14
Indonesia	20.73	9.12	13
Iran, Islamic Republic of	15.43	9.12	3
Ireland	7.05	1.34	11
Israel	7.78	0.74	9
Italy	7.05	1.34	11
Jamaica	16.10	4.47	13
Japan	6.08	0.62	12
Jordan	16.32	3.18	16
Kenya	32.25	10.18	15
Korea, Republic of	15.55	5.20	15
Kuwait	3.90	0.29	4
Latvia	5.23	0.67	4
Lebanon	13.13	5.89	4
Lesotho	17.40	0.00	1
Lithuania	4.14	0.38	5
Luxembourg	7.05	1.34	11
Madagascar	6.73	0.69	7
Malawi	19.71	4.69	16
Malaysia	12.59	2.94	13
Mali	15.66	2.50	5
Malta	7.54	0.96	5
Mauritania	22.42	6.38	10
Mauritius	31.02	6.88	13
Mexico	16.28	5.41	18
Mongolia	8.20	0.00	1
Morocco	28.15	8.34	17
Mozambique	15.74	1.25	5
Namibia	24.40	0.00	1
Nepal	17.73	4.27	9
Netherlands, The	7.05	1.34	11
New Zealand	6.99	3.67	8
Nicaragua	11.02	6.64	10
Niger	18.30	0.00	1
Nigeria	30.14	5.25	16
Norway	4.88	1.22	9

Table 4A.3 (continued)

Country	Average unweighted tariff rate		
	Mean	Standard deviation	No. of observations
Oman	4.12	1.58	9
Pakistan	60.37	14.50	18
Panama	9.96	1.70	5
Papua New Guinea	17.06	5.43	5
Peru	26.48	13.34	19
Philippines, The	23.96	8.1	21
Poland	12.90	3.37	12
Portugal	7.05	1.34	11
Qatar	3.75	1.37	4
Romania	14.20	4.38	7
Russian Federation	11.24	2.48	5
Rwanda	34.53	5.69	4
Samoa	9.00	0.00	1
Saudi Arabia	9.58	4.36	12
Senegal	13.10	1.78	8
Sierra Leone	29.82	8.31	6
Singapore	0.30	0.16	15
Slovak Republic	7.10	0.91	5
Slovenia	11.00	0.69	3
Somalia	29.67	5.98	3
South Africa	11.86	6.43	13
Spain	7.05	1.34	11
Sri Lanka	24.52	8.09	13
Sudan	35.90	21.05	5
Suriname	24.82	10.15	5
Swaziland	15.10	0.00	1
Sweden	7.05	1.34	11
Switzerland	1.59	2.19	8
Syrian Arab Republic	20.57	13.34	6
Taiwan, China	17.94	9.31	13
Tanzania	25.58	5.03	14
Thailand	30.72	10.83	11
Togo	15.25	2.95	4
Trinidad	18.33	1.06	6
Tunisia	27.55	2.47	16
Turkey	21.26	9.32	12
Uganda	16.87	6.89	7
Ukraine	9.83	0.67	3
United Kingdom	7.05	1.34	11
United States	5.93	0.69	12
Uruguay	21.27	11.95	16
Venezuela	19.59	8.32	15
Vietnam	13.50	2.03	4
Yemen, Republic of	20.73	4.94	3
Yugoslavia	11.84	0.09	5
Zambia	20.17	7.85	9
Zimbabwe	16.39	6.23	11
Total	17.65	14.12	1,255

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Comment

Douglas A. Irwin

This chapter tackles a broad but topical subject—the cross-country empirical relationship between trade liberalization and within-country wage inequality, particularly in developing countries. This relationship, and hence the subject of this paper, is somewhat open ended because there is no strong theoretical result that influences our prior belief about what the relationship should be.

One could use the Stolper-Samuelson theorem to suggest that abundant factors of production should benefit from trade liberalization, but the mapping between this theorem and the messy complexity of developing countries is problematic, to say the least.

This chapter uses two measure of wage (not income) inequality, one relating to occupational wages and the other relating to industry wages. These inequality measures are related to a direct measure of a country's average tariff (not "openness" as measured by trade volumes and commonly employed in other studies).

The authors find weak evidence that a reduction in the average tariff rate is associated with higher interoccupational wage inequality in poor countries and somewhat strong evidence of an association with greater interindustry wage inequality. Although the authors are suitably cautious in interpreting their results, I would reinforce this caution. At one point, the authors write that a tariff reduction "contributes to" increased wage inequality. Since establishing a strong causal relationship between the two measures was not the primary object of the paper, I think "association" is a better characterization of the findings. Many factors drive wage inequality, and tariff policy is simply one (a measurable one) among many.

In addition, if there are difficulties in attributing changes in inequality within a country over a given time period to a particular policy measure, these difficulties are aggravated when considering the cross-country evidence. (The United States experienced growing wage inequality in the 1980s, and yet the average tariff did not change at all during the decade.)

At the same time, the results—their general tendency as well as their weakness—do not come as too much of a surprise. As Gordon Hanson's paper (chap. 10 in this volume) points out, six studies of six different countries all found the same general results—that greater openness leads to greater income/wage inequality. Thus, it appears that country studies have uncovered an empirical regularity.

This regularity, however, is itself a bit of a paradox. Given the Stolper-

Douglas A. Irwin is the Robert E. Maxwell '23 Professor of Arts and Sciences and a professor of economics at Dartmouth College, and a research associate of the National Bureau of Economic Research.

Samuelson theorem, which is drilled into the minds of every international economist, we would expect to see the skill premium fall for skilled workers in developing countries with trade liberalization. But perhaps someone should inform those workers that the skill premium increases with globalization. This is because Mayda and Rodrik (2005) examine surveys of pro- and antitrade views around the world and find that, in developing countries, higher levels of education are associated with antitrade views, consistent with Stolper-Samuelson. Yet, ironically, the evidence indicates that those with higher levels of education are precisely those benefitting from more trade.

Several broader points deserve mention as well.

- Sometimes I think we are missing the big picture. In low-income countries, about 60 percent of labor force is in agriculture; most of the rural poor are in agriculture. Yet our data sets usually cover just manufacturing or industry. I think if we are interested in inequality in developing countries, the urban-rural inequality or agriculture-nonagriculture wage gap is much more important than wage inequality within manufacturing (which could be a small part of the story). By focusing exclusively on manufacturing, we might be missing a big chunk of the economy and a big part of intranational wage inequality.
- A paper by Shang-Jin Wei and Yi Wu (2001) gets at this by measuring the urban-rural wage differential for 100 or so Chinese cities (urban areas and adjacent rural counties) over the period 1988–93. The central finding is that cities that experience a greater degree of openness in trade also tend to demonstrate a greater decline in urban-rural income inequality. Thus, globalization has helped to reduce, rather than increase, the urban-rural income inequality. What they suggest is that this pattern in the data suggests that inferences based solely on China's national aggregate figures (overall openness and overall inequality) can be misleading. What I would suggest is that raising rural, agricultural incomes is a key part of reducing inequality, and trade reforms (agricultural or land policy liberalization) may promote this process. Traditionally, trade policies have been strongly antiagrarian in developing countries.

To conclude, the literature on globalization and income inequality includes several country case studies. This paper attempts a cross-country examination of the relationship between tariff policy and inequality, and in some sense it confirms what we have learned from the country studies. Yet because these findings, for developing countries, conflict with the basic Stolper-Samuelson theorem prediction, there is a paradox waiting for more discussion and analysis in future work.

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