ESSAYS ON TRADE AND EQUITY

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Essays on Trade and Equity

Essays over internationale handel en gelijkheid

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by

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Gracias a la vida, que me ha dado tanto.

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Chapter 1

Introduction

Globalization is a central topic of current economic and public debate. The term globalization is associated with the increase of trade volumes, investment flows, outsourcing activities and the expansion of multinational enterprises. However, a commonly accepted definition does not exist and other characterizations intertwine different contemporaneous trends, such as, increased transport facilities, faster and cheaper communication technologies, changes in labor conditions, and even cultural change. No matter which definition is used, there is a common urge to know how globalization affects our economic and social conditions.

Trade liberalization is a main component of economic globalization and as such, there is much interest about the positive and negative effects it may cause. However, trade policy reforms have been a recurring public issue for over a century and economic integration between developed countries had already expanded significantly during the postwar period. The interesting new element associated with globalization is the ever-increasing integration of developing countries into the process, which has been clearly noticeable since the 1980s. In addition, the recent spell of wage inequality and unemployment in OECD countries has added interest to the consequences of trade openness.

As with any critical and heavily disputed public issue, many diverse and contradictory opinions exist. Mainstream economists and financial international organizations claim that trade liberalization and capital mobility have increased economic growth in those countries that have successfully integrated into the world economy. Moreover, these improved economic conditions have allowed these developing countries to reduce poverty (World Bank, 2002). On the other hand, many different observers[1] claim that economic, political and social conditions in developing countries are worsening as a consequence of these enhanced trade volumes and foreign direct investment activities. In OECD countries, some economists and politicians blame globalization for the recent increase in the wage dispersion between skilled and unskilled workers.

Who is right? The debate is still open and as with any complicated and controversial issue, it would probably not vanish soon. Moreover, global trends do not affect each country identically and economic globalization can be better described as a set of opportunities and threats that we must deal with. Indeed, it seems clear that the current trend of increased international relations is only increasing with time. This does not mean that it cannot revert, as it already did at the beginning of the last century, but reinforces our need to understand it better in order to act accordingly.

On its own, income inequality is a pervasive topic in the public agenda. We are concerned about the distributional impact of economic policies and the consequences they may have for poor households. Many economic reforms and most structural adjustment programs are directed to enhance economic growth as an instrument to reduce poverty and increase general living standards. In the meantime, these policies also change – in different degrees– how personal income is finally assigned. While not all economic policies worsen income inequality and complementary redistributional instruments exist, the ensuing trade-off between these efficiency and equity issues is a central concern for policy makers and the general public. Another example of distributional distress is the considerable academic and political interest in the recent surge of wage inequality between skilled and unskilled workers.

Encompassed in this rather general framework, the concrete contribution of the present compilation of articles is to analyze some of the interactions between international trade and income inequality. Trade openness is fundamental to global economic integration and its distributional impact is of considerable interest in the global agenda. However, in this

[1] Here we can include anti-globalization activists, international NGOs, but also prominent economists as the Nobel Prize winner Joseph Stiglitz (2002).

particular topic we find again the same divergent opinions that predominate in the larger globalization debate. Some of the contradictory views are based on methodological and definitional differences and thus, it is important to start our discussion by delimiting the scope of our study.

The particular income distribution level experimented by each country is a complex phenomenon that includes economic and non-economic issues. For instance, factor endowments, technologies, market structure and consumer preferences are some economic factors that determine income inequality. However, the final outcome is also conditioned by non-market institutions, such as the political system, governmental policies, the factor ownership structure, social and historical conditions and the legal system. Thus, as economists we have limited analytical tools to explain how income distribution is determined (Atkinson, 1997). Moreover, trade liberalization is only one of many inequality determinants and it is generally not considered as a main factor. Therefore, in this dissertation we analyze the marginal effect of international trade on the distribution of income.

A further delimitation is that we study only country-specific inequality. When Sen (2002) refers to global inequality and poverty, he claims that it is the central economic issue related to globalization. Indeed, global inequality is higher than in any particular country and absolute poverty affects between a quarter and half of the world population, depending on the definition used. These figures are disturbing, but global redistribution is a topic that exceeds the scope of the present work. Country-specific inequality is still the main priority for national governments and households seldom evaluate inequality beyond their own country. Thus, within-country inequality remains the main element of the distributional debate, although it may not be the most pressing issue from a global perspective.

It has been acknowledged that an important part of the globalization debate between economists and non-economists can be attributed to the definition and measurement of the different variables, and the framework of analysis.[2] For instance, an important distinction to be made is which kind of inequality we are measuring: absolute or relative. Almost

[2] Kanbur (2001) has observed that some of the globalization disagreements can be divided in two broad groups. Group A or "Finance Ministry" and Group B or "Civil Society". Among academics, economists would generally be included in Group A. Ravallion (2003) also analyzes some of the methodological discrepancies between different researchers.

always, economists refer to the latter, which is implicit in most of the common inequality indicators (e.g. Gini coefficient, Atkinson indexes, and share ratios). On the other hand, Ravallion (2003) argues that many non-economists think of inequality in absolute terms.[3] A shift between these two ways of perceiving inequality creates very different results. For example, any growing economy with constant relative inequality experiences increasing absolute inequality. In what follows we study only changes in relative inequality and do not consider absolute inequality at all.[4]

An important complication present in trade liberalization analyses is to separate its effects from other economic policies. Trade liberalization is not easily defined and it is commonly pursued in conjunction with other economic and structural reform policies, i.e. competition policies, public finance reforms, labor market reforms and privatization of public utilities. Moreover, we treat trade openness and trade liberalization as equivalent concepts, both referring to the changes in the level of openness. When possible we try to analyze the isolated impact of trade liberalization on income inequality. Nonetheless, it is not always possible to separate the impact of specific economic policies and our results acknowledge these limitations.

Finally, it is relevant to highlight that we are analyzing global trends between trade liberalization and income inequality. Our theoretical models and empirical tests are built to isolate general relationships between trade and inequality using cross-country and panel data. While these trends are very important to understand, they do not replace country-specific analysis that account for national characteristics and particular economic conditions. For instance, Devarajan and van der Mensbrugghe (2000), Hertel *et al.* (2000), and Ianchovichina *et al.* (2000) have applied computational general equilibrium models to examine the impact of trade reforms on poverty and income distribution for specific countries. However, our results provide important policy lessons for the implementation of medium and long-run development strategies and provide a broad understanding of the distributional effects of openness.

[3] For example, Amiel and Cowell (1999) report that 40% of the students in experiments done in the UK and Israel think of inequality in absolute terms.

[4] The absolute case is in fact not very interesting from an economic point of view. Sustained growth, even when not substantial, requires steady relative inequality declines to maintain absolute inequality constant. Provided that national inequality is generally stable, absolute inequality will be increasing almost in every country with positive growth rates.

Once the scope of this dissertation is delimited, we focus our attention on our main topic and in the following paragraphs we present an outlook of the thesis.

The relationship between trade reform and equity is an old topic of economic research. In their seminal paper of 1941, Wolfgang Stolper and Paul Samuelson presented a theorem that directly linked changes in tariff protection with the functional distribution of income using the Hecksher-Ohlin-Samuelson framework of international trade. The Stolper-Samuelson theorem remains the main theoretical result in trade theory related to income distribution. In Chapter 2 we present an extensive survey of the literature that followed and tested its empirical consequences. The linkage between trade and the personal (or household) distribution of income is a less studied issue, although some recent contributions acknowledge its increasing importance.

Complementing the trade literature, there has been a recent surge in the topic of income distribution that has prompted a new wave of publications (Atkinson, 1997). There is a sizeable literature regarding the interaction between income distribution and other economic variables, such as growth, poverty, trade and economic policy. Most of the theoretical models have turned their attention to the interactions between growth and inequality (Aghion, Caroli and García-Peñalosa, 1999; Bénabou, 1996; Edwards, 1997), the impact of trade on inequality (Bourguignon and Morrisson, 1989; Cline, 1997; Wood, 1994) and the linkages between the economic and political systems (Alesina and Rodrik, 1994; Persson and Tabellini, 1995).

In Chapter 2 we present a thorough literature survey of the particular relation between trade and income inequality, for both the functional and the personal distribution. There we outline the recent trends in national and international inequality and overview the recent debate concerning the impact of trade on wages. This debate focused on the functional distribution of income and in addition, we survey those studies that explored the relations with the household distribution.

A crucial limitation in the study of income inequality is the coverage and quality of the available data. An important step forward was achieved with the introduction of a cross-country inequality data set by the World Bank (Deininger and Squire, 1996). Moreover, this dataset prompted the appearance of a series of econometric studies. Many of these studies have focused on the growth-inequality relationship (Deininger and Squire, 1998; Barro, 2000), while others have been assessing the evolution of national and international inequality (Li, Squire and Zou, 1998; Sala-i-Martin, 2002a and 2002b).

Although the dataset compiled by the World Bank has been extensively used, it has important caveats that have been recently addressed by Atkinson and Brandolini (2001). A distinctive characteristic of inequality observations is that they differ on concept measured (i.e. income or expenditure), reference unit (e.g. person or household) and the primary source of information. Given the scarcity of national income surveys, the grouping of observations with different definitions becomes a critical concern of any inequality dataset. One of the main critiques of the Deininger and Squire dataset has been made precisely on the method used to group heterogeneous observations.

In an attempt to construct a dataset that can overcome some of the problems associated with this fact, in Chapter 3 we propose a new methodology to group heterogeneous inequality observations. The new dataset is based on existing secondary inequality datasets, but our grouping procedure yields distinctive and new inequality time-series. This approach diminishes the measurement error inherent to inequality data sets and by extension, to the outcome of inequality empirical tests. As an added contribution, we use parametric estimations of the Lorenz curve to calculate alternative inequality indexes and absolute poverty estimates. The inequality dataset constructed in Chapter 3 is the basis to the empirical tests that follow in Chapters 5 and 6.

The main objective of Chapter 4 is to overcome the relative dearth of theoretical models that link trade to the household distribution of income. In particular, we construct a theoretical model associating tariff changes with personal income inequality. We use an ownership matrix to map general equilibrium results from the functional to the personal income distribution. In addition, we employ a Sen-type social welfare function (Sen, 1974) to associate different inequality indexes with particular social welfare functions, and this allows us to relate tariff changes directly to social welfare. Next, we embed income inequality indexes into a general equilibrium trade model. Within this general theoretical setting, we use two specific standard trade models (Hecksher-Ohlin-Samuelson and Ricardo-Viner) to obtain particular ownership matrices and trade results. This last step allows us to directly analyze the impact of tariff changes on the household distribution of income and this provides a number of observations that are latter empirically tested in Chapters 5 and 6. Finally, we use the social welfare effects of our general equilibrium setting to explore how the inclusion of distributional concerns by the policy-makers can affect the political economy of tariff formation.

Chapter 5 is a combination of the theoretical model outlined in the previous chapter and our inequality dataset. Here we empirically test the propositions that result from both trade models in a cross-country regression centered on 1994.[5] Our results are compatible with the predictions of the Stolper-Samuelson theorem: higher tariffs are associated with increased income inequality in poor labor-abundant countries and with lower inequality in rich capital-abundant countries. Furthermore, the relationship is monotonically decreasing, suggesting that in middle-income countries import protection does not have an unambiguous effect on inequality. Thus, using a full range of countries where endowment abundance is related to per-capita income levels, we reconcile the Latin American experience with the Stolper-Samuelson predictions. Previous studies had found that trade liberalization had worsened the distribution of income in this region, contradicting the theorem (Davis, 1996; Robbins, 1996a and 1996b). However, most of the region is characterized by middle-income economies, which are not labor-abundant in a global sense but only with respect to developed economies.

The last chapter works again with our new inequality dataset, but now we use cross-sectional time-series econometric estimations to run an empirical reassessment of the well-known Kuznets curve. In 1955, Kuznets observed that income inequality tends to increase in the early stages of development and decreases later. Thus, the evolution of inequality within countries follows an inverted-u pattern. In the recent literature that employs the Deininger and Squire dataset, the curve has been regarded as an empirical regularity (Higgins and Williamson, 1999; Barro, 2000). Our own empirical tests confirm these previous findings. However, when we take into account that the quality of the income data is intrinsically better for richer countries and worsens with per capita income levels, we find that the inequality pattern is better described by a cubic form. Concretely, we find the familiar inverted-u curve in the initial income range, but for higher income levels (reached

[5]This year is selected due to the availability of tariff data. Comprehensive tariff data for a wide selection of countries only exists until the late 1980s.

by developed countries only until the 1980s and 1990s) inequality is again increasing. This result is a consequence of the latest surge of inequality in OECD countries, which has been widely documented (Gottschalk and Smeeding, 2000; Atkinson, 2003).

Additionally, in this chapter we test if trade liberalization has affected income inequality when panel estimations are used. In this case, the lack of tariff data forces us to use other openness indicators (e.g. the Sachs and Warner index, adjusted trade volumes and import duties data). While these indicators have important shortcomings, we expect that they can capture some of the trade liberalization effects. However, we do not find a significant effect of openness on income inequality.

We conclude in Chapter 7, where we present a summary of the main findings of this dissertation and key policy implications that can be drawn from our results. It is important to remark that each chapter has been written as an independent self-standing paper and thus, there are some overlapping themes between particular chapters.

Chapter 2

Literature Survey on the Effects of Trade Liberalization on Income Inequality

2.1 Introduction

The equity effects of globalization have become a central topic of academic and policy debate. While there is not a conclusive definition of this concept, it is associated with increased trade, foreign direct investment, the expansion of multinational enterprises, the diffusion of new communication technologies and the internet, among other global trends. This increase in global integration is not a new phenomenon, but as all significant events it is something that affects society as a whole. This is testified by the increase in anti-globalization protests, as well as the intensive political and economic debate that surrounds many globalization issues.

This paper surveys the literature concerned with the distributional effects of trade liberalization. Trade openness is an essential part of globalization and its interrelationship with income inequality is a relevant issue in the global agenda. The increase in the volume and scope of international trade has been an essential part of the recent surge in the patterns of global integration. Many observers have pointed out that openness has deteriorated world poverty and increased income inequality. In contrast, other sectors have refuted such effects. Among the last, economists have in general presented an optimistic view of the effects of trade liberalization.

However, it is important to keep in mind that this topic can be analyzed from different perspectives. Sen (2002) claims that inequality is the central economic issue related to globalization. However, he defines inequality as the current high levels of global inequality and poverty, and finds less interesting the marginal changes that globalization may have caused. Indeed, global inequality is higher than in any particular country and absolute poverty affects between a quarter and half the world population, depending on the definition used. In particular, Sen argues that the main issue is how the benefits from globalization are distributed, not if everyone is simply better-off with globalization. Nonetheless, within-country inequality is still the main concern of national governments and most citizens evaluate inequality only with respect to their local and national community.

The interaction between trade and income inequality is a complex problem and economists have a partial role to play. The outcome of the market economy is conditional on several economic factors, as endowments, technologies and market structure. But these are not the only determinants and inequality is much influenced by non-market institutions. E.g., political, social, legal and other institutions have critical distributional effects. Therefore, as economists we can analyze some of the mechanisms at play, but not all. In particular, although we find the remarks of Sen relevant, we limit our survey to the marginal changes in inequality caused by trade openness. The length and number of papers analyzed can give an idea of the large literature that is contained within these limits and the broader issues relating to globalization and worldwide inequality are beyond our analytical scope.

With these considerations in mind, we can also highlight that we deal with the economics of globalization. It has been acknowledged that an important part of the debate between economists and non-economists can be attributed to the definition and measurement of the different variables, and the framework of analysis.[1] For instance, an

[1] Kanbur (2001) has observed that some of the globalization disagreements can be divided in two broad groups. Group A or "Finance Ministry" and Group B or "Civil Society". Among academics, economists would generally be included in Group A. Ravallion (2003) also analyzes some of the methodological discrepancies between different researchers.

important distinction to be made is which kind of inequality we are measuring: absolute or relative. Most economists usually refer to the second, which is implicit in most of the common inequality indicators (e.g. Gini coefficient, Atkinson indexes, and share ratios). On the other hand, Ravallion (2003) argues that many non-economists think of inequality in absolute terms.[2] This creates very different inequality results indeed. For example, any growing economy with constant relative inequality experiences increasing absolute inequality. In this survey, when we refer to income inequality we are exclusively dealing with relative inequality and do not consider absolute inequality at all.[3]

Furthermore, globalization has been associated with multiple international phenomena, many of which are not directly related to trade liberalization (i.e. competition policies, market power issues, international capital movements, environmental and social standards). While some of these policies are implemented simultaneously with openness, when possible we try to analyze the isolated impact of trade liberalization on poverty and inequality. Nevertheless, one of the conclusions of this study is that the simultaneity of policy reforms makes this task a difficult one and in general, we have to discuss the effects of entire economic reform packages. Moreover, we treat trade openness and trade liberalization as equivalent concepts, both referring to the changes in the level of openness.

Overall, the main results and conclusions are conditional and in some cases tentative. There are important theoretical cavities and empirical tests are blurred with data limitations, econometric specification problems and measurement error issues. Hence, many studies directly contradict each other and consensus is not reached in many of the issues involved. Nevertheless, some general ideas do remain strong and are generally accepted, despite the noise surrounding the results.

We start our analysis with an overview of national and international income inequality trends. An outline of the main empirical facts provides a necessary departure point in the analysis of the relationships between income equity and trade policy. In section 3 we present a general theoretical framework and then we survey the well-studied debate on

[2]For example, Amiel and Cowell (1999) report that 40% of the students in experiments done in the UK and Israel think of inequality in absolute terms.

[3]The absolute case is in fact not very interesting from an economic point of view. Sustained growth, even when not substantial, requires steady relative inequality declines to maintain absolute inequality constant. Provided that national inequality is generally stable, absolute inequality will be increasing almost in every country with positive growth rates.

how trade liberalization affects wages. This section is finished with an empirical survey on how openness influences the personal distribution of income. Section 5 presents a summary of our main findings.

2.2 The evolution of national and global inequality

2.2.1 Inequality measurement and data problems

Inequality measurement is a theoretical and empirically difficult exercise. On one hand, inequality is represented by the second moment or dispersion of incomes and there are several different indexes that can be used to summarize this information into a single value (i.e. Gini coefficient, Atkinson indexes). Additionally, the use of different indexes has different welfare implications (Dalton, 1920; Kolm, 1969; Atkinson, 1970).

Empirically, there is a dearth of consistent and systematical performed income surveys for most of the countries. This fact considerably limits the number of comprehensive inequality time series for many regions. Finally, many inequality observations are not strictly comparable, since we can measure inequality of different income definitions (e.g. Net and Gross) and also, of consumption and expenditure information.

Several attempts have been made to bridge these limitations and produce large and relatively comparable inequality panel data. The most notorious is the dataset presented by Deininger and Squire (1996). Using three basic quality criteria, these authors gathered single inequality series for a large number of countries. This dataset has been widely used in the empirical literature and hence, most of the studies that involve inequality are dependent on the quality of this dataset. Although the Deininger and Squire (DS) dataset reduced the implicit measurement errors, it did not solve all of them. Atkinson and Brandolini (2001) explain some of these unresolved shortcomings from inequality data. Thus, it is important to bear in mind that the empirics of inequality are permeated by measurement error problems.[4]

[4]For a comprehensive analysis of inequality data measurement issues, see Chapter 3.

2.2.2 National and international inequality trends

Within-country inequality is characterized by heterogeneous experiences. However, there are some patterns that have been identified. Using fragmented historical evidence for a few developed countries, Kuznets (1955) observed an initial increase in the levels of inequality in the early stages of development and afterwards, a steady decrease. This observation gave birth to Kuznets inverted-u curve. In the last decades, however, the experience in OECD countries has reverted the decreasing tendency expected for rich economies. Gottschalk and Smeeding (2000) and Atkinson (2003) identify and inequality U-pattern for these countries after the Second World War. The lack of relevant historical data and relevant inequality time-series for many developing countries, does not allow a similar analysis for this broader group of countries.

However, the most relevant observation concerning income inequality is that within-country variation is much less important than between-country variation (Li *et al.*, 1998). Latin America and African countries usually have the higher inequality levels, while countries from the former communist block and OECD countries experience significantly lower income inequality. Using the DS dataset, these authors argue that income inequality is a relatively stable phenomenon and there are no systematic patterns in country experiences. In addition, the significant between-country difference suggests that inequality determinates change slowly and differ between countries. For instance, initial endowments, development levels, educational attainment, economic policies and demographic factors have all been identified as important determinants of income inequality and we treat these issues in detail in different sections of this survey.

2.2.3 Global inequality

Combining within-country and between-country inequality weighted by population, one can obtain global inequality. This is, the dispersion of incomes among all the individuals in the world. Lindert and Williamson (2001) report that global inequality has risen in the last two centuries. This increase has been driven mainly by between-country inequality changes, spurred by accelerated growth from OECD countries. Within-country inequality, as explained above, has changed episodically during this time and no clear trend emerges.

This conclusion is also reached by Bourguignon and Morrisson (2002), who report that inequality has been increasing since the 1820s, but has been relatively stable after the Second World War.

However, when analyzing global inequality in the last two decades, the results are mixed. In general, one can find studies that claim that inequality has increased and others reporting the contrary. Nevertheless, this major divergence is explained by measurement issues (Ravallion, 2003) and when sound measurement methods are used, the conclusions are less vague. For instance, Svedberg (2001) and Milanovic (2002) review most of the studies and have summarized the main differences. They claim that the critical issue involved is how to measure income and inequality. Some studies use GDP per capita adjusted by PPP and others do not adjust and use instead GDP per capita calculated with exchange rates. Moreover, inequality is sometimes measured by quintile shares (assuming no within-quintile inequality), while others do not even assume within-country inequality. Other important distinctions are provided by the time period used, the inequality indicators and the weight and selection of countries.

A first attempt to create a consistent estimation of global inequality was given by Milanovic (2002) who based his estimations using only national household surveys and adjusting income by PPP. With this methodology he can include about 84% of the world's population. However, he can only study two years: 1988 and 1993, and finds that the Gini coefficient has increased from 0.63 to 0.66 between both years.

However, the generally accepted result is presented by Sala-i-Martin (2002a,b). He uses PPP adjusted income, several comparable inequality indicators and the DS dataset, which provides significant world coverage and a large panel database. He finds that within-country has been slightly increasing but that this effect is upset by the strong decrease in between-country inequality. Thus, he reports that global inequality has been declining between 1980 and 1998. Cross-country inequality is driven mainly by the experience of China, which has had large growth rates in the past two decades.[5] The different outcomes

[5] A limitation in the analysis is that he assumes no within-country inequality for those countries without inequality information and constant inequality for those years were there is no inequality observations. Therefore, his estimates are crude and approximate. However, the possible bias created by these limitations is not expected to change the general results, given the strong influence of China and the relative small weight of the missed observations for which the assumptions are made.

between Sala-i-Martin and Milanovic can be mainly explained by the use of consumption and income, respectively, as the main inequality concept measured.

2.3 Trade and income distribution

The interest on the effects of trade liberalization on income inequality has been revived by two important events. The first one, which has received much of the attention, has been the increase in the wage differential in the United States.[6] The second has been the significant trade liberalization processes in many developing countries in the last two decades. The analysis of both events has been concentrated on the empirical viability of the Stolper-Samuelson theorem. In addition, this has increased the number of alternative explanations that have been proposed to associate trade with income inequality.

This section begins with a small general equilibrium model in which we outline some determinants of the personal distribution of income. From this theoretical framework we can better understand how trade policy can affect income inequality. Then we introduce the Stolper-Samuelson theorem, which is the main theoretical instrument used to analyze the link between trade and income distribution. However, the theorem is aimed at the impact on the functional distribution of income and therefore, we first focus on this aspect. Afterwards, we report the empirical evidence linking openness with the household distribution of income.

In addition, we present a survey of the vast literature dedicated to the connection between trade liberalization and wage inequality. With this background we have a better understanding on how the theorem fares empirically.

2.3.1 Theoretical reference framework

The following section is based on the work of Bourguignon and Morrisson (1989, 1990) and Spilimbergo *et al.* (1999). We assume perfect competition in the factors and final goods markets, well-behaved production and utility functions, I individuals and J factors of

[6] This phenomenon has also been observed in the United Kingdom. In other OECD countries, the main concern has been the increase in the unemployment rates of unskilled workers (Desjonqueres *et al.*, 1999)

production, a closed economy, and no government. We can then construct a simple model relating income distribution with economy-wide endowments. First, aggregate supply is given by:

$$Q = F(E) \tag{2.1}$$

where the vector Q represents total output, the vector F represents production functions, and E is the vector of total factor endowments. With perfect competition in the factor and final goods markets, flexible factor prices and full employment, we obtain:

$$W = P \cdot F'(E) \tag{2.2}$$

where P is the vector of domestic prices, $F'(E)$ is the vector of marginal productivities and W is the vector of factor prices. Here we assume constant returns to scale. The demand for final goods shown in equation (2.3) closes the general equilibrium system.[7]

$$P = P(Q) \tag{2.3}$$

Plugging (2.1) in (2.3) yields:

$$P = P(E) \tag{2.4}$$

Substituting equations (2.1) in (2.2), and assuming that preferences are equal for every individual but productivity may differ across factors, we get:

$$W = W(E, F'(E)) \tag{2.5}$$

We can think of the vector $F'(E)$ as a measure of the productivity of each factor and an indicator of technological progress. We call this vector ϕ, to express this more general idea of factor productivity, thus:

$$W = W(E, \phi) \tag{2.6}$$

[7] In this case the utility function is homothetic. If it wasn't, then the final income distribution will affect aggregate demand.

At this point we drop our closed economy assumption and assume the country is relatively small and domestic prices are determined by international prices: $P = P^w + T$. Where P^w is the vector of world prices and T is the vector that represents the degree of distortions to international trade created by transport costs, tariffs and other barriers. To leave the government out of the analysis, we will assume these are transport costs and other non-governmental trade barriers that create a difference between the domestic and the international price.

Thus, the analogous of equation (2.4) will give:

$$P^w = P^w(E^w) \tag{2.7}$$

where E^w is the vector of world endowments. Following (2.2) we can now construct:

$$W = P^w(E^w) \cdot \phi \cdot T \tag{2.8}$$

Equation (2.8) needs the Factor Price Equalization (FPE) condition to hold. However, this condition is violated, among others, by the presence of non-traded goods and it is hard to sustain empirically (Leamer and Levinsohn, 1995). Therefore, we do not assume FPE and the internal factor endowments determine some final goods prices and also influences the internal factor prices. Under these conditions we obtain:

$$W = W(E^w, E, T, \phi) \tag{2.9}$$

This equation determines the vector of factor prices and from here we can obtain the functional distribution of income. In addition, to move towards the personal distribution of income we can use the following equation:

$$y_i = \sum_j w_j(E, E^w, T, \phi) E_j \omega_{ij} \tag{2.10}$$

where y_i is the personal income of individual i, w_j is the factor price of j, E_j is the total endowment of j in the economy. Finally, ω_{ij} is the share of factor j owned by individual i, with $\sum_i \omega_{ij} = 1$.

We use the Gini coefficient as a representation of the income distribution of the country. Additionally, S is the matrix of coefficients ω_{ij} that represents the ownership structure of the economy. As noted by Spilimbergo *et al.* (1999) the matrix S is determined by historic conditions and may differ quite dramatically from country to country. We can now focus directly on inequality:

$$Gini = g(Y) = g\left(E, E^w, T, \phi, S\right) \tag{2.11}$$

where total income $Y = \sum_i y_i$.

Although (2.11) includes several variables, it is clearly an incomplete model to explain income inequality. However, as pointed by Atkinson (1997) and Spilimbergo *et al.* (1999), such a complete model is not yet available.

There are several ways in which the model can be enhanced. Bourguignon and Morrisson (1989) drop the assumption of flexible factor prices and full employment to obtain a Keynesian approach, where demand plays a key role and has an interdependent relationship with income distribution. Trade also affects income distribution via the exogenously determined foreign demand. As well, there may be some dynamics between the final distribution $g(Y)$ and the aggregate demand function. We can also extend equation (2.11) by introducing imperfect competition, economies of scale, governmental intervention, economic growth dynamics and political economy issues. Some of these inequality determinants are mentioned in following sections.

However, the lack of a formal model to analyze the rich variety of interrelations is a main theoretical limitation in the study of income inequality. In particular, since equation (2.11) is a partial theoretical representation of the link between trade openness and inequality, we can be omitting variables that may change our theoretical results. In particular, we expect that governmental policies are both crucial for trade outcomes and the final household distribution of income.

In any case, our objective is not to explain how income distribution is determined. Thus, we retain this simple model and we use the vector T as a measure of trade liberalization. Hence, equation (2.11) becomes a clear theoretical representation of the relationship between trade and income distribution, although we must bear in mind its limitations.

For the purpose of this survey, it is used to classify the different papers that try to explain such relationship.

2.3.1.1 The Stolper-Samuelson theorem

The main theoretical instrument used to asses the relationship between trade and income distribution is the Stolper-Samuelson theorem.[8] This theorem is directly derived from the Heckscher-Ohlin-Samuelson (HOS) model. The standard assumptions of the model and its definition are detailed by Francois and Nelson (1998):

- Rational behavior by households and firms.
- Complete, perfectly competitive markets.
- Two countries with balanced trade.
- Both countries possess identical tastes, represented by identical homothetic demands.
- Two production factors, endowed in fixed quantities in each country.
- Both factors have uniform quality, are perfectly mobile between sectors and perfectly immobile between countries.
- Two goods, each requiring strictly positive inputs of both factors to be produced.
- Both countries share the same technology, represented by linear homogeneous, twice differentiable, and strictly concave production functions.
- Factor-intensity: At all relevant factor prices, each good uses more intensively one of the factors.
- Factor-abundance: For given endowments, each country is taken to be relatively abundant on one of the factors.
- International trade in goods is costless.

[8] Stolper and Samuelson (1941). Bhagwati (1959) restated the original theorem to solve some theoretical difficulties.

The theorem states that an increase in the relative price of one of the goods will raise the return to the factor used intensively in the production of that good, relative to all other prices; and lower the return to the other factor, relative to all other prices. It is important to remark that the real factor return is increasing for the favored factor and thus, there is an unambiguous gain from the price increase. This result is possible given the "magnification" effect of the price increase, i.e. $\frac{\partial w}{\partial p} > 1$. [9] Although some of the assumptions are considered strong and sometimes violated in reality, the theorem is one of the most elegant and intuitive results derived from general equilibrium models.

The distributional implications of trade policy are easily derived from the theorem. For instance, trade liberalization is generally associated with a decrease in tariffs, which in turn affects directly the prices of goods and hence, changes relative factor prices and inequality.[10]

In principle, the theorem is directly related to the functional distribution of income. Nonetheless, it is usually generalized to include the personal distribution. Consequently, there has to be some assumption regarding factor-abundance and the matrix ownership to generalize its effects. In particular, it is assumed that developing countries are relatively abundant in labor or unskilled labor and that these are equally distributed in the country. On the other hand, capital or skilled labor are abundant but not equally distributed in developed countries. Thus, trade liberalization in developing countries increases the returns to labor and reduces the returns to capital. Since labor is equally distributed, then the overall effect is a reduction in inequality. In the developed world, it has the opposite effect: an increase in the returns to capital and higher inequality.

With regard to our reference framework, the theorem gives a clear theoretical link between final goods prices, factor endowments and factor prices. The simplicity and strong intuitive conclusions of the theorem are very appealing and the Stolper-Samuelson remains as the most accepted theoretical approach to the relationship between trade and income distribution.[11]

[9] This is a direct consequence of the zero-profit condition of the general equilibrium system.

[10] For a complete exposition of the application of the Stolper-Samuelson theorem to trade, see Deardorff and Stern (1994) and Cline (1997).

[11] However, the original article was initially treated as a theoretical curiosity given its strong assumptions and was regarded as impractical to deal with concrete trade policy (Deardorff and Stern, 1994; Cline,1997)

Nevertheless, there are other factors at work which can influence the results. For developed countries, redistributive policies and the endowment of asset earnings can have important effects on the personal distribution. In developing countries, unemployment, trade composition and the endowment of human capital are all variables that can hinder the results (Litwin, 1998). Another difficulty involved is the already mentioned strong and demanding assumptions of the theorem. The departure from some of these assumptions can undermine or even reverse the Stolper-Samuelson results. In the following section we address the main considerations that have been used to challenge the theorem.

2.3.1.2 Theoretical issues revising the Stolper-Samuelson results

The direct and indirect empirical tests of the theorem are plenty. We analyze in detail these studies below, but at present we can state that there are many studies with mixed results and some clearly invalidate the theorem. For instance, Davis (1996) presents a short review of studies for developing countries were he states that there is enough empirical evidence to call the case an anomaly. Hence, many papers have focused their attention at explaining the possible theoretical issues that can clarify this apparent anomaly. However, there is not a clear agreement on the empirical evidence and the following departures from the main model only illustrate the theoretical possibilities that have been studied.

It is important to remember that some papers use the functional distribution of income (i.e. wage differentials) and others use personal income distribution indicators (i.e. Gini coefficient and Lorenz curves) to test the theorem. Hence, not all the explanations apply to each case.

There are several theoretical propositions in the literature to explain how the Stolper-Samuelson results can be offset. Some of them are not critical to the theorem, but others are. These can be divided into three broad groups: the theoretical explications that deal with the functional distribution of income; market structure issues; and those authors that use theoretical explanations to address changes in the personal distribution of income. We deal with each in the following sections.

Theoretical explanations addressed to the functional distribution

In the first group we can name several explanations that refine and/or generalize the theorem by dropping one or more assumptions.

Factor intensity reversals. The Stolper-Samuelson theorem requires technology to be identical across countries and that each good does not change its factor-intensity. Analytical work on production functions with constant elasticity of substitution (CES), showed that in the face of a sufficiently large shift in relative factor prices, goods can switch over from being intensive in one factor to being intensive in the other (Bhagwati and Dehejia, 1994). They also mention that empirical work has shown that this can be a distinct possibility.

The Metzler paradox. When the foreign offer curve is inelastic, an increase in the tariff may actually decrease the price of the protected good. This is achieved through a large improvement in the terms of trade. Thus, the tariff increase actually lowers the returns to the scarce factor (Metzler, 1949). However, the data does not suggest that this has happened (Francois and Nelson, 1998).

Higher dimensions of factors and goods. When the number of factors and goods is increased, the definition of factor-intensity and factor-abundance is ambiguous. Yet, Ethier (1984) demonstrated that with the same number of goods and factors, every good is a "friend" to some factor and an "enemy" to some others. Therefore, the increase in the price of a friendly good will raise the price of the factor to which the good is friendly and decrease the price of the factor to which the good is an "enemy". Nevertheless, there are few chances of identifying global or natural friends and enemies. The alternative that can be used is to identify friends and enemies on average and hence, the strong results of the theorem are undermined, but still valid in a broad sense.

Intermediate goods. As explained in Francois and Nelson (1998), the introduction of inter-industry linkages alters the mechanism that transmits the effect of goods prices to factor prices. In particular, the changes in goods prices are transmitted first to value added prices, and then to relative factor prices. In this case, the theorem will hold when goods are still assumed to be homogenous. However, the recent literature on trade with inter-industry linkages emphasizes on intermediate product differentiation and in this setting, the Stolper-Samuelson result may no longer hold.[12]

Non-traded goods. The presence of national goods with prices not determined by world prices can attenuate the results of the theorem. A change in world prices affects the price of non-traded goods through the market clearing conditions, but in this case the price and output can change to restore equilibrium. Thus, the production and intensity shifts that yield the wage increase are not assured (Winters, 2000a).

Specific Factors. This is the well-known Ricardo-Viner trade model. In this setting we have three production factors and two are specific to each sector. The third factor, usually labor, is mobile between sectors. Under these conditions, the distributional results are more complicated than in the Stolper-Samuelson theorem. The wage of the specific factor increases when the relative price of that good rises, with the inverse result for the other factors price. Labor prices increase, but proportionally less than the increase in the final goods price and in general there are no simple generalizations concerning the changes in real wages.

Hence, in this setting it is hard to draw straightforward conclusions about the distributional impact of tariff reductions. The distribution of land and capital, which are usually taken as the specific factors, is not easily generalized and the impact on labor wages can be insignificant in some cases.

Factor immobility across sectors. In this case, the identity of the industry in which the factor was employed determines the changes in factor prices. This issue relates to the Ricardo-Viner specific-factor model. However, the inconvenience can be solved by using a weaker version of the "friends and enemies" argument. Hence, assuming that capital is sector specific and labor is mobile, each good is friend to its own specific capital and enemy to the other (Deardorff, 1993).

Non-diversification. Bhagwati (1959, 1998) presents the case when the non-diversification assumption is dropped. In the 2x2 model, when a country specializes its production in one of the goods, both factor prices increase as the price of the specialized good raises.

[12]The net effect of trade on wages will depend critically on global variety and scale effects related to specialization of differentiated goods (Francois, 1996; Puga and Venables, 1997).

He labels this the "lifting-all-boats" effect, since the asymmetrical effects on factor prices central to the Stolper-Samuelson result disappear.

Technological change. In the HOS model, technology is assumed to be constant and equal among sectors and countries. Relaxing this assumption yields several possibilities. First, if technologies are different among countries, Desjonqueres *et al.* (1999) argue that what remain critical are the rates of technological progress. However, they show that the Stolper-Samuelson results still hold in general.

Secondly, when global technology is changing then the results of the theorem can be reversed. In particular, sector-biased technological change has been the main argument in explaining the wage dispersion process in the United States. This issue is analyzed in depth when we review the wage inequality debate.

Capital Mobility. Another of the assumptions that can be relaxed is that of immobility of factors between countries. In particular, this rules out foreign direct investment (FDI). Some authors have acknowledged the important distributive effects of outsourcing by Multinational Enterprises (MNEs) in developing countries.[13] Given the sharp increase in the volume of FDI from the North to the South in recent years, this is a persuasive explanation for the apparent Stolper-Samuelson anomaly in the South. We analyze this topic further in the section concerning market structure issues.

Diversification cones. Finally, we present a compelling theoretical explanation to revise the Stolper-Samuelson results in developing countries. In a world with wide divergences in relative endowments, global FPE is ruled out. Instead, countries are grouped in "diversification cones". Inside each cone, countries have similar relative factor endowments and produce the same range of goods. Hence, there is local FPE and specialization in certain goods. In this setting, countries compete within their own cone and trade liberalization will increase the price of the locally abundant factor in the cone[14] and the theorem's results must be revised. The main idea of the theorem is still valid, but the reference group for which to compare relative factor abundance is no longer the global economy, but each country's own diversification cone.

[13] See for example, Feenstra and Hanson (2004).
[14] Dixit and Norman (1980).

Using this idea, Davis (1996) modifies the original HOS model by including many countries, three goods and two factors. In particular, a LDC can be labor-abundant in a global context, but capital-abundant in their diversification cone. In this case, the change in factor prices is the opposite to the anticipated changes predicted by the conventional generalization of the Stolper-Samuelson theorem.

Theoretical explanations dealing with the personal distribution

As mentioned above, the Stolper-Samuelson was originally intended to deal only with the functional distribution of income. Nonetheless, some authors extend the basic theorem's results to analyze the impact of trade on the personal distribution of income.

The following set of papers explains some shortfalls of the Stolper-Samuelson theorem to deal with the personal distribution of income. In relation to our reference framework, this group of authors mainly looks after the dynamic interaction between the relative initial endowments in the economy and the ownership matrix. Using these interactions, they explain the influence of trade liberalization on the personal distribution of income.

THE OWNERSHIP STRUCTURE. Spilimbergo *et al.* (1999) state that the ownership matrix is mainly determined by the initial relative endowments and that it varies across time and countries. They observe that some factors can be accumulated without restrictions (land and capital), but others have a natural limit (skills and education). Therefore, when a country is mainly endowed with the first kind of factors, income distribution can be very unequal. On the other hand, economies endowed with the second kind of factors are expected to be more egalitarian.

Therefore, the initial endowment is not only influencing inequality through the Stolper-Samuelson mechanism, but it can also determine the initial level of inequality. Moreover, unequal distributions can create political and institutional mechanisms to maintain the *status quo* and the interrelations that this generates are further analyzed below.

DYNAMIC FRAMEWORK. Fisher (2001) incorporates a dynamic framework to the analysis. This is a major improvement, since it is then possible to study the interaction between income distribution and the ownership matrix. This author uses a set of two factors: unskilled labor that is equally distributed to all agents and wealth, which is unequally distributed. By wealth he refers to an array of assets: human capital, capital holdings, land, natural resources and any other asset that can be used as a production factor and yield a monetary return. Each agent lives a single period and earns the wage of his unskilled labor and the wealth he was bequeathed, then he has a random marginal propensity to consume and bequests to his sole descendant. He uses the Lorenz curve to study the personal income distribution and concludes that long-run inequality is determined by the interest rate. The intuition is that in a steady state all assets will receive the same return as physical capital. Thus, a decrease in the interest rate diminishes the relative importance of wealth with respect to unskilled labor and improves the distribution of income.

The impact of trade on personal income distribution will depend on the behavior of the interest rate after trade liberalization. To model trade, he uses a dynamic specific-factors model. He concludes that trade liberalization in a land-abundant country will increase inequality and decrease it in a labor-abundant country. This result can be used to explain the empirical evidence that land-abundant Latin America has faced an increase in inequality after trade liberalization. Furthermore, when trade liberalization is complemented with capital mobility, the previous results are counterbalanced and in the long run the country's inequality is equal to world inequality.

Political and Institutional Explanations. The last group of contributors pays more attention to the direct effect of political and institutional interactions with the ownership matrix. They do not focus on how the factor endowments affects S, but instead, they analyze the changes that trade liberalization and globalization create on the political and institutional conditions of the country and how this affects income distribution through S. This is a more heterodox point of view, but cannot be disregarded, given the importance of institutional conditions in the distribution of income. When the government is introduced to the equation, it is obvious that institutional changes are important in determining income inequality –through redistribution policies, market interventions and labor market regulations.

First, a direct link is provided by the way in which the government redistributes tariff revenues. In our reference framework, we have not explicitly dealt on how this can be done, although one can assume that it is redistributed in a way in which distribution is not altered. However, once the government is in the equation, such an assumption must be revised. In many developing countries tariff revenues are important income source for the government and thus, its redistribution can have a measurable impact on income inequality. In addition, trade liberalization significantly reduces this income source and the previous benefactors are hurt.

Another point of view is voiced by Cornia (1999), who presents some economic and specially, non-economic analysis. He states that there has been a recent surge in inequality worldwide[15] and that this episode seems to be related to changes in economic policies and ideology, which reflect a shift towards liberalization and globalization. He gives a number of stabilization and structural adjustment policies, related to the "Washington Consensus" that may have caused this increase. Among others, financial liberalization, the rise of financial rents, privatization and the distribution of industrial assets, changes in labor market institutions, and changes in the tax and transfer system that have eroded the redistributive role of the state.

Accordingly, Robinson (2000) stresses the institutional and political determinants of inequality and how globalization interacts with these. He numbers three main relationships between globalization and income distribution. First, the economic behavior that determines factor prices (i.e. the Stolper-Samuelson effect). Second, he considers that property rights may be treated as endogenous and argues that globalization affects the income distribution by altering the ownership of assets. Third, he says that globalization influences the equilibrium of institutions and policies at a given time.

He presents historical and political evidence suggesting that the impact of globalization on inequality depends much on the institutional and political situation of each country and region. He concludes that political institutions primarily determine income distribution in Latin America, while trade and globalization play a complementary role to this, but are not the main influence. In other words, the ownership matrix and the interactions created by private and public interactions play a crucial role in the whole analysis. Implicit in his argument, is the fact that some production factors, such as land and natural resources,

[15] He uses a 77-country database and some national and regional studies. However, as mentioned above, Sala-i-Martin (2002a,b) presents a more convincing methodology, which contradicts the perception that global inequality is increasing.

have large non-competitive profits that create political and institutional pressures to alter the ownership matrix.

Finally, there is a sizeable literature concerned with the endogenous determination of tariffs.[16] Under this setting, the interaction of political institutions with trade policy creates an integral relationship between both variables. Thus, the political institutions influence on income inequality through tariffs can be increased or tampered, depending on the specific political characteristics of each country.

Of course, such type of relationships can include other policies which also affect trade and inequality. For instance, competition policy and the political and legal issues concerned with FDI and MNEs. It suffices here to acknowledge such relationships and the potential impact it can have on the outcome of trade policy on the distribution of income.

2.3.1.3 Market structure issues

An important critique to the traditional HOS model has been its inadequate explanatory power in some areas. As explained by Helpman and Krugman (1985) standard trade theory has failed to account for empirical evidence in developed countries in four main issues: trade volumes, trade composition, the volume and role of intra-firm trade and direct foreign investment, and finally, the welfare effects of trade liberalization. They argue that these theoretical shortfalls can be explained by incorporating increasing returns and imperfect competition in the analytical framework. In this way they explain the large trade volumes among countries that are relatively similar, two-way trade in goods with similar factor intensity, the little resource reallocation that some trade liberalization episodes have created and the positive welfare effects of increased productivity due to trade.

Although they only mention one interaction between economies of scale, trade and income distribution, their analysis suggests a large scope of possible interactions. In the following paragraphs we detail some of the arguments presented in the literature.

First, we must remember that the zero-profit condition is critical to the unequivocal results of the Stolper-Samuelson theorem. The introduction of imperfect competition and

[16] See for example, Grossman and Helpman (2002).

other market structure issues may render this condition unattainable. Accordingly, the results of the theorem may not hold.

Economies of scale

We can distinguish two approaches to model scale economies. The "old" method uses economies of scale that are external to the firm but internal to the industry. In this setting, we can still work with perfect competition. The "new" approach works instead with imperfect competition, since economies of scale are internal to the firm.

With the "old" approach it is straightforward to observe that economies of scale can reverse the Stolper-Samuelson results. An increase in production raises the marginal productivity of both factors and hence, both factor prices can increase (Panagariya, 1980). However, the final effect on the real factor returns depends on the parameters of the general equilibrium model.

Using the new approach, Krugman (1981) and Helpman and Krugman (1985) employ economies of scale to give an example of the influence of trade on the distribution of income. In general, the Stolper-Samuelson theorem is still at work, but the introduction of economies of scale and product differentiation adds additional effects. The final outcome over the relative factor payments depends on three aspects (Oliveira-Martins, 1994):

- Initial relative factor endowments
- The degree or extent of economies of scale
- The elasticity of substitution between varieties (both domestic and imported) of a given product

The first point is the familiar Stolper-Samuelson result. The bigger the endowment differences, the volume of trade will increase and each country will specialize more in the good which intensively uses the abundant factor. The second and third effects are consequence of the incorporation of economies of scale and may reverse the first effect. As the economy moves towards free trade, the market demand will increase and economies of scale can be exploited, increasing the returns to both factors. The third effect, which measures the preference for product diversity, determines the level at which the economies of scale may operate.

Krugman (1981) finds that if products are sufficiently differentiated, both factors improve their wages. Moreover, this result also holds when countries have a sufficiently similar initial factor endowment. Through these findings, he explains the lack of distributional effects produced by North-North trade liberalization.

Accordingly, Oliveira-Martins (1994) observes that there is a trade-off between economies of scale and product diversity. When there is strong product differentiation, there is less competition and it is easier to exploit economies of scale. As products tend to be more homogenous, competition increases and potential gains from scale economies are hampered. He also argues that intra-industry trade does not substantially affect factor wages. Instead, it is inter-industry specialization which causes the Stolper-Samuelson results.

Based on this theoretical framework, Oliveira-Martins empirically tests the relationship between trade and wages in the OECD countries. He finds that industries with low product differentiation have a higher import penetration. Besides, the wage differential is bigger than in industries with high product differentiation. As expected, industries with strong differentiation usually have scale economies and higher industrial concentration. In this example, market power and economies of scale tamper and may reverse the Stolper-Samuelson results.

Pro-competitive effects of trade

When firms in a specific sector have market power, they are producing less than in the perfect competition benchmark. If trade increases competition in the sector, output increases and this benefits the wages of the relatively abundant factor (Bhagwati and Dehejia, 1994; Richardson, 1995). Additionally, the non-economic profit generated in non-competitive industries can be distributed in different ways. Mainly they are taken by the capital owners, but in highly unionized sectors, it can be shared with the workers. Both issues affect relative wages and income inequality.

Hanson and Harrison (1999) find evidence that the trade reforms in Mexico and Morocco reduced the profit margins in both countries. Currie and Harrison (1997) also reported this result for Morocco. Thus, the reforms brought pro-competitive effects to these economies and under these conditions, factor prices share the adjustment strain together with profit margins. In Morocco it was purely profit margins which decreased, leaving wages unaltered. For Mexico, the changes came mainly in a reduction in the wages of protected and highly unionized workers.[17]

A similar argument was advanced by Borjas and Ramey (1995). They based their explanation on the statistically strong correlation between the wage differential and the imports of durable goods in the United States, the fact that workers in the durable goods industries were mainly unskilled and highly unionized, and that there is a high concentration and profits in these industries. Before trade liberalization and the increase of imports (or decrease in exports), the owners shared the high profits with the workers, through negotiations with the unions, and the workers had above-average wages. With the increase in trade, profits decreased, as well as the wages for unskilled workers. This explanation also allows the inclusion of the effects of unions on the distribution of non-competitive profits, which was stressed by Freeman (1991) as an important wage determinant.

However, this analysis has been criticized and has important shortcuts. The relative weight of the workers from the durable goods sector with respect to the total working force does not seem to be high enough to explain the whole behavior of wages in the United States. In addition, there is a relatively small percentage of unskilled labor force that is unionized in the manufacturing sectors in the United States (Bhagwati, 1998). Furthermore, Lawrence and Slaughter (1993) do not find a decrease in the premiums paid to unionized workers in manufacturing sectors. Nevertheless, this is an interesting market structure approach to the problem and stresses the importance of profit margins for the distributional effects of trade.

Oligopolistic competition

More recently, Neary (2001) uses an oligopolistic competition model to explain the wage differential episode. He assumes that there is a domestic and a foreign firm, which compete in two stages: first they decide their investment levels and then the production quantities. With trade liberalization, the domestic firm is no longer protected and both firms engage in aggressive investment to cut down the other firm. Thus, competition fostered by trade

[17] This is consistent with the findings of Revenga (1997).

liberalization creates investment incentives for both oligopoly firms. Provided that this investment is skill-intensive, the relative demand for skilled workers will increase and this will widen the wage differential. The results of this model are consistent with the three stylized facts of the wage debate, which are presented in the corresponding section below. Furthermore, this model provides a promising use of imperfect competition models to explain the relationship between trade and income distribution.

Incorporating market structure into the theoretical reference framework

By allowing non-competitive profits, we can enhance equation (2.11). Using equation (2.8) we include a non-competitive profit function vector (π) that is determined by firm markup levels (m). For instance, $\pi = \sum_n \pi_n(m)$, with $\pi_n(0) = 0$ and $\pi'_n(m_n) > 0$; where $n = 1, \ldots, N$ is the number of firms in the economy. Then we have:

$$W = P^w(E^w) \cdot \ T \cdot \ \phi \cdot \delta\pi \tag{2.12}$$

where δ is a vector of profit distribution that represents the bargaining power or institutional mechanism through which each factor owner receives part of the firms' non-competitive profits. Among others, this includes the presence and power of unions, and labor and capital market institutions. However, it excludes specific factors of production that may be responsible for the firms' profits (i.e. patents, copyrights), which will be incorporated in the factor productivity vector ϕ and will have an independent return.

Assuming that the profit function is identical across industries and countries, income distribution will be given by:

$$g(Y) = g(E, E^w, T, S, \phi, \delta, m) \tag{2.13}$$

When we incorporate economies of scale, the factor productivity will be a function of the economic activity of the firm λ_n, the production function and endowments: $\phi_n = \phi(\lambda_n, F_n(E_n))$, with $\frac{\partial \phi_n}{\partial \lambda_n} > 0$, when there are increasing returns to scales and $\frac{\partial \phi_n}{\partial \lambda_n} < 0$ with decreasing returns. Under imperfect competition, equation (2.9) changes to:

$$W = W(E^w, E, T, \phi, \delta, m, \lambda) \tag{2.14}$$

With overall increasing returns to scale we have $\frac{\partial W}{\partial \lambda} > 0$; and $\frac{\partial W}{\partial \lambda} < 0$ with decreasing returns. Incorporating both imperfect competition and economies of scale gives:

$$g(Y) = g(E, E^{w}, T, S, \phi, \delta, m, \lambda) \tag{2.15}$$

By reasons previously explained, equation (2.15) is still an incomplete attempt to describe how income distribution is determined. For our purposes, it theoretically incorporates market structure issues in the link between trade and income inequality.

Multinational enterprises

Multinational firms also play an important role in the link between trade and income inequality. On the HOS setting firms are not explicitly treated, but there is much intra-firm trade in world markets and MNEs are important international trade actors with an ever increasing influence (Helpman and Krugman, 1985). Since the 1960s, MNEs began to penetrate domestic markets in a global scale and strategic alliances among firms have increased trade in technology (Richardson, 1995). According to Sachs and Shatz (1994), the empirical evidence points to a significant relation between trade volumes and the presence of MNEs in developing countries.

Outsourcing activities conducted mostly by MNEs is one of the main explanations employed to understand the recent episode of wage dispersion. Two papers by Feenstra and Hanson (1995a, 1995b) study the trade reforms in Mexico and find evidence of an increase in the wage differential. Moreover, this rising wage gap is associated with changes internal to industries and even internal to plants. They explain this phenomenon by introducing Foreign Direct Investment (FDI) and the outsourcing of local production by MNEs. They depart from the HOS model by introducing intermediate goods and the full specialization of countries in some goods. In this framework, without factor price equalization between countries, the North specializes in inputs that are relatively intensive in skilled labor and the South specializes in inputs that are relatively intensive in unskilled labor. Then the flow of capital provided by the MNEs shifts some unskilled-intensive production (from the viewpoint of the North) to the South, were it is regarded as skilled-intensive. This mechanism increases the wage differential in both regions.

They also find that FDI growth is positively correlated with the relative demand for skilled labor, which supports the hypothesis that outsourcing by multinationals has been a significant factor in the increase in the relative demand for skilled labor in Mexico and they expect this hypothesis to hold for other countries, as well. These papers give support to the idea that MNEs are playing an important role in the functional distribution of income, through an increase in the relative demand of skilled workers in both developed and developing countries.

Closely related to the outsourcing hypothesis, is the skill-enhancing-trade hypothesis proposed by Robbins (1996b). He argues that trade liberalization accelerates capital inflows and technology transfers from the North to the South. As a consequence, these inflows increase the productivity of skilled labor and this widens the wage differential in developing countries.

Markusen and Venables (1996) arrive at similar conclusions. They adapt a formal model of MNEs and scale economies to address distributional effects. They find that the increased participation of MNEs can raise the wage differential in the North and, when the endowment differences are big enough, also in the South. In essence, MNEs change the trade flows from final goods to intermediates and services. This change breaks the factor price and final goods price linkage central to the Stolper-Samuelson results. Accordingly, multinationals export firm-specific knowledge-capital and their production is fragmented between skilled and unskilled-labor-intensive activities in national and foreign plants. Besides the wage differential result, they also find that MNEs may export low skilled jobs to the South and that trade barriers do not protect unskilled labor in the high-income countries.

Although the exact effect of outsourcing is open to debate in the North[18], it is still regarded as an important factor that can explain wage differentials.

To summarize this section, we found important theoretical arguments to take into consideration when analyzing the effect of trade on income inequality. The Stolper-Samuelson remains the best analytical tool available to examine the links. However, there are important theoretical reasons that can invalidate the Stolper-Samuelson results. In particular,

[18] We discuss this further in the wage inequality debate section.

technological change and outsourcing have received special attention in the literature. Additionally, caution must be used when shifting from the functional to the personal distribution of income. In particular, the importance of country-specific and political variables to determine household inequality must be controlled when conducting empirical tests.

In the following sections we turn our attention to the empirical studies that test the link between trade and inequality. First, we give a review of the trade and wage inequality debate, which is focused on the functional distribution of income. Later, we present the empirical results that directly test trade liberalization against the personal distribution of income.

2.3.2 The trade and wage inequality debate

There is a vast literature analyzing the effects of trade on wage inequality. In this context, some of the studies have been used to test the Stolper-Samuelson theorem and the effect of trade on the functional distribution of income. The surge of studies was encouraged by the sharp wage inequality increase in the United States that began in the 1980s (Bound and Johnson, 1992; Katz and Murphy, 1992).

As exposed by Feenstra and Hanson (1995b) while there is a general agreement that the wage inequality increase was caused by a shift in the relative demand for skilled labor, the sources of this shift were controversial. On one hand, some argued that skilled-biased technological change had increased the demand and the returns to skilled labor.[19] Alternatively, some authors claim that trade liberalization and increased North-South trade have been responsible.[20] This trade argument relied, in general terms, on the Stolper-Samuelson theorem: the unskilled workers in the North will have a relative wage decrease with respect to the relatively abundant skilled workers. With respect to our theoreti-

[19] Leamer (1994) argues that it is sector-biased technological change which raises the relative productivity of skilled labor. However, Krugman (1995a) forcefully demonstrates that when the small-country assumption is dropped and technological change affects international prices, then it is factor-biased technological change that matters.

[20] It is important to remind that trade *per se* is an endogenous variable in a general equilibrium trade model. Thus, the trade explanation is based on the change of an exogenous variable that induces more North-South trade. The possibilities include tariff reductions and other trade policy instruments, drops in transport costs, shifts in foreign tastes and technology (Deardorff and Haikura, 1994; Slaughter, 1998).

cal framework, the skill-biased technological change will be represented by a change in the vector ϕ and the trade argument is represented by shifts in tariffs and the initial endowment vectors (T, E, E^w).

2.3.2.1 OECD countries

As mentioned before, the main debate has focused almost entirely on the experience of the United States. However some studies have included other OECD countries which also present an increase in the wage differential or were unemployment has risen. We analyze first the methodological debate and afterwards, the empirical evidence.

Theoretical and methodological issues

Besides the technology change and trade arguments, there were some alternative explanations to the wage inequality rise. Immigration was mentioned to have increased the supply of unskilled labor and decreased the wage of these workers. Thus, immigration was complementing the trade effect (Borjas 1994; Borjas, Freeman and Katz 1992; Williamson 1998).

Others blamed the labor market institutions. The OECD wage data showed that those countries with more flexible labor markets had greater wage inequality. One of these issues was the decentralization of wage determination and the power and coverage of unions (Freeman 1991).[21] While the United States and the United Kingdom had greater wage inequality, continental Europe had greater unemployment. This suggests that the labor market institutions affected the final outcome of the labor market shocks, but did not generate these shocks (Wood, 1995; Neary, 2001).

At the end, labor supply issues and labor market institutions were left out of the main debate, which was centered on the skilled-biased technological change and trade liberalization arguments. Both were regarded as key influences, but the magnitude of each became a critical part of the debate. In particular, the concern was on how to

[21] He acknowledges, however, that deunionization was not the critical factor explaining the surge in wage inequality. In addition, Davis and Haltiwanger (1991) and Bound and Johnson (1992) do not find a significant role for deunionization.

estimate the impact of both phenomena. This has created an analytical dispute that has been implicit in most of the literature.

In general terms, one can observe a confrontation of the labor economists (using mainly partial equilibrium analysis and econometric methods) with trade economists (using mainly general equilibrium analysis and the HOS trade model). Several authors convincingly point out that a general equilibrium approach is the appropriate tool to tackle the problem.[22] Nevertheless, this may complicate the empirically testing and hence, partial equilibrium and an econometric approach can also yield useful insights (Richardson, 1995).

The starting point of the debate has been the correlation of trade volumes, trade deficits and wage inequality. Labor economists have used the factor content of trade analysis to link the wage dispersion with trade. This analysis looks at the factor intensities of exports and imports. In the case that the skill to unskilled labor intensity in exports is higher than in imports, then increased trade volumes are associated with an increased demand for skilled workers. Hence, increased trade widens the wage differential. Complementing this insights with the increased trade volumes and trade deficits in the United States since the 1980s, has lead some authors to state that the wage differential had been caused by trade liberalization.[23]

However, this approach has been criticized on several grounds. First, it lacks a clear theoretical framework to link the mentioned variables and this qualifies any conclusion drawn from these simple relationships. In fact, there is an implicit flaw in the analysis, which is easily observed using the HOS trade model. For the Stolper-Samuelson effects to be in play, the relative price of skilled and unskilled labor intensive goods must change. Trade volumes are in this case an endogenous variable and can be affected by several factors and thus, the behavior of such an endogenous variable does not establish causality.

Therefore, the factor content of trade analysis, while looking at trade volumes and not at relative prices, has been regarded as an invalid method to explain wage inequality caused by trade.[24] Nonetheless, Krugman (1995) argues that the approach is valid when

[22] Richardson (1995), Bhagwati (1998) and Francois and Nelson (1998).
[23] Borjas, Freeman and Katz (1992), Murphy and Welch (1992), and Wood (1994, 1995).
[24] Lawrence and Slaughter (1993), Bhagwati and Dehejia (1994), and Bhagwati (1998).

dealing with small trade shares, as is the case in North-South trade. In this context, using the factor content of trade can answer the question on how the increased volumes of imports from the South changed the wages in the North, compared to the case where trade had not occurred.

Further criticism to the labor economists come from two additional results of the HOS model (Leamer, 1993; Robbins, 1996b). Labor economists often net out relative supply shifts. However, labor supply changes do not affect relative wages and demand shifts induced by trade are directly observable in wage changes. Secondly, exchange rate variations also leave relative wages unchanged. This observation invalidates those studies that focused on the dollar revaluation (i.e. Murphy and Welch, 1992; Revenga, 1992).

Empirical evidence

In this section we turn our attention away from the methodological issues and deal with the empirical evidence. The relevant empirical facts can be summarized as follows:

- the relative price of unskilled-intensive goods has not declined significantly[25]
- the ratio of skilled to unskilled employment has increased in all sectors (not only in the skill-intensive sector)
- the wage differential has risen in both North and South

Using the first fact in combination with the HOS model, several authors conclude that trade does not have much influence on the wage gap.[26] Moreover, if trade was the cause of the wage dispersion episode, then the unskilled to skilled employment ratio must decrease. Given the second empirical fact, this is not the case.[27] Lastly, the Stolper-Samuelson effect should reduce the wage dispersion in developing countries, not increase it.[28]

The first fact was initially presented by Lawrence and Slaughter (1993). They claim that the relative price of investment goods, which are taken to be skill-intensive goods,

[25] As we illustrate below, this is still disputed by some authors, but is generally accepted.
[26] Lawrence and Slaughter (1993), Bhagwati and Dehejia (1994), and Bhagwati (1998).
[27] Krugman and Lawrence (1994).
[28] We analyze further this empirical fact in the following section.

was falling instead of increasing. However, Sachs and Schatz (1994) refute this conclusion and claim that the price of skill-intensive goods was driven by computer prices.[29] Once these are removed, the relative price is increasing.[30] Yet, they do not find that the fall in the relative price of unskilled-intensive goods is large enough to account for the significant wage dispersion.

This last point can be explained by another argument of labor economists, i.e., that the smooth transition of labor from one sector to the other, implicit in the general equilibrium approach, is not reasonable. They remark that there are adjustment costs and that these costs can cause wage differentials. For instance, skilled labor can shift more smoothly to another productive sector and obtain earnings similar to those he enjoyed before. This may not be the case for unskilled workers. Therefore, a small change in relative prices can cause a significant change in the wage differential in the short run (Richardson, 1995). For instance, Slaughter (1998) acknowledges this problem and suggests that such labor-market frictions can create a lag in the price-wage relationship that can seriously affect the empirical testing.

On the other hand, the biased technology explanation is consistent with both an increase in the wage gap and constant final goods prices. Leamer (1994) and Richardson (1995) demonstrate how in a general equilibrium setting, trade prices and biased technological change are the causes of factor price changes in the long run. Empirically, Lawrence and Slaughter (1993) find a significant and positive relation between the growth of total factor productivity (TFP) and the intensive use of skilled labor, which further supports the biased technological change hypothesis.[31]

However, there is also an important shortcoming with the technological change explanation. As the second empirical fact shows, the skill intensity in both sectors has increased and this is incompatible with the skilled biased-technological change argument.[32] In an

[29]The price of computers fell sharply due to extraordinary productivity increases, which are hard to measure and thus, it is hard to obtain the "effective price", i.e., the net effect of price changes and total factor productivity changes. Nonetheless, since the introduction of personal computers has been mentioned as a critical part in the recent technological spell, it is a crucial decision to exclude them.

[30]This last result depends on the price definition used and in some cases the changes are not statistically significant.

[31]Although this is again challenged by Sachs and Shatz (1994), which find they TFP changes actually favored low-skill labor.

[32]See Neary (2001) for a thorough exposition.

attempt to account for this fact, Francois and Nelson (1998) use a combination of technological change that is both sector and factor biased. They show that in this case, the skilled-intensive sector becomes more skilled intensive while the other sector is more unskilled intensive. Yet, the empirical evidence suggests that both sectors became more skill-intensive and their explanation is not satisfactory. Moreover, Leamer (1994) argues that technological change has in fact decreased the wage differential. In general, there are several limitations on the empirical testing of the nature and magnitude of technological changes. These limitations make it very difficult to check if technological change has in fact been the culprit (Deardorff, 1998).

Alternatively, some authors combine both effects to explain the wage gap increase and the changing ratio of skilled to unskilled labor.[33] On the other hand, Neary (2001) disregards both the trade and technical change arguments as the causes of the wage dispersion. He argues that none of the explanations comply with the three stylized facts and he uses his oligopolistic competition model (discussed above) as an alternative explanation.

To qualify the impact of trade, Krugman (1995a, 1995b) focuses on the relative importance of North-South trade. Using the fact that manufactured imports from newly industrialized economies (NIEs) to OECD countries has only increased from almost zero to 2%, he assess a significant but rather small impact of trade on factor prices.

Finally, the introduction of outsourcing and MNEs has also played an important role in the debate. As explained above, Feenstra and Hanson (1995a, 1995b) have argued in favor of outsourcing as an additional trade induced mechanism and an important determinant in the wage episode. Nonetheless, outsourcing has been downgraded as an explanation of the wage differential in the United States (Lawrence and Slaughter, 1993; Berman, Bound and Griliches, 1994; Slaughter, 1995). One of the reasons is that the magnitude of outsourcing is too small to account for the wage differential increase. Other reason is that South unskilled labor is apparently a weak price substitute for American unskilled workers (Slaughter, 1995). However, Feenstra and Hanson (1995a, 1996 and 2004) observe that

[33] Sachs and Schatz (1994), Leamer (1993).

when a broader definition of outsourcing is taken[34], it can explain an important share of the wage changes.

Going one step further, Feenstra and Hanson (1997) use a regression linking industry prices and wages to disaggregate the effects of trade and skill-biased technological change.[35] Trade is taken as the broad outsourcing definition introduced before and the respective technical change as expenditures in high-technology capital.[36] Their estimations predict correctly the wage changes and they find that trade has a significant effect, but technological change is the most important factor to account for the increase in the wage differential.[37] However, Slaughter (1998), surveying this and other studies that used the same technique, concludes that the results are sensitive to the exact estimating strategy.

We conclude this section with Table 2-1. Here we use an arrangement from Cline (1997) to summarize and frame each position in the debate.

The general perception is that the debate and empirical evidence has been favorable for the skilled-biased technological change explanation (Richardson 1995; Cline 1997). Trade liberalization is accepted to have had a significant influence on the wage inequality increase, although the magnitude seems to be relatively small compared to the effect of the skill-biased technological change (Slaughter, 1998). However, when complemented with the outsourcing argument, trade does seem to have a stronger effect.

For our purposes, the debate is corroborating the Stolper-Samuelson theorem. Although the weight attributed to trade in the increase in the wage differential is relatively small, it is a significant influence and it has the expected direction.

2.3.2.2 Developing countries

Most of the empirical evidence for this group of countries, especially from Latin America in the 1980s, reports different results from those expected using the Stolper-Samuelson

[34] One that includes all imported intermediates or final goods, which are used to produce or are sold under an American brand.

[35] They regress the change in the industry price on the level of factor shares, where the estimated coefficients are interpreted as the expected wage changes.

[36] Of which computers is the only significant component.

[37] They reach this result when they adjust industry prices to obtain the "effective price".

Table 2-1: Overview of the trade and wage inequality debate

	Labor Economists	***Trade Economists***
Trade	Borjas, Freeman & Katz (1992) Murphy & Welch (1992) Revenga (1992) Borjas & Ramey (1995) Burtless (1995)	Leamer (1993, 1994, 1996) Sachs & Shatz (1994, 1996) Wood (1994, 1995, 1998) /1 Feenstra & Hanson (1995a/b, 1997) /2 Deardorff (1998) Francois & Nelson (1998)
Technology (no trade)	Mincer (1991) Davis & Haltiwanger (1991) Bound & Johnson (1992) Berman, Bound & Griliches (1994)	Lawrence & Slaughter (1993) Bhagwati & Dehejia (1994) Deardoff & Hakura (1994) Krugman & Lawrence (1994) Krugman (1995a/b) Baldwin & Cain (1997) Bhagwati (1998) Desjonqueres *et al.* (1999)

Notes: The 'trade' row includes papers giving a moderate and high importance to trade, while the 'technology' row includes papers favoring skilled-biased technological change and give a low trade effect.
1/ Wood is a development economist
2/ These authors refer to trade as outsourcing by MNEs and FDI.

theorem. This has prompted Davis (1996) to call the case an anomaly. Nevertheless, it is important to keep in mind that developing countries are more difficult to study, as exposed by Robbins (1996b):

> "Less developed country studies often face greater data limitations, less perfect understanding of non-trade influences on their labor markets, and the difficulty of controlling for important non-trade changes."

A first attempt to analyze economies outside the OECD was made by the World Bank (1987). They analyzed the cases of Singapore, Hong Kong and Korea and reported that the theorem was corroborated. With this evidence they encouraged trade policy reforms in developing countries during the 1980s.

However, the most studied developing country case has been Mexico (Feenstra and Hanson 1995b; Hanson and Harrison, 1995, 1999; Feliciano 2001). These studies found evidence that trade liberalization has in fact increased instead of declined wage inequality, as was expected from the Stolper-Samuelson theorem. However, Hanson and Harrison

(1999) argue that the decline in tariffs was highest in the unskilled labor intensive sectors.[38] Under these conditions, the Stolper-Samuelson theorem holds, since the price of unskilled labor intensive goods is reduced and this diminishes the wage of these workers. In particular, these sectors were highly protected before the reforms and had significant profit margins in the domestic markets. The tariff reductions increased exports and reduced the profit margins. Outsourcing from MNEs was another explanation common to all these papers related to the Mexican experience. As explained above, the capital inflow that FDI brings, has been described as skilled-biased and thus, can also explain the increase in the wage dispersion.

Robbins (1996a, 1996b) studies nine Latin American and East Asian countries[39] and reports that wage inequality rose after trade reforms. In addition, he finds that labor supply shifts and real exchange rates do affect relative wages, in contradiction to the HOS model. Therefore, he nets out these supply changes on wages to identify the demand shifts. To explain the apparent Stolper-Samuelson anomaly, he employs his skill-enhancing-trade hypothesis. He supports this claim by finding a positive and significant relationship between his estimated labor demand variable and the imported capital stock.

Wood (1997) separates both regions and claims that in Latin America the theorem fails empirical evidence, but is supported by the East Asian experience. He introduces land (also associated with natural resources) as a third production factor and employs it to explain the different inequality outcomes of the trade liberalization episodes in Latin America and Southeast Asia. He notices that Latin American countries are relatively land-abundant, while Southeast Asian countries are relatively labor-abundant and the developed North is capital-abundant. However, he does not find any strong empirical evidence to support that this endowment differences can explain the inequality results. Instead, he suggests another plausible reason of the different outcomes in both regions. The 1960-70s, when Southeast Asian countries began to liberalize trade, was different from the 1980-1990s, when Latin America began to liberalize. In the second time period, there was an increase in the world supply of unskilled labor after the rise of trade volumes experimented by China and other low-income Asian countries. As a result, one can use

[38] They also claim that this was the case in Morocco.

[39] Argentina, Chile, Costa Rica, Colombia, Malaysia, Mexico, the Philippines, Taiwan and Uruguay.

an argument related to the diversification cones hypothesis. Thus, this event changed the cone in which some middle-income countries were and their comparative advantages were shifted from products intensive in unskilled labor to products of medium-skill intensity. Subsequent trade liberalization in these countries reduced the relative wages for unskilled workers. Secondly, technical progress between the 1970s and the 1990s was biased against unskilled workers, in accordance with the skill-enhancing-trade hypothesis.

Unfortunately, Robbins and Wood do not analyze changes in the relative price of final goods, nor do they check for the sectoral composition of the tariff reductions. Thus, it is not clear whether the Stolper-Samuelson results are an important influence or if the effect is going in the right direction.

Williamson (1998) introduces economic history into the debate and argues that the previous globalization process in the late XIX century was as big or bigger that the on in late XX century. He finds evidence to support the Stolper-Samuelson theorem in the previous case. He also says that immigration, together with trade liberalization, plays a central role in the wage convergence process.

In a recent report, the World Bank (2001) cites a number of case studies that suggest that openness over the last decade is partially responsible for increasing relative wage dispersion. They argue that relative endowments, previous trade protectionism patterns and technological change can be influencing these results.

With respect to the influence of labor market institutions in developing countries, initial studies did not find a significant influence on the wage inequality results. Harrison and Revenga (1995) claimed that the empirical evidence for Mexico and Morocco suggested that labor market restrictions were not responsible for the labor market response to trade reforms. However, more recent papers have stressed the importance that these institutions exert in developing countries through changes in labor demand elasticities (Rodrik, 1997; Turrini, 2002).

Thus, besides the three effects already identified (e.g. trade liberalization under a diversification cones context, FDI and technology inflows),.labor market institutions becomes an additional factor to explain the increased wage inequality in the South.

Summarizing the studies presented here, there are mix results regarding the evidence on the Stolper-Samuelson results for developing countries. Most studies that disregard the theorem lack a robust specification in the analysis. Those that support it, are few and case specific. In general, there is a lack of representativeness of other developing countries regions (Africa and South Asia) and further research is necessary to obtain a better understanding of the effects of trade on wage inequality in the South.

2.3.3 Trade and the household distribution of income

Contrasting the trade and wages debate, the empirical tests on the effect of trade on the personal distribution are based on cross-country and panel data analysis. In these studies it can no longer be taken as granted that inequality has increased, as was the case in the wage dispersion debate. Now, we must associate trade openness with possible distributional changes and this introduces an additional and important problem: how to measure trade liberalization. Thus, we first deal with this issue and afterwards present the empirical literature that links trade openness with the household distribution of income.

2.3.3.1 Trade openness measurement

There are a number of important analytical and practical problems involved when measuring trade openness. As explained by Berg and Krueger (2003), the main concern is about policies that distort market allocation and there can be many instruments that can achieve this; among others, tariffs and non-tariffs barriers (NTBs), and discriminatory exchange rates. However, one must also keep in mind that there are possible trade policy substitutions that cannot be monitored by a single variable indicator.

Direct policy measures

This approach directly uses some policy measures that restricts trade, such as average tariffs and NTBs. The problem here is how to aggregate meaningfully across goods. Simple averages can overestimate the weight of goods that are relatively unimportant for a country, but have high tariffs (Dollar and Kraay, 2004). Since tariffs affect trade flows, weighting the tariffs by trade volume can also be inappropriate because tariff levels have different effects on goods and moreover, different welfare effects. In addition, there is no necessary relationship between official and collected tariff rates.

On the other hand, NTBs are extremely hard to quantify. Thus, most of the indicators focus on easily quantifiable barriers, but usually exclude the most difficult ones to measure. Additionally, non-tariff coverage ratios do not discriminate adequately between highly and less restrictive barriers (Rodríguez and Rodrik, 2001).

Moreover, tariffs and NTBs can be substituted in a way that low average tariffs can be compatible with high protectionist overall policies. Finally, it is not clear if these direct measures explain changes in trade volumes. For example, some studies have found that both measures have little correlation with trade volumes (Dollar and Kraay, 2004), while Rodrik (2000) does find a significant and negative relationship between average tariffs and import/GDP ratios.

These are important critics and limitations, but there is no ideal openness indicator. Therefore, these direct measures can be informative of trade policy and in contrast with other indicators, provide non-binary values.

Policy-based measures

Another approach is to take into account the possible substitution between measures. Sachs and Warner (1995) try to capture any substitution policies by looking at five different trade policies. They construct a dummy variable for openness based on those countries that pass each of the following policy tests:

- An average tariff rate below 40 percent
- NTBs covering less than 40 percent of trade
- A black market exchange rate premium below 20 percent on average during the 1970s and 1980s
- The absence of a socialist economic system
- The absence of an extractive state monopoly on major exports

If any of the previous criteria is not met, the economy is considered to be closed and the index is assigned a value of zero.

This approach has several drawbacks. First, being a dummy variable one cannot differentiate degrees of openness.[40] Second, some of the index components do not relate directly to trade liberalization. For instance, Hanson and Harrison (1999) remark that the black market premium measures factors other than trade policy. There has also been criticism in the use of the marketing board component and the independence of the tariff and quota measures with regard to the aggregated measure. Finally, Berg and Krueger (2003) claim that the measure can be seen as a product of its time, since there are new policies that restrict trade (e.g. sanitary and technical standards, contingent protection).

To sum up, the Sachs and Warner index is suitable to assess broader economic reforms and the degree of liberalization of the economy, but it does not seem to properly isolate trade policy.

Outcome-based measures

Finally, openness can be measured using trade volumes. The common indicator is the ratio of exports plus imports with respect to Gross Domestic Product, i.e.$\frac{(X+M)}{GDP}$. This is a useful indicator of trade outcomes, but not necessarily of trade policy (Dollar and Kraay, 2004). In fact, this particular measure can be influenced by a number of country characteristics that are not related to trade policy at all. E.g. level of development, country size, geographic factors and resource endowment. An alternative is to regress $\frac{(X+M)}{GDP}$ against these non-policy determinants of trade and take the residual as a measure of openness. However, Berg and Krueger argue that the empirical models used in these regressions are still not sufficiently accurate to identify the residual with policy.

An interesting approach is to use changes in the volume of trade over time (Dollar and Kraay, 2004). In this way, the geographical and country specific variables, which are assumed to be constant, drop out of the indicator. Of course, other economic policy variables can also be changing with time, but at least the geographical ones are not. However, Rodrik (2000) has criticized this indicator for its indirect relationship with trade policy. He argues that policy makers cannot control the levels of trade, only the tariffs and NTBs.

[40] Yet, some authors use the average of the indicator for a given period to obtain a value between zero and one.

Another refinement is to take into consideration the effects of openness on relative prices. An expected increase of manufacturing productivity from trade increases the relative price and relative importance of the non-tradable sectors. This effect can bias the outcome index towards lower "real openness". Alcalá and Ciccone (2004) suggest correcting this bias using variables adjusted by purchasing power parity (PPP).

Summarizing, there is no ideal openness indicator and each group of measures has its own problems and limitations (Edwards, 1998). A strategy to overcome this problem has been to test the robustness of the results using different openness indicators.[41] Nevertheless, the measurement error induced by the openness indicators can be significant and most be considered when conducting empirical research.

2.3.3.2 Empirical studies analyzing the effects of trade on the personal distribution of income

The trade and wages debate was centered exclusively on labor wages and excluded capital and land rents. Thus, a generalization from this partial functional distribution to the personal income distribution may be misleading.

Early studies that focused on the household distribution from the 1960s and 1970s (Bhagwati, 1978; Krueger, 1978) produced mixed results that did not state a clear relationship between both variables. More recently, Bourguignon and Morrisson (1990) and Spilimbergo *et al.* (1999), used the general equilibrium theoretical approach outlined in Section 2.3.1 to test if trade has a significant impact on income inequality. Bourguignon and Morrisson prove the importance of initial endowments on the personal distribution of income. They find that there is higher inequality in land and natural resource abundant countries. This finding is corroborated in most of the following studies. In addition, they conclude that trade protection increases inequality. However, their sample is restricted to small and medium-sized developing countries, it only covers one year (1970) and the quality of the variables they use is dubious.[42] Hence, their results can only be considered as tentative (Fischer, 2001).

[41] Among others, Levine and Renelt (1992); Edwards (1997, 1998); and Greenaway *et al.* (1998).

[42] In particular, the ILO inequality dataset lacks many of the quality criteria implemented later by Deininger and Squire (1996).

Spilimbergo *et al.* significantly improve the empirical settings by using more countries, a 28 year period and better constructed variables. They also find that factor endowments can significantly account for income inequality. Land and capital abundant countries are more unequal, while countries with relative high average skill levels are less unequal. Additionally, when trade is introduced in the equation, factor endowments are still a significant explanatory variable for income inequality. Overall, openness increases inequality, but the effects depend on the initial factor endowments. Trade decreases inequality in capital-abundant countries and increases inequality in skill-abundant countries. These last results directly contradict the Stolper-Samuelson theorem and are rather puzzling, but are later refuted by the other studies. As with the rest of studies that follow, they use the DS inequality dataset, which allows panel data regressions. In addition, these authors construct their own openness indicator, based on an adjustment of the volumes of trade for geographical variables and factor endowments.

Edwards (1997) finds no evidence linking openness to increases in inequality. He pays special attention to the methodological problems of measuring openness. Thus, he uses six alternative openness indexes to test the robustness of his results. Inequality is defined as the average Gini change from the 1970s to the 1980s and it is regressed against the openness indexes and other control variables. Yet, he does not include initial endowments in his regressions and his study is not directly comparable with the previous. Dollar and Kraay (2002) use a similar approach, include endowment variables and do not find a systematic relation between different openness indicators and the income of the poorest 20%. Although they claim that the income share of the first quintile is highly correlated with the Gini coefficient, this introduces additional measurement errors to the already highly problematic data on inequality. This problem is acknowledged by these authors and thus, we do not consider it very informative on inequality, but on poverty.

In a related study, Litwin (1998) uses a sample of developing countries and finds a negative relation between trade openness and inequality. She includes endowment variables and trade composition in the regressions. The endowment of human capital is an important determinant of income inequality and can even reverse the positive openness effects. Trade composition is also important and those countries with relatively more manufactured exports than primary goods exports, experience better income distributions. Thus, she concludes that trade openness can reduce inequality when human capital accumulation is increased. This directly improves the income distribution and indirectly, by changing the composition of trade towards manufactured goods.

Some studies use the Kuznets curve tradition to check for openness effects and hence, include a quadratic GDP specification. Higgins and Williamson (1999) find limited support for the positive impact of trade openness on income inequality. Nevertheless, even when they acknowledge and find empirical significance for factor endowments, they do not jointly regress them with their openness variables. Besides including GDP per worker to check for the Kuznets hypothesis, they also use cohort size as an inequality determinant. They find both relationships significant and robust to the inclusion of other control variables. For instance, they conclude that the demographic factors proxied by their cohort size variable is extremely important to explain inequality differences between countries.

Besides the GDP quadratic specification Barro (2000) includes schooling variables, and dummies for Sub-Saharan Africa and Latin America as important inequality determinants. He finds a significant and positive influence of openness on inequality. Openness is measured by the relative volume of trade filtered by country size, but the results are robust to the inclusion of other indicators. To analyze the impact of endowments, he uses the fact that rich countries are characterized by a relative abundance of physical and human capital and developing countries are relatively abundant in labor or unskilled labor. When openness interacts with GDP, the relationship is negative and thus, this result conflicts with the Stolper-Samuelson theorem. However, the results are not robust to the econometric specification and are valid only when excluding fixed country effects.[43]

Fischer (2001) finds that after trade liberalization income inequality increases. The effect is relatively small and is moderated by factor abundance: in land-abundant countries the increase is higher than in labor-abundant countries. Hence, he can explain the dissimilar income inequality results of trade liberalization in Latin America and East Asian countries. Nevertheless, there are several limitations in his empirical approach: he does not include additional control variables, uses only the Sachs and Warner openness index, and the Gini coefficient data does not comply with the quality standards of the Deininger

[43] He uses seemingly unrelated regressions (SUR) in this case. Fixed-effects estimations do not show significant coefficients for the openness and interactive openness-GDP variables.

and Squire dataset. Additionally, he finds that labor-abundant countries are less unequal, a result that is not supported by other empirical studies.

In a related study, Lundberg and Squire (2003) also find a significant and positive relation between the Gini coefficient and the Sachs and Warner openness index. They use a three-stage least squares (3LS) technique, arguing that the use of fixed effects estimation does not exploit the significant between-country inequality differences. However, they are not directly concerned with the effects of openness and state that the Sachs and Warner index is not appropriate when exclusively focusing on trade liberalization. In addition, they use a large set of control variables, which are chosen to reflect determinants of growth and inequality.

Finally, we survey a group of studies done by World Bank researchers. These set of papers, however, where not intended to assess directly the effect of trade openness on inequality, but anyway, present interesting results. Dollar and Kraay (2004) regress changes in the volume of trade with respect to GDP, against changes in the Gini coefficient. They control for changes in average incomes and some macroeconomic indicators.[44] They do not find any significant relationship between changes in trade and inequality, but they do not include any endowment or demographic variables.[45] Analyzing country case studies, the World Bank (2001, 2002) finds mixed results in inequality experiences in countries that have performed significant outward oriented reforms. In particular, they find decreasing inequality in the Philippines and Malaysia and increasing inequality in most Latin American countries and China. Nevertheless, they refer to countries which have integrated sharply into the world economy and they do not explicitly separate trade liberalization from other outward oriented and domestic market liberalization reforms. The limitations in both the regression and case study analysis from this group of papers obscure the role of trade liberalization in any inequality episode.

Surveying the literature on sectoral and micro studies, Berg and Krueger (2003) conclude that there is no systematic relationship between openness and income inequality, when growth is not included.

[44] Government consumption, inflation, financial development and the rule of law.

[45] They use the DS data set and check the robustness of the results using different income and openness definitions.

Table 2-2: Survey of empirical studies on trade and income inequality

Authors	ID Data	Sample	Dep. var.	Trade var.	Control var.	Results
Bourguignon &Morrisson (1990)	ILO	small & medium LDCs only 1970	Income shares	Dummy for efective protec. in manuf. sector >30%	GDP per capita factor endowments ownership var.	Endowments are important Openness decreases inequality
Edwards (1997)	DS	43 countries Two decades: 70s and 80s	Gini coefficient: Decadal average difference	Seven indexes & Reform dummy	Schooling Macro var.	Schooling reduces inequality No openness effect No trade reform effect
Litwin (1998)	DS	only LDCs	Gini coefficient Income shares	(X+M)/GDP Sachs &Warner	Endowments Kuznets Trade composition	Endowments are important Openness increases inequality, conditional on endowments.
Spilimbergo et al. (1999)	DS a/	34 countries 320 obs 1965-1992	Gini coefficient Income shares	Own constructed index Alternative indexes	GDP per capita factor endowments ownership var.	Endowments are important Openness increases inequality, conditional on endowments.
Higgins & Williamson (1999)	DS b/	73 countries 219 obs Four decades: 1960-1990s	Gini coefficient: Decadal average Quintile ratio gap	Sachs &Warner Alternative indexes	GDP per worker Rich/Poor dummy Cohort size	Demographic factors important Support for cond. Kuznets-cuve Openness reduces inequality, but the effect is weak
Barro (2000)	DS b/	Var. sample Three decades: 1965-1985	Gini coefficient: Decadal average Income shares	Own constructed volume index Alternative indexes	GDP per capita Schooling Regional dummies	Openness increase inequality and inverserly with GDP size Support for cond. Kuznets-curve
Fischer (2001)	Barro &Lee	Five year averages 1965-1995	Gini coefficient	Sachs &Warner	Factor endowments	Openness increases inequality, but it is dependent on endowments

Notes: a/ They drop 10 observations, so within country is either expenditure or income-based.
b/ They use dummy variables to correct for differences in income and reference unit definitions

A summary of the results is presented in Table $2-2$. Although the results are mixed, more studies conclude that trade increases inequality. However, the relationship is conditional on factor endowments and other economic variables (e.g. GDP per capita, schooling and cohort size). This outcome conditions the results of those studies that do not control for these variables and makes it a necessary step in future empirical tests. In any case, the major setback of the previous studies relies on the inequality dataset used. Some studies use inequality data that does not comply with the DS quality criteria. Furthermore, the DS dataset has been recently criticized by Atkinson and Brandolini (2001) and this qualifies the results based on this dataset.

Therefore, there is a major scope for improving the empirical testing. The introduction of a new income inequality dataset seems necessary to check the results drawn from the DS dataset. Additionally, the specification of the empirical tests can also be improved.

2.4 Summary

This survey has documented the main debates and findings of the literature that studies the effects of trade liberalization on income distribution. This research field is characterized by the presence of controversial topics, some of which are still part of ongoing academic debates. In addition, much of the empirical research is characterized by important methodological problems and data limitations. However, from the large amount of studies that have been published, we can highlight some important observations.

The main theoretical results linking trade openness and income inequality are derived from the Stolper-Samuelson theorem. This extension of the HOS trade model predicts that trade liberalization in rich capital-abundant countries deteriorates the real income of workers and benefits capital-owners. In poor labor-abundant countries these results are inverted. These predictions are used to determine the possible influence of import tariff changes on the functional distribution of income. Moreover, these basic results are usually generalized to encompass broader trade liberalization policies. Furthermore, assuming that labor is equally distributed and more abundant in developing countries, the Stolper-Samuelson results can also be used to predict the influence of trade liberalization on household inequality.

However, the Stolper-Samuelson theorem is based on simple and strong assumptions. When some of these assumptions are relaxed or changed, the main results no longer hold. Therefore, many studies have focused on testing the theorem's predictions. Given the large amount of empirical studies, we divided the discussion between those analyses that dealt with the functional or the household distribution of income.

In the first case, the OECD experience documented in the trade and wages debate seems to support the Stolper-Samuelson results, although other important factors are at play, i.e. technological change and outsourcing activities. In developing countries the experiences are driven by country specific characteristics, which are crucial to understand particular results. The intrinsic heterogeneity between developing countries makes it difficult to discern a clear general relationship. When subdivisions are used, some distinct patterns emerge. For instance, the South East-Asian experience tends to confirm the Stolper-Samuelson results. On the contrary, trade and inequality experiences in Latin

American countries do not follow the theorem's predictions. A common explanation to this apparent anomaly is that these economies are not labor-abundant in a global sense, but rather belong to an intermediate specialization cone.

Concerning the influence of trade liberalization on the personal distribution of income, the empirical results are mixed. Most of the countries in the world do not possess adequate inequality time-series and income data is associated with intrinsic measurement error problems. In addition to these important data limitations, some studies do not control for country endowments or income inequality determinants. Therefore, the various outcomes are not conclusive and there is scope for improved empirical tests.

Chapter 3

The Construction and Interpretation of Combined Cross-Section and Time-Series Income Inequality Datasets[1]

3.1 Overview

There is a sizeable literature regarding the interaction between income distribution and other economic variables, such as growth, poverty, trade and economic policy. Beginning with Kuznets (1955), the theoretical work has steadily grown, and recently there has been a surge in the topic, reflected in a new wave of publications (Atkinson, 1997). Complementing this recent literature, the introduction and easy availability of a cross-country inequality data set by the World Bank (Deininger and Squire, 1996) has launched a series of influential econometric studies.

While at the core of most recent work in the area, the inequality dataset compiled by Deininger and Squire, henceforth DS, has been recently criticized by Atkinson and Brandolini (2001). They forcefully argue for the need to assess the use of such "secondary"

[1] **This chapter is based on "The Construction and Interpretation of Combined Cross-Section and Time-Series Income Inequality Datasets" by Francois and Rojas-Romagosa (2004c).**

datasets mechanically and to deal more systematically with the measurement problems involved. This paper focuses on the empirical and theoretical difficulties related to income inequality measurement, analyzes the characteristics of "secondary" datasets and presents a new methodological approach to reduce the measurement error problems common to inequality information.

Substantial difficulties arise in the empirical measurement of inequality. The most notorious is the lack of an institution and agreed procedures that can assure data quality and consistency over time and across countries. In other words, an equivalent to the United Nations System of National Accounts, which provides macroeconomic statistics that are constructed by national agencies and are reasonably consistent over time and countries. In the absence of such an institution, some organizations have constructed "secondary" datasets, of which the best known are DS, the World Income Inequality Database (UNU/WIDER-UNDP, 2000) and the Luxembourg Income Study (LIS). These datasets compile available national inequality statistics and perform quality assessments of all the data observations. This has been an important first step towards the creation of internationally comparable inequality time series.

Deininger and Squire (1996) compiled a large amount of inequality observations for the entire world and classified each observation following three quality criteria. More recently, the World Income Inequality Database (WIID) has extended and updated the DS dataset, using similar quality criteria. Throughout the paper, we use this larger compilation of data provided by the WIID as our main inequality data source.

Beyond quality criteria issues, there are additional problems that increase the measurement error present in national series and in international inequality comparisons. In particular, national inequality statistics generally include observations that differ on concepts measured (i.e. expenditure, gross and net income), reference units (e.g. household, person, family) and/or sources. Subsequently we refer to these three distinctive characteristics as the inequality data definitions and we consider an inequality series to be consistent when these definitions are identical for all observations. Although some countries have relatively extensive and consistent time series, the general rule is that inequality observations are sparse and differ on definitions over time. Hence, to create relatively extensive inequality time series that can be used in econometric studies, it is often necessary to assume the comparability of some of the definitions.

Deininger and Squire assumed that all definitions are broadly comparable and used their "high quality" observations to construct the most consistent inequality time series for each country. However, they cautioned about the potential problems of the comparability assumption and as an alternative they advised the use of dummy variables to adjust and account for different definitions. Using this approach, they provide a single inequality series for a wide number of countries, which can conveniently be used in empirical research.

While convenient, we argue that these simplifying assumptions (i.e. the complete comparability of definitions and sources), introduce serious distortions and noise into the data. Furthermore, we agree with Atkinson and Brandolini and consider that the use of dummy variables is not an acceptable solution to this problem. Thus, a main contribution of this paper is the testing of the comparability of different concepts. We use the results of this comparability analysis to generate consistent inequality series, for which the measurement error is reduced.

Our goal is to diminish the measurement error implicit in currently available datasets (DS, WIID, LIS) and to provide an augmented version of these data that we hope is more reliable. However, we do not address all measurement issues. Most of the problems embedded in inequality data, including the comparability assumptions, have been extensively addressed by Atkinson and Brandolini for a sample of OECD countries[2] and this paper does not attempt to correct all of the limitations at hand.

We also explore conceptual issues of measurement. There are important theoretical considerations with regard to inequality measurement. While there are several indicators that measure inequality, there is no consensus in favor of any particular index.[3] Although the Gini coefficient is the most commonly used indicator, there are many more inequality indexes that can be used. Alternatively, distribution share analysis (i.e. quintiles and deciles information) can also be employed. The main difficulty involved is that an

[2]Throughout this paper we refer to OECD countries as: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, Luxembourg, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, United Kingdom and United States. Before 1990, Germany refers to the Federal Republic of Germany.

[3]A comprehensive survey of the topic can be found in Cowell (2000).

inequality index reduces the information provided by a whole distribution function to a single value. This simplification is helpful, but necessarily disregards information. There are several inequality indexes suggested in the literature, but an appropriate choice is to select an index that complements the information provided by the widely used Gini coefficient. In particular, since this coefficient is more sensitive to transfers in the middle of the income distribution and less sensitive to transfers at the extremes, it is convenient to include an index that can solve for this limitation. Such an index is provided by Atkinson (1970), which has an inequality aversion measurement parameter that controls for the sensitivity of transfers at the extremes.

However, there is an important practical difficulty involved. In order to estimate alternative indexes, we need the whole income distribution. Yet, most sources of inequality data only report Gini coefficients and grouped income shares. This means that we must first estimate the income distribution from grouped data and then estimate the inequality indexes. Specifically, we use parametric estimation of Lorenz curves to approximate the entire income distribution, and then use these estimates to calculate four different Atkinson indexes and poverty rates. These alternative inequality indexes complement the information provided by the Gini coefficient.

The paper is organized as follows. Section 2 explains the difficulties involved in dealing with inequality data. In the following section we assess the comparability assumptions and discuss the resulting assumptions needed in order to group different definitions and sources. In Section 4 we explain how we estimate the Lorenz curves and Atkinson indexes from grouped data, and in addition, we discuss how we have calculated poverty ratios. Once we have obtained the new inequality dataset, we compare it with the DS series and also compare the results provided by the Gini and the Atkinson indexes. Likewise, we analyze how international and intertemporal inequality has changed over time. This set of results and comparisons is presented in Section 5. We conclude in Section 6.

3.2 Problems when dealing with inequality data

We divide the tasks involved in building a cross-country inequality dataset into two main groups. The first group includes data compilation and quality control. These issues are relatively well addressed by existing datasets. The second group includes those issues that are not yet convincingly tackled: the intertemporal and international comparability and consistency of inequality data.

3.2.1 Secondary datasets

A "secondary" dataset is a summary of national information that is drawn from household income studies and micro-datasets produced by national surveys. The two most used datasets are the Deininger and Squire (DS) and the World Income Inequality Database (WIID). The WIID was constructed itself based on the DS dataset and has expanded this dataset and aggregated new available information. Thus, this is the largest and most exhaustive compilation of inequality data available. It provides up to 5067 data observations, for different definitions, coverage and quality ratings. Therefore, we take it as our starting point and main source of information.

The secondary datasets provide two important advantages. They compile most of the available inequality data into one source, and they check for the quality of each observation. The quality controls used to filter information from the primary to the secondary datasets eliminate unreliable data and inequality observations that are not representative of the whole country. Deininger and Squire (1996) used three quality controls:

> "The statistics were selected by requiring that they be from national household surveys for expenditure or income, that they be representative of the national population, and that all sources of income or expenditure be accounted for."

The WIID quality ratings are very similar to those of DS. However, there are some important differences.[4] In particular, WIID considers as reliable data some of the observations that did not have a clear reference to the primary source, while DS did not

[4] These quality criteria differences introduce some divergences between our dataset and DS, which are not accounted for by the comparability assumptions we use later.

consider these observations in their "high quality" dataset.[5] The second important difference is the inclusion of observations based on monetary income, which is not used in DS because it does not account for all sources of income.[6] Finally, missing income concepts are not accepted by WIID and this implies that we do not consider some of the DS observations. The reliable data ratings of the WIID are labeled as OKIN and from a total world population of 150 countries, we have OKIN data for 141 countries.

The main difference between DS and WIID, however, is that the last does not identify a single time series for each country. Instead, the researcher has the full available information and has a wide range of series to choose from. The disadvantage is that there is no clear indication on how to use or join inequality observations with different definitions and/or sources.

Finally, it is important to mention that another source of inequality data, which we do not use, is from the micro data-sets provided by the Luxembourg Inequality Study (LIS) and the Living Standards Measurement Study of the World Bank. We do not consider their use here because their coverage, in terms of time and countries, is very limited. In addition, they are usually difficult to access and obtaining summary statistics is very time consuming in an already burdensome process.[7]

3.2.2 Definition inconsistency

We follow the WIID and classify each data observation into six characteristics: concept measured, reference units, area coverage, population coverage, data sources and quality ratings. Technically there are other distinctive characteristics that may significantly alter the inequality values, such as: survey methods, sample characteristics, income issues included (imputed rents for own-occupied houses, insurance premia, interests and dividends) and the time period considered.[8] Nevertheless, since our main interest is to obtain

[5] As explained later, the inclusion of these observations significantly increases the number of inequality data points in the 1960s and 1970s. Although the measurement accuracy may be reduced, it provides a valuable extension in the time series (Barro, 2000).

[6] We justify the inclusion of these observations in the following section. Mainly, the data included are from rich or middle-income countries, for which non-monetary income is not expected to be significant.

[7] Nonetheless, these data sets may be a preferable source of information for single country inequality analysis or limited cross-country analysis.

[8] Some of these issues can induce substantial measurement errors of their own. For example, imputed rents can represent a significant share of household income in some countries and if it is included in the

a reliable cross-country dataset, we do not deal with these measurement issues, which are generally country-specific, and we focus instead on the broad characteristics of the inequality observations. Thus, our resulting dataset diminishes the measurement error embedded in inequality data, but does not entirely eliminate this problem.

From the main six characteristics, we follow DS and select only data that considers the entire population and has a national coverage. Moreover, we use the quality criteria provided by the WIID and select only the observations labeled as OKIN. After this first filtering of the information, we remain with three characteristics: concept measured, reference unit and source. Since there are multiple combinations of concepts and recipients, and usually more than one source per country, we have what Atkinson and Brandolini (2001) refer to as a "bewildering variety of estimates". That is, a number of generally discontinuous series, with differences in one or two definitions and usually from different sources. The problem can be better understood by looking at Table 3-1.

For this particular country we have seven different series, two income concepts (gross income and gross monetary income), two recipients or reference units (person and household) and six sources. However, the number of series and definitions involved can be larger in other countries. In total, there are five different concepts and as many as nine different income recipients.[9] Additionally each data point provides the Gini coefficient and sometimes distribution shares in quintiles or deciles of population. In Table 3-1 the observations with distribution shares are indicated by the data in boxes. Moreover, the Gini coefficient can be given by the primary source or directly estimated from the distribution shares when available.

These characteristics of the inequality data leave several questions to be answered. Can we mix different definitions of income concepts and income recipients? Can we mix different sources? If yes, how do we mix them? Which data observation should we choose when there is more than one available for a given year? How many series should we analyze per country? In order to have an inequality dataset that can be readily used for

household survey it can create an important source of distortion in the comparability of different inequality observations. Ravallion and Chen (1997), and Atkinson and Bourguignon (2000) discuss further on these points.

[9] The concepts are gross income, gross monetary income, net income, net monetary income and expenditure. The most common reference units are person, household, household per capita, household equivalent, family and family equivalent.

Table 3-1: Chile, Gini series with different definitions, 1968-1996.

Income concept:	Income Gross						Monetary Inc. Gross	
Recipient:	Household			Person			Household	
Source:	UN	Fields	SH	Mideplan	Paukert	IADB	SH	DS accept
Series:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
1968	$\boxed{45.64}$				$\boxed{44.00}$			45.64
1969								
1970								
1971		$\boxed{46.00}$						46.00
1972								
1973								
1974								
1975								
1976								
1977								
1978								
1979								
1980				53.21				53.21
1981				53.46				
1982				56.98				
1983				54.49				
1984				55.85				
1985				54.91				
1986				55.69				
1987				56.72				
1988				54.50				
1989				$\boxed{57.88}$				57.88
1990			54.70	53.18			55.65	
1991				55.38				
1992			$\boxed{52.19}$	$\boxed{50.70}$		52.00	53.08	
1993						50.00		
1994			55.58				57.42	56.49
1995								
1996			56.37				57.24	

Note: Observations in boxes represent data with distribution share information
Source: WIID, version 1, Sept. 2000

empirical research we must answer these questions. The questions are relevant if we want to analyze country time series, as well as cross-country analysis. Given that especially in developing countries there are not many inequality observations, there also has to be some assumptions regarding the combination of the different definitions in order to obtain at least one series per country.

3.2.3 Deininger and Squire approach

DS assume that all definitions are broadly comparable and instead, focus on the quality of the observations. Thus, they freely mix the different definitions, regardless of income concept or reference unit. However, they acknowledge the potential measurement errors that this approach may cause and recommend the use of dummy variables to deal with the problem.

This strategy allows them to present a single time series for each country, which is very convenient for empirical studies. Moreover, when there is more than one observation per year that satisfies their three quality criteria, they choose the observation which is consistent with the rest of the series. In other words, they try to maintain the same income concept, recipient unit and source when possible. Their final inclusion criterion is that the observation originates from an official publication. In the last column of Table 3-1, we present their "high-quality" data observations, which are labeled as DS-accept. In this particular case, they mix household and person reference units and three different sources. The 1994 observation they use is not included in our series because it does not properly define the type of income it uses. This lack of information accounts for the WIID considering the observation not reliable (NOOK) and we do not use it.[10]

Nevertheless, this grouping procedure has been strongly criticized by AB (Atkinson and Brandolini, 2001). Using a sample of OECD countries, they show how inconsistent it is to create such series. In many cases the constructed DS-accept series significantly modifies the level and even the trend of some inequality series, in comparison with series that use only consistent income concepts. Furthermore, AB demonstrate for OECD countries, that the use of dummy variables is not enough to render some definitions comparable. In particular this is the case for net and gross income, as well as for income based and expenditure based observations.[11]

DS defend the grouping of net and gross income by assuming that in developing countries, where there is not enough data to compare both definitions, redistribution is not important and thus, gross and net income are comparable. Yet, AB stress the inconsistencies this mixing yields for many OECD countries for which both types of income are available. Compared with the Luxembourg Inequality Study (LIS), which adjusts inequality data to make it international comparable within the OECD countries, AB find that the rankings provided by DS are very different from those of the LIS.

Similarly, for the case of expenditure-based and income-based observations, DS acknowledge that both concepts are significantly different. To correct for this problem, they suggest the use of a fixed adjustment to render both concepts comparable. In their dataset they find that expenditure-based observations are on average 6.6 points below

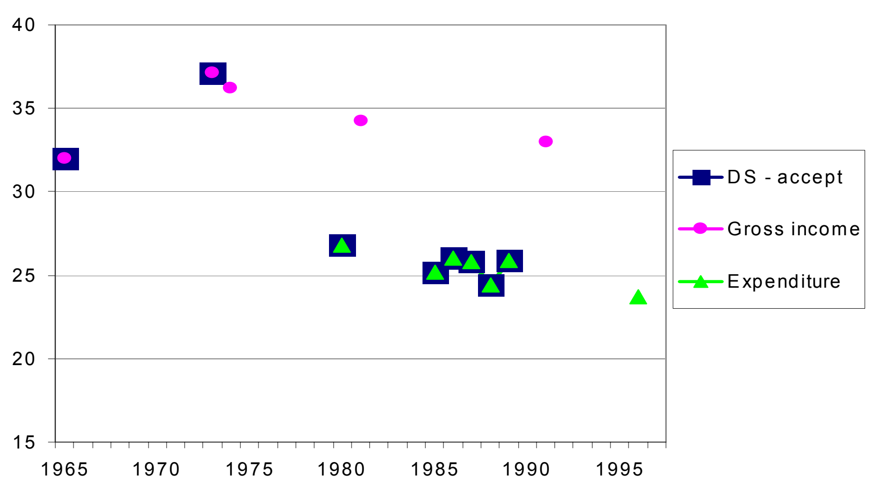
[10]For the rest of the MIDEPLAN observations, they are not DS-accept because there was not a clear reference to the primary source. However, WIID consider this data to be reliable and it provides more observations to be used.

[11]In the next section, we use a bigger sample and reach the same conclusions.

income-based Gini coefficients. However, this particular value is conditional on the sample they analyze. We use the same procedure and compare the inequality levels of both concepts only for those countries where both are available, but in our sample we find that the average difference is three points.[12] Therefore, using the fixed value proposed by DS increases in average, around 3.6 points the levels of inequality for the countries in our sample. We conclude that the estimate of the true difference can be unreliable and the use of fixed adjustments introduces arbitrary noise in an already problematic dataset. In short, we agree with AB and conclude that the use of fixed adjustments is not enough to reconcile both definitions.

The treatment of different reference units by DS is also problematic. DS collapse the numerous definitions into two categories: person-based and household-based observations After this rough grouping, they compare both definitions and conclude that they are not significantly different. However, not all the person-based definitions are comparable, nor are all the household-based ones. This also raises questions about their comparability assumptions.

Figure 3-1: Spain, Gini coefficient series



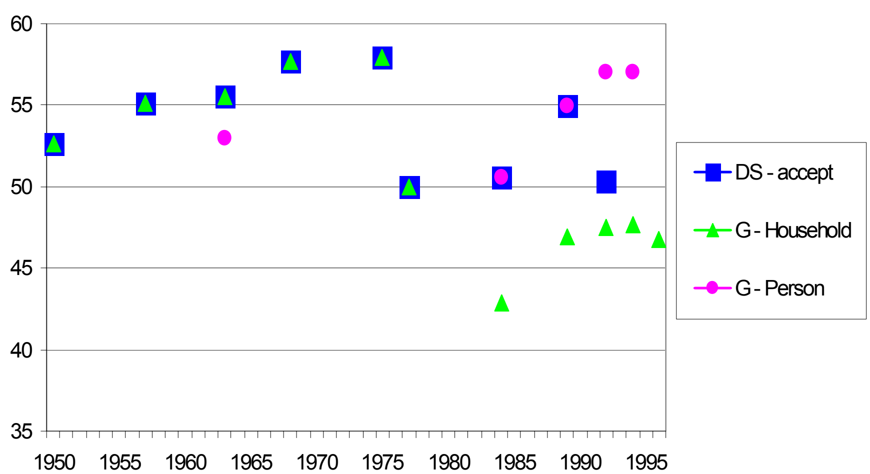
We use two examples to highlight in detail the main problems involved when grouping heterogeneous series. In Figure 3-1 we plot two inequality series for Spain, each differs in

[12]The sample difference is given by the larger compilation provided by the WIID and by the differences in some of the quality criteria we explained before. In our case, we can directly compare 19 countries that have both income and expenditure information, of which 58% belong to the OECD.

the concept measured (e.g. gross income and expenditure). Each series has different levels and no significant trend. However, DS use two gross income observations to expand the expenditure series and this alters significantly the inequality conclusions. First, the DS has a time trend, with a considerable decrease in inequality from the 1970s to the 1980s. Secondly, the combination of both concepts significantly changes the inequality levels. In particular, using the gross income series for the 1980s yields inequality observations of around 8 points higher than the DS-accept values. Such a dramatic level variation substantially changes the country's international inequality ranking, as we show in Section 5.

In Figure 3-2 we plot two series for Mexico that differ in their reference unit (household and person), but have the same gross income concept. Both series have distinctive time trends and significantly different levels in the last years. DS combine both reference units and again, freely mixing different definitions changes the inequality results. In particular, for the 1980s, the levels of both series are quite different –around 8 points– and this introduces an important modification in the inequality results. Furthermore, DS report an increase in inequality from 1977 to 1989, while the consistent gross-household series shows the opposite result. Finally, the DS dataset fails to indicate any time trend at all for the entire period, while the gross-household series has, at least, a decreasing trend.

Figure 3-2: Mexico, Gini coefficient series



Although such examples are not widespread, they certainly introduce noise in the data that increases the measurement error and may affect the overall empirical results. While the corrections proposed by DS may sometimes work, on other occasions they may distort the data further or leave the discrepancies unaltered. In our last example from Mexico, DS do not recommend any correction for reference unit differences and the problem shown above persists. For the case of Spain, the use of the particular 6.6 adjustment value reduces the inconsistencies, but in other cases it does not help.

In summary, although the DS dataset was a very important step forward in the study of inequality data, it has significant limitations that increase the measurement error and may seriously alter the empirical results of those studies that use this inequality dataset. In particular, though it is quite reasonable to use the kind of quality control considerations they introduced, the evidence suggests that grouping different definitions to create single country time series is unreliable and the use of fixed adjustments does not correct this problem.

3.3 Controlling for definition comparability and series grouping

Since there are several concepts and reference units, we still need to have a methodology to collapse the various series available for each country. This is necessary to obtain consistent and comparable series that can be used in cross-country and time series studies. To collapse further the existing series provided by the WIID, we take three steps. First, we group those series which have the same definitions (concept and reference unit). The second step is to make a comparability analysis and judge which definitions and conflicting sources can be mixed and how. The final step is to construct the national series, using consistent groupings and standard procedures based on these results.

3.3.1 Grouping series with the same definitions

It is straightforward to group those series with identical definitions. The main difficulty in this first step is to deal with different sources. In some cases, we can have a year where two sources report observations with the same definitions. This is the case for Chile in 1992, shown in series 4 and 6 of Table 3-1. If both sources have series that can be analyzed[13] we run the same comparability tests as below. If this is not possible, we choose observations following this preference ordering:

- Observations with income share information
- LIS data
- DS accept data
- The source with the longest time coverage.

Since one of our main purposes is to compare the Gini coefficient with alternative indexes, we need the income share information to construct such indexes. The Luxembourg Income Study (LIS) is a project that has created a micro-data of social and economic information. It has been used to explicitly compare cross-country inequality information and thus presents adjusted data for such purposes (Atkinson, *et al.*, 1995). Finally, the last two ordering preferences ensure consistency in the series and the previous considerations contemplated by DS.

In our example for Chile, we have two conflicting sources: MIDEPLAN and IADB. Using the four decision criteria stated before we prefer the MIDEPLAN observation, since it provides income share information. After doing this first grouping, we have collapsed seven series into three (see Table 3-2).

We use the same standard procedure with all the countries to collapse series with the same definitions. However, this first step is insufficient given that many countries remain with several series (i.e. Sweden has up to 14) and further grouping procedures are necessary.

3.3.2 Comparability analysis

Li *et al.* (1998) have compared values of the Gini coefficient with different definitions for countries and years where estimates are available. However, this procedure is biased

[13] As explained below, this requires that both series have a common sample of at least three observations in a time span of five or more years.

Table 3-2: Chile, grouped Gini series, 1968-1996.

Income concept:	Income Gross		Monetray Inc. Gross	
Recipient:	Household	Person	Household	
Source:	3s	3s	SH	DS accept
Series	(1)	(2)	(3)	(4)
1968	45.64	44.00		45.64
1969				
1970				
1971	46.00			46.00
1972				
1973				
1974				
1975				
1976				
1977				
1978				
1979				
1980		53.21		53.21
1981		53.46		
1982		56.98		
1983		54.49		
1984		55.85		
1985		54.91		
1986		55.69		
1987		56.72		
1988		54.50		
1989		57.88		57.88
1990	54.70	53.18	55.65	
1991		55.38		
1992	52.19	50.70	53.08	
1993		50.00		
1994	55.58		57.42	56.49
1995				
1996	56.37		57.24	

Note: "3s" refers to three different primary sources
Source: WIID, version 1, Sept. 2000

towards the sample of countries with available data. A more satisfactory procedure is to compare observations available for the same country and the same year, as was done by DS.

The existing literature does not offer a consistent and standard comparison procedure. DS limit themselves to comparing the average difference between different definitions. In the case of income and expenditure, they try to check the correlation of the differences with some explanatory variables. In our view, this procedure is not satisfactory. For two series to be comparable, and thus, interchangeable if one data observation is missing, we need much more than average differences. We need two series which have a very similar trend and level. In the case the level is not comparable, we need that the difference between both series is relatively constant over time and only in such cases does it makes sense to freely mix different series. In other cases, grouping definitions that are not comparable can seriously alter the level and/or trend of the series. Adjusted data in this way at best increases the measurement error and at worst can invalidate the empirical results.

In addition, to clearly single out what we are comparing, the series should differ in only one of the definitions. For example, if a Gini coefficient series for net income and household seems comparable with a series of gross income and person; it could be because indeed gross and net are comparable or because net and gross are different but a combination of income concepts and reference units produce the similarities.

3.3.2.1 Comparability procedure

To analyze different definitions and sources, we use the following procedure:

- We use only Gini data for the same country and the same year when they differ in only one of the definitions (concept or reference unit) and have a common sample of at least three observations in a time span of at least five years.

- We estimate the simple correlation between both series. If the correlation is negative we conclude that the series are not comparable.

- We check if both series are normally distributed and run hypothesis tests for equal mean and equal variance (i.e. a t-test and an F-test). If the variance is significantly different (at a 5% significance level) then we conclude that the series are not comparable. If the mean is significantly different, we test if there is a constant difference between them.[14]

- When the series are positively correlated and have the same variance, they move in the same direction over time. If instead, the mean is not equal we use the average difference between series.[15] Furthermore, we check for one, three and five points differences in the means (which are some values reported as the average difference between series with different concepts).

- To complement the hypothesis of equal mean and variance, we take OLS regressions on the equation: $S_1 = \beta\left(S_2 + c\right) + \varepsilon$, where S are the series and ε the error term.

[14] For the few cases where the series are not normally distributed, we use the ANOVA F-statistic to test for equality of means and the Levene and the Brown-Forsythe methods to test for the equality of the variances. Again, if the variance is significantly different we conclude that the series are not comparable.

[15] We round the values to the closest integer to simplify the procedure. In some cases we need to use half points in order for the series to be comparable.

We run a Wald coefficient test to check the null hypothesis that $\beta = 1$. When the means of S_1 and S_2 are not the same c is the average difference of the series, otherwise $c = 0$. To check how sensitive the series are to absolute differences in the mean, we also test the null hypothesis when c is ± 1, 3 and 5. Note that in this case the inclusion of c is equivalent to the use of a fixed adjustment or the use of a dummy variable.

- In summary, we consider two series to be comparable when they have a positive correlation, not a significantly different variance and we cannot reject the null hypothesis that $\beta = 1$, when $-1 \leq c \leq 1$. For other values of c we consider the series to be comparable, but with a constant absolute difference between them. In this last case, we must add c to make the series compatible.

Through this comparison procedure we attempt to assure that both series have statistically the same time trend and the same level (or an absolute constant difference). This ensures that when freely mixing two series we do not alter the trend or level of the resulting series. This procedure was also employed to compare some series that differed only in the source.

3.3.2.2 Comparability assumptions

Once we have pared comparable series and followed the above procedure, we can study which definitions can be mixed. The results of this analysis provide the basis for establishing the comparability assumptions we use later. This allows us to consolidate series and reduce the number of definition combinations available for each country. In total we have 179 possible comparable pairs for 38 countries, 14 of which are OECD countries with 107 pairs and 24 are developing countries with 72 pairs. The results for all the comparable definitions are summarized in Table 3-3.

In this table, the first column shows which variables are being compared. The next two columns indicate the number of pairs compared and the percentage that belongs to OECD countries. The next column is very important, since it shows the percentage of series that are not comparable, either because there is a negative correlation between the series or because they fail the equal variance test. The columns labeled $c\pm 1$ and $c\pm 2$ show

Table 3-3: Comparability results for all definitions

	Sample		Comparability Results					DECISION
	Obs	OECD	Not Comparable	c ±1	c ± 2	average c	percentage comparable	
Income vs Monetary Income	7	43%	0%	43%	71%	-2	29%	**uncertain**
Income vs Expenditure	19	58%	42%	16%	32%	3	26%	**no**
Gross vs Net	36	78%	31%	19%	28%	3	39%	**no**
Person vs HH per capita	8	0%	0%	88%	88%	0	88%	**yes**
Household vs HH equivalent	36	100%	25%	19%	28%	4	36%	**no**
Household vs Person/HHpc	23	26%	9%	61%	61%	0	61%	**uncertain**
Household vs Family	8	75%	0%	50%	63%	2	38%	**uncertain**
Household vs Family eq.	18	100%	6%	28%	44%	5	17%	**no**
Household eq. vs Family eq.	16	100%	13%	81%	88%	0	81%	**yes**
Family vs Person	2	50%	0%	100%	100%	1	100%	**uncertain**
HHe/Fe vs Person/HHpc	6	83%	17%	67%	83%	0	67%	**uncertain**

Notes: The OECD column corresponds to the percentage of observations from these countries.
The Not Comparable column presents the percentage of observations with a negative correlation and/or different variance.
The c±1 column shows the percentage of comparable pairs with a fixed adjustment (c) of +1 or -1.
The next column is the equivalent when the ±2 range is used.
The Average c column reports the mean difference between definitions for all observations.
The following column shows the percentage of comparable pairs when this average c is used.
The last column reports the comparability result for each pair of definitions.

the percentage of comparable series when the absolute average difference is less than one and two, respectively. The following column shows the sample average difference and the following column reports the percentage of series that are comparable when this average difference is applied. The final column indicates the decision regarding the comparison of definitions.

In Table 3-3 we show all the 11 possible comparisons. We analyze the six most relevant pairs and we use them to illustrate how we reach the final decision regarding the comparability assumptions. The remaining five couples follow the same procedure and we just mention the final decision.

Three main considerations were taken into account when deciding which series could be comparable. The first criterion was the percentage of non-comparable couples. A high percentage indicates that the considered series had different trends and hence, provide a very bad substitute for a missing series. The second criterion is the percentage of comparable series when no fixed adjustment is applied and when the average difference is applied. These percentages show how good it can be to mix series with or without a fixed adjustment. Finally, we prefer combinations with large samples and, for some cases, the percentage of OECD observations is also relevant for the analysis.

Income and Monetary Income. In this case, we have four developing countries and the USA, which offer seven comparable pairs. Of these 43% are comparable series with an average difference (c) smaller than ± 1 and 71% are in the range of $c < |2|$. Almost one third of the pairs are comparable when using the average difference of -2. Moreover, all the pairs are positively correlated, have equal variance and an average difference of -2.

This result is theoretically consistent, since monetary income excludes own-produced consumption and it should report a higher level of inequality. Since all the series are in principle comparable it seems reasonable to freely mix both definitions. Nonetheless, we are uncertain of which fix adjustment (average difference) to apply. If we use the average difference of -2 only 29% of the series are comparable. Another inconvenience is the small sample of only seven observations. Thus, we are uncertain about this comparison couple.

It is important to note that DS did not take into consideration data observations defined for monetary income. They argue that the consumption of own produced goods is an important source of revenue for poorer households and not taking into account this kind of consumption can skew the indicator towards more inequality. However, our results suggest that for some cases, a fix adjustment can render both definitions comparable. In particular, we mix both definitions only in rich and middle-income countries[16], for which one does not expect this kind of consumption to be important. Therefore, we find it reasonable to include data based on monetary income for these countries and by doing so we can expand the available number of observations.

On the other hand, the example provided by DS to exclude monetary income is not compelling. Although later they conclude, as we do, that income and expenditure-based are not readily comparable, in their paper they actually compare the monetary income

[16] Australia, Brazil, Costa Rica, Hong Kong, Panama, Russia, United States and Venezuela.

and expenditure observations for Greece in 1974, as a way to associate income and monetary income. Since for this year the difference in both series is of roughly six points they conclude that both definitions cannot be mixed. However, when applying our comparability analysis to Greece, we conclude that both series are indeed comparable when adding three points to the expenditure series.

Income and Expenditure. In this case 42% of the observation couples are not comparable at all, i.e., they have a negative correlation or significantly different variance. Moreover, if we were to use a fixed adjustment to compare series, only 26% are comparable for the average difference of 3 points. If we were to use the average difference found in DS of 6.6, then only 11% of the series are comparable. In other words, in almost one out of ten cases the fixed adjustment significantly alters the level and/or trend of the series, overwriting information contained in the original data.

Since expenditure does not take into account income that is saved, we expect it to give less unequal values. On the other hand, expenditure information can take into account income smoothing by borrowing or lending. Thus, we also expect lower inequality values from expenditure surveys. However, in developing countries the main limitation is that many households do not know their actual income or their knowledge is incomplete. This limitation can be explained by the presence of significant own-produced consumption, temporal and/or irregular monetary income sources. Therefore, it is easier to survey their consumption (expenditure) and this has become a common practice in poor countries.

Nevertheless, given the very high percentage of non-comparable cases, the most likely possibility is that both series are providing different inequality information. Therefore, we conclude that it is not reasonable to mix income and expenditure definitions, not even when using fixed adjustments.

Gross and Net Income. For this case, almost a third of all the series couples are not comparable. The average difference is of three points, but only 39% of the series are comparable when such a fixed effect is applied. Although this is not such a clear case as the previous, we also reject comparing gross and net income.

One argument used in DS to compare both definitions is that in developing countries the difference should not be big, assuming that their redistribution systems have a small impact on incomes. However, of the eight comparisons that come from non-OECD countries, 25% are not comparable, 50% are comparable with $c < |1|$ and 38% are comparable when the average difference of three points is applied. Therefore, even when the sample is still small and all the non-OECD countries involved are middle-income countries, there is not much evidence that gross and net income are equivalent.

On the other hand, grouping series with net and gross income does make a significant impact in OECD series. In particular, AB show in detail how damaging the combination of both definitions is to the information on levels and trends contained in the original series for these rich countries.

Mixing different reference units. Grouping person and household per capita is probably the clearest case in favor of mixing definitions. In 88% of the cases we can freely mix both series. In addition, we do not need to adjust for any fixed effects. Although the sample of eight is small, the evidence is strong. Therefore, we assume that both series are comparable and we evaluate this grouped definition (person-household per capita) with other income recipient definitions.

On the other extreme, comparing household and household equivalent does not seem reasonable. A quarter of the observations are not comparable and when the average difference of four points is applied only 36% of the series are comparable.

The comparability of household with person/household per capita is an uncertain one. The non-comparable percentage is relatively small, but only in 61% of the cases are the definitions comparable. In fact, household and person are the most common reference units and this particular comparability assumption is very important and we deal with it latter.

Finally the remaining five comparability results are the following: household and family (uncertain), household and family equivalent (no), household equivalent and family equivalent (yes), family and person (uncertain), household equivalent/family equivalent and person/household per capita (uncertain).

In summary, for the eight couple of reference unit definitions that were tested, only one was clearly non-comparable. These results suggest that one may mix reference units in many cases, but not in all.

3.3.3 Grouping the data

Using the results of the previous section we can directly group those series with comparable definitions and reduce the number of series per country. However, since the three main concepts (e.g. net income, gross income and expenditure) cannot be mixed, then inevitably we have more than one inequality series for some countries. Moreover, we associate each concept with the most common reference unit and this union creates our three main resulting series: Gross Income-Household, Net Income-Household and Expenditure-Person.

The presence of three series may seem inconvenient when conducting empirical research, but different income concepts may offer different information about inequality behavior and using more than one concept can increase the available information that we can use. For example, evaluating both gross and net income inequality measures provides important information concerning the redistribution policies of some national governments. In addition, trade theory makes direct predictions about gross, not net factor incomes.

On the other hand, separating series because of different reference units is not very compelling. Any difference provided by dissimilar reference units are mostly explained by demographic factors. When the size of the household changes according to the income classes, one can expect different inequality results from household and individual information. In addition, the number of adults in different income classes can provide divergences in the inequality results. However, these demographic factors vary across countries.

An additional advantage of our comparability analysis is that it offers country-specific information. For those countries were we conducted the tests, we have and indication of whether specific national series with different definitions may be comparable or not. This information is used to group series in that particular country, even if the overall analysis resulted in a verdict of non-comparable definitions. For instance, in a country were gross and net income are comparable with a fixed adjustment, we can use this information to expand one or both series.

Finally, we have to take a decision concerning those definitions that yield an uncertain result. The approach we take is to have two broad types of series. First, we construct "basic" series in which we are confident of the definition groupings used. Hence, we only mix definitions for which we have strong evidence that they may be comparable and/or definitions that are comparable for that particular country.

We then construct "extended" series. These series have more observations, but we use definition groupings that are less reliable and we combine definitions for which we are somewhat uncertain about their comparability. For example, the most common "basic" series is the data based on gross income and household units, and this series is mostly complemented with observations based on monetary income and person units to create the "extended" series.

In summary, to finally collapse the remaining series, we perform the following standard procedure:

- We group those reference units for which we are certain about their comparability
- We use the country specific information to further group series. In particular, to expand the most common series: Gross Income-Household, Net Income-Household and Expenditure-Person. This includes grouping series with a fixed adjustment, when the evidence supports this type of comparability assumption.[17]
- We also take advantage of the LIS data (only available for OECD countries) to adjust the series in those cases in which they are comparable. E.g., if we have a net income-household series from both the LIS and another source, and both series are comparable with a fixed adjustment, then we adjust the series to have the levels reported by the LIS data.

[17] When some observations are adjusted to make them consistent with the rest of the series, we have a problem with the Atkinson indexes (that we calculate later). Generally, the difference between Gini coefficients with different definitions is not the same for the Atkinson indexes. To solve this problem we use two standard approaches. A direct approach is applied when we have three or more comparable Atkinson observations and then, we directly apply the average difference. The indirect approach is to estimate the average difference between Gini and the particular Atkinson index for both series and then adjust the Atkinson by the difference between both averages.

- Finally, we use definitions for which we are uncertain about their comparability to create the "extended" series.

In brief, for each country we can have one or more basic series relating to gross income, net income or expenditure and in some cases, extended series, which expand the series by including uncertain definition comparisons.

Even though we apply this same procedure for each country and in every case we try to be as objective as possible, we face the same dilemma expressed by DS and "decisions concerning the inclusion or exclusion of certain observations are always based on some judgment and arbitrariness".

In our case, our results can be replicated using the WIID database. This allows other researches to review our procedure and make their own changes if necessary.

Table 3-4: Chile, final Gini series, 1968-1996

Income concept:	Income Gross		
Recipient:	Household/Person	Person	
Source:	3s	3s	DS accept
Series	(1)	(2)	(3)
1968	45.64	44.00	45.64
1969			
1970			
1971	46.00		46.00
1972			
1973			
1974			
1975			
1976			
1977			
1978			
1979			
1980	54.71	53.21	53.21
1981	54.96	53.46	
1982	58.48	56.98	
1983	55.99	54.49	
1984	57.35	55.85	
1985	56.41	54.91	
1986	57.19	55.69	
1987	58.22	56.72	
1988	56.00	54.50	
1989	59.38	57.88	57.88
1990	54.70	53.18	
1991	56.88	55.38	
1992	52.20	50.70	
1993	51.50	50.00	
1994	55.58		56.49
1995			
1996	56.37		

Source: WIID, version 1, Sept. 2000

In our previous example for Chile, from an initial number of seven series, we can collapse the inequality information into two "basic" series (see Table 3-4). Here we used the fact that for Chile the household and person series are comparable with an adjustment of 1.5 points. This allows us to expand the Gross income-Household series from 6 to 15 observations.[18]

When compared with the DS accept series the levels are not significantly different, but in this example, we find different time trends. In section 4 we compare in detail the inequality trends reflected in our dataset against those present in the DS dataset.

3.3.3.1 Characteristics of the three main series

When our grouping methodology is applied to all the countries, we still have several series, especially in OECD countries. However, the three main series have comprehensive world coverage and can be readily used for empirical analysis. The main statistics of these series are presented in Table 3-5.

Table 3-5: Characteristics of the three main inequality series

	Gross Income Household		Net Income Household		Expenditure Person		Total for the three series		DS-accept
	Gini	Share data	Gini	Share data	Gini	Share data	Gini	Share data	
BASIC series									
Countries	49	38	27	25	69	63	145	126	115
Observations	427	326	288	241	189	159	904	726	693
Average obs.	8.71	8.58	10.67	9.64	2.74	2.52	6.23	5.76	6.03
OECD countries	29%	39%	70%	72%	1%	2%	23%	27%	17%
EXTENDED series									
Countries	95	70	47	43	85	75	227	188	115
Observations	634	445	433	376	254	205	1321	1026	693
Average obs.	6.67	6.36	9.21	8.74	2.99	2.73	5.82	5.46	6.03
OECD countries	17%	23%	43%	44%	1%	1%	16%	19%	17%

Although all the series have a smaller sample and coverage than the DS-accept series, we still have a satisfactory representation. Moreover, the extended gross-household series is fairly comparable in number of observations and OECD representation to the DS-accept series.

The net-household series seems to be better suited to analyze OECD countries, while the expenditure-person series consists of a majority of developing countries. The gross

[18] Note that for international comparisons we only use the first series. The second series is only used when analyzing inequality in Chile.

income series is better balanced between rich and poor countries. These sample differences between concepts can also be observed in the average series length. OECD countries have longer series and this is reflected in the average number of observations in the net income series. In contrast, developing countries and the expenditure-person series have shorter series, and in many cases just one or two observations.

Additionally, we have at least one of the three main series for 141 countries. This is an improvement with respect to the 115 country coverage of the DS dataset. The share data is also sparser than the Gini coefficient observations and again, only the net income series has almost the same number of share data and Gini observations, due to the better quality of the OECD inequality data.

3.3.3.2 Series length and panel data analysis

Another important decision regards the length of the series we choose to analyze. Given the sparse amount of inequality information in many countries, once we have created the basic and extended series, some countries end up with only one or two data points per series.

Subsequently, when we analyze inequality trends and when we compare our dataset with DS, we use only countries with at least one series with three or more observations, in a time span of at least five years. This allows us to study cross-country inequality trends and use panel data analysis. This set of countries with long series includes 80 countries and the main characteristics are reported in Table 3-6. In the Appendix we present the summary statistics of this subset of countries with longer series

Our full dataset, which includes countries with only one or two observations, can be used to conduct cross-country analysis for specific years of our three main series. It also allows for decade or five year averages that can be used as a panel database. Since there is a majority of expenditure-person series for developing countries in the full dataset, this can give a better representation than the longer series.

In summary, we have six main series: three basic series, with consistent definition comparability and three extended series with less reliable comparability assumptions.

Table 3-6: Characteristics of the three main inequality series, countries with three or more observations

	Gross Income Household		Net Income Household		Expenditure Person		Total for the three series		DS-accept
	Gini	Share data	Gini	Share data	Gini	Share data	Gini	Share data	
BASIC series									
Countries	38	34	20	20	22	21	80	75	66
Observations	413	322	279	235	127	106	819	663	634
Average obs.	10.87	9.47	13.95	11.75	5.77	5.05	10.24	8.84	9.61
OECD countries	37%	47%	75%	75%	5%	5%	38%	43%	29%
EXTENDED series									
Countries	57	47	32	31	30	28	119	106	66
Observations	580	420	407	354	182	146	1169	920	634
Average obs.	10.18	8.94	12.72	11.42	6.07	5.21	9.82	8.68	9.61
OECD countries	26%	32%	50%	52%	13%	11%	29%	32%	29%

In constructing these series, we have used the recommendations made by Atkinson and Brandolini (2001):

> "The use of simple dummy variable adjustments for data differences is not appropriate. Over time, the net and gross income distributions may behave differently, as may the distributions for households and for families. It is necessary to piece together information from different sources, informed by an awareness of their relative strengths and weaknesses. All of this points to the need for a blend of quantitative and qualitative analysis, and the avoidance of mechanical use of the (secondary) datasets."

3.4 Alternative inequality indexes

Given our inequality series, we now turn our attention to a completely different topic. In this section we estimate Lorenz curves from distribution share or grouped data. This allows us to construct the Atkinson indexes and estimate headcount poverty ratios.

Thereafter, we can use these estimates to conduct some tests. First, we can compare our Gini estimates, which are drawn from the Lorenz curve, with the Gini coefficients reported by the primary sources. Secondly, we can test if the inequality information provided by the Gini coefficient is similar to the one offered by the Atkinson indexes. The later can be considered a test of the robustness of the Gini coefficient as an inequality indicator.

3.4.1 Inequality measurement

Most of the inequality observations provided by the secondary datasets are given by Gini coefficients. Formally, this inequality measure is given by:[19]

$$G = 1 + \frac{1}{n} - \frac{2}{n^2\overline{y}}\left(y^1 + 2y^2 + ... + ny^n\right) = 1 + \frac{1}{n} - \left[\left(\frac{2}{n^2}\sum_h \frac{hy^h}{\overline{y}}\right)\right]$$

where income is arranged so $y^1 \geq y^2 \geq ... \geq y^n$. This inequality index is widely accepted and used. However, there are many other inequality indexes and there is no theoretical prerogative to prefer any.

Inequality is associated with the variance of the income distribution and this creates two basic measurement complications. First, as with any distribution, it is not a single-valued variable. Second, even when the concepts of Lorenz-dominance and Generalized Lorenz-dominance (Shorrocks, 1983) are widely accepted as ways to impartially rank two different distributions, in many cases the Lorenz-curve intersects at least once, and this method yields an incomplete ranking of distributions.

To solve both problems, inequality indexes are used to rank distributions in these indeterminate cases and to provide a single-valued variable that can be used in empirical models. However, since all inequality indexes have a specific method to weight and rank incomes from different levels, there is no objective inequality index and any inequality indicator has built-in social preferences. Moreover, many inequality measures are implicitly based on a social welfare function.[20]

In particular, when the Lorenz curves intersect, different indexes can provide different inequality information and this makes the choice of the index important for the results. For instance, the Gini coefficient is more sensitive to changes in the middle of the income distribution and it is less sensitive to movements at the extremes. On contrary, the family

[19] A summary with definition of the variables is presented in the Appendix to Chapter 4.
[20] Dalton (1920), Kolm (1969) and Atkinson (1970).

of Atkinson indexes is precisely more sensitive to changes at the extremes and thus, this index is a very convenient complement to the Gini coefficient.

Formally, the Atkinson index (A) is defined as:

$$A = \begin{cases} 1 - \left[\frac{1}{n}\sum_h \left(\frac{y^h}{\overline{y}}\right)^{1-\theta}\right]^{\frac{1}{1-\theta}} & \text{if } \theta \neq 1 \\ 1 - \prod_h \left(\frac{y^h}{\overline{y}}\right)^{\frac{1}{n}} & \text{if } \theta = 1 \end{cases}$$

where y^h is the income of household h, $\overline{y}$ is the average income, and the level of sensitivity is conveniently provided by the inequality aversion parameter (θ), which defines each Atkinson index.

To illustrate the differences between both inequality measures, in Figure 3-3 we plot the Lorenz curves for Bulgaria in 1978 and 1996.[21] Both curves intersect once and this points to an important change in the distribution of income. For instance, in 1996 both the lowest and highest deciles increased their income share[22], while intermediate deciles experienced a relative decrease. The Gini coefficient, however, did not change (26.5). In contrast, the Atkinson index with an inequality aversion parameter of one, decreased more than a point (from 11.5 to 10.4), reflecting the gain of the lowest income quintile against medium income households.

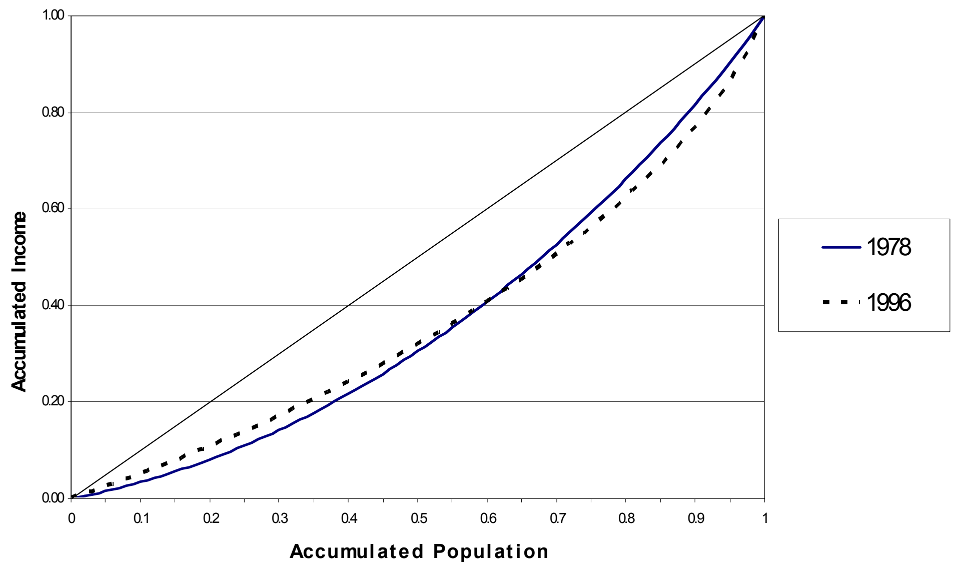
Therefore, by using both the Gini coefficient and the Atkinson indexes we can be more certain about variations in the whole income distribution. If both inequality measures move in the same direction our conclusions are more robust. If both measures behave differently this is an indication that the choice of a particular inequality index is important, since the weighting assigned to different parts of the distribution is relevant. Thus, the information given by both indexes is complementary and gives us a better understanding on how the income distribution is behaving.

We use four different values of θ to obtain more information on the inequality trends (0.5, 1, 1.5 and 2). It is known that for values above one, the Atkinson index is very sensitive to abnormally low incomes (Cowell, 1995). With this in mind, we estimate the four values to have a broader picture of how θ affects the levels and trends of inequality.

[21] The curves are estimated using the technique we describe below.

[22] This can be observed by a steeper curve in these population segments.

Figure 3-3: Bulgaria, Gross-Household series, Lorenz curves for two different years



The most common used values of θ is 0.5 and 1 (Atkinson, *et al.*, 1995; Burniaux, *et al.*, 1998). In the macro literature, the conceptually equivalent risk aversion parameter is estimated to be less than one.

Finally, some studies do not rely entirely on indexes and use share data directly to asses inequality behavior. Indeed, one can compare the ratio of the first and fifth quintile or the extreme deciles to obtain inequality information. These ratios provide information on the gap between the richest and poorest households of the population and can be used to assess inequality dispersion.

However, this method has some drawbacks. The most relevant is that it does not consider the distribution within income shares. The lowest quintile income share can remain unchanged, even if the poorest individuals are worst off. On the other extreme, the highest quintile share can also remain constant even when the richest individuals are much better off. Such intra-share changes in inequality are not measured by this kind of ratios, but are taken into account in an indicator like the Atkinson index. Moreover, there is no clear indication of which ratios are to be used and employing the extreme shares does not assure that we are comparing poor and rich individuals, since many poor people can be represented by middle shares in countries with widespread poverty. Finally, the ratios completely ignore the behavior of the middle income households. It can be the case that the ratio of the extreme quintiles is unchanged while the middle income shares are diminishing and thus, the distribution of income is being polarized.

For the reasons listed above, we do not use such ratio measurements in this paper and instead focus on inequality indexes that provide information for the whole income distribution. To obtain information about poverty and how poor individuals are faring with respect to the rest of society, we can directly estimate the extent of poverty from the Lorenz curve. This estimation is more useful than just assuming that a particular share is representative of the poor and provides us with better information in order to assess poverty.

Additionally, it is also of practical importance to know the actual shape of the Lorenz curve, which can be directly used to asses and compare inequality. Although this information cannot be directly employed in econometric models, it provides useful information for country-specific inequality analysis and greater detail on the actual inequality experience of each country.

Finally, we could use our country results to estimate the world's Lorenz curve. However, this task has additional limitations (i.e. lack of inequality data in many countries) that require further assumptions, which exceed the scope of this paper. On the other hand, this estimation has already been done by Sala-i-Martin (2000a,b) and we do not expect our estimations to alter the results found in these papers.[23]

3.4.2 Parametric estimation of the Lorenz Curve from grouped data

To construct inequality measures from grouped data we must first obtain the Lorenz curve. There are two approaches to obtain the Lorenz curve from grouped data: simple interpolation and methods based on parameterized Lorenz curves. As explained by Datt (1998) the second method is preferred for its relative accuracy.

[23] Although our country-specific Lorenz curve may be better estimated, in general the trend of reduced global inequality driven by high growth rates in China, can hardly be offset by such estimation improvements.

3.4.2.1 Valid Lorenz curves

Parametric estimation implies choosing a specific functional form and then estimating the underlying parameters. After the parameters are obtained, the Lorenz curve can be easily calculated. Nevertheless, in order to be considered as a legitimate Lorenz curve a functional form must comply with certain conditions.

If p is the cumulative proportion of population and $L(p)$ is the cumulative income share of group p. $L(p)$ is a valid Lorenz curve if and only if:

$$L(p) \geq 0 \quad \forall p \in (0,1) \tag{3.1}$$

$$L(0) = 0 \tag{3.2}$$

$$L(1) = 1 \tag{3.3}$$

$$L'\left(0^{+}\right) \geq 0 \quad \text{for } p \in (0,1) \tag{3.4}$$

$$L''(p) \geq 0 \tag{3.5}$$

3.4.2.2 Functional forms

There is a large literature concerning Lorenz curves estimation and there are many proposed functional forms. Some models are better suited for specific distributions and others perform better on typical distributions. However, given that income distributions can differ widely across countries and time, not one functional form is always preferable. To deal with this fact, we use the most popular functional forms and for each case, we choose the one that gives a better estimation.

We describe next each functional forms and the parameter constraints that assure a valid Lorenz curve.

The General Quadratic Lorenz curve.[24] In this model the Lorenz curve is given by:

$$L_{GQ}(p) = -\frac{1}{2}\left[bp + e + \left(mp^{2} + np + e^{2}\right)^{\frac{1}{2}}\right] \tag{3.6}$$

[24] **Villaseñor and Arnold (1989).**

where $e = -(a+b+c+1)$; $m = b^2 - 4a$ and $n = 2be - 4c$. The parameters to be estimated are then: a, b and c. In order for (3.6) to represent a valid Lorenz curve we must have: $m < 0$, $e < 0$, $c \geq 0$ and $a + c \geq 1$

Pareto Family of Lorenz Curves. A group of functional forms has been derived from the well-known classical Pareto Lorenz curve. The main difference between these models is the number of parameters employed.

- P0: Classical Pareto. This functional form is given by:

$$L_{P0}(p) = 1 - (1-p)^{\pi} \tag{3.7}$$

A valid Lorenz curve is obtained when: $0 < \pi \leq 1$.

- P1: Ortega *et al.* (1991):

$$L_{P1}(p) = p^{\alpha}\left[1 - (1-p)^{\pi}\right] \tag{3.8}$$

where the necessary conditions for a valid Lorenz curve are: $0 < \pi \leq 1$ and $\alpha \geq 0$. If $\alpha = 0$ then P1 reduces to P0.

- P2: Rasche *et al.* (1980). Here we have:

$$L_{P2}(p) = \left[1 - (1-p)^{\pi}\right]^{\beta} \tag{3.9}$$

where the necessary conditions are: $0 < \pi \leq 1$ and $\beta \geq 1$.

- P3: Sarabia *et al.* (1999). Combining P1 and P2 they propose:

$$L_{P3}(p) = p^{\alpha}\left[1 - (1-p)^{\pi}\right]^{\beta} \tag{3.10}$$

where $0 < \pi \leq 1$, $\alpha \geq 0$ and $\beta > 0$ assure a valid Lorenz curve.

Kakwani and Podder (1973). They suggest the following functional form to estimate the Lorenz curve:

$$L(p) = p^{\alpha} \exp^{\beta(p-1)} \tag{3.11}$$

A valid Lorenz curve is obtained when $1 < \alpha < 2$ and $\beta > 0$

The Beta model.[25] The is Lorenz curve is given by:

$$L_B(p) = p - \theta p^{\gamma} (1-p)^{\delta} \tag{3.12}$$

where θ, γ and δ are the parameters of the model to be estimated and we need for a valid curve that: $\theta > 0$, $0 < \gamma \leq 1$ and $0 < \delta \leq 1$. However, in many cases $L_B(p)$ fails condition (3.1) even when the parameters have the right values. This is an important shortcoming of the Beta model, but we consider it here since it is one of the best performers (Datt, 1998) and the negative values it produces in the lower tail of the distribution can be easily detected.

Sarabia et al. (1999). They propose a four-parameter functional form to correct for the Beta model problem. We name their equation the BS model:

$$L_{BS}(p) = p^{\lambda+\gamma} \left[1 - \theta (1-p)^{\delta}\right]^{\gamma} \tag{3.13}$$

where $0 \leq \theta \leq 1$, $\gamma \geq 1$, $0 < \delta \leq 1$ and $\lambda \geq 0$ assure that $L_{BS}(p)$ is a valid Lorenz curve.

3.4.2.3 Estimation and selection of the Lorenz curve model

In total we can use seven different models to estimate the Lorenz curve. Some of these can be linearized to use ordinary least square estimations, but others cannot. Therefore, we have employed a non-linear estimation program in the General Algebraic Modeling System (GAMS) software to test the seven parametric models. We also check if each model complies with the conditions to be taken as a valid Lorenz curve. When more

[25] **Kakwani and Podder (1976) and Kakwani (1980).**

than one model yields a valid Lorenz curve we use the standard procedure adopted in the literature and choose the model that yields a lower sum of squared residuals.

In our view, this non-linear estimation of parametric models is an improvement with respect to existing software. The POVCAL software (Chen *et al.*, 1998) only estimates linearized models of the General Quadratic and Beta models. In some cases both models fail to provide a valid Lorenz curve and in addition, this software does not correct for Beta models that generate negative values at the bottom of the distribution.[26]

Furthermore, our GAMS program calculates the underlying income distribution associated with the estimated Lorenz curve. Using this information it directly estimates the Gini coefficient, the Atkinson index for the four different θ values and the poverty ratios. Nonetheless, not all the series present in the WIID database have grouped data information. Therefore, the number of Gini and Atkinson indexes estimations we obtain is lower than the number of Gini coefficients provided by primary sources. This limits the analysis but provides additional information not present in the source Gini coefficients.

3.4.3 Poverty estimation

It is straightforward to conduct poverty analysis once the entire income distribution is estimated and this procedure has the advantage of not relying on the strong assumption that the poor people are well represented by the lowest quintile or decile.

To estimate poverty ratios, we use the official World Bank absolute poverty lines of one and two dollars a day (Ravallion *et al.*, 1991). The income levels are taken from the PPP-adjusted GDP values of the Penn World Tables version 6.1.[27]

However, the use of GDP data as an income indicator is problematic. First, inequality data is drawn from household surveys and there is a substantial discrepancy between the national income reported from these household surveys and that from national accounts data. The difference is mainly explained because GDP not only includes private consumption, but also private investment and government spending. Secondly, the poverty lines

[26]Nevertheless, we use POVCAL to estimate the cumulative income shares $L(p)$ when these are not provided directly by the source and instead, the grouped data is presented by income classes or the income data is associated with mean income and/or upper limit values.

[27]The poverty lines were reported in 1985 values and the PWT data is in 1996 dollars. Thus, the equivalent annual income of \$1/day is \$532 and for \$2/day is \$1064 (Sala-i-Martin, 2000a,b).

were calculated using mean consumption levels in poor countries and therefore, include only the most basic consumption needs and it does not take into account public services or investment.

Following these considerations, there are two main approaches to estimate absolute poverty. The World Bank (Chen and Ravallion 2001, 2004) uses consumption and inequality data both drawn from household surveys. On the other hand, we follow Sala-i-Martin (2000a,b) and estimate poverty using inequality data from household surveys, but per capita income from national accounts data. This later approach allows us to have larger series. In addition, it indirectly accounts for governmental expenditure and other non-private consumption sources of goods and services for the poor.

The two methods can yield significantly different poverty estimates for a given poverty line. Yet, in recent articles (Chen and Ravallion, 2004; and Ravallion, 2004) it is shown that both methods produce very similar results when the World Bank method uses the \$1/day poverty line and the other method uses \$2/day. Moreover, Ravallion (2001) finds that, with the exception of the transition economies of Eastern Europe and the Former Soviet Union, growth rates of national accounts measures are not systematically different from growth rates of household survey measures.

To sum up, using a \$2/day poverty line we obtain poverty estimates roughly equivalent to the \$1/day absolute poverty based on consumption and we expect that this equivalence does not change over time.

Formally, the headcount poverty ratio is defined as the number of individuals with an income below the poverty line in relation to the total population:

$$PR_z = \frac{\int_0^z I(y)dy}{\int_0^\infty I(y)dy} \tag{3.14}$$

where z is the poverty line and $I(y)$ is the distribution function of income y.

We estimate the poverty ratio using a GAMS program similar to the one used in our previous section. Nonetheless, since we are now primarily interested in the lower tail of the income distribution, we select the model that provides a valid Lorenz curve that fits best the lower quintile of the distribution.

Moreover, we use a discrete version of the previous formula. For instance, we divide the population in a thousand q units and estimate the income of each unit using the formula:[28]

$$I(q) = GDP_{pc} * IS(q) * 1000 \tag{3.15}$$

where $GDPpc$ is gross domestic product per capita, $IS\,(q)$ is the income share of unit q. The poverty ratios are given by the sum of the number of units with an income below the two poverty lines ($1/day and $2/day), divided by the total population. The total number of poor can easily be obtained by multiplying the poverty ratio by the total population. With a similar procedure, it is also straightforward to estimate other poverty indexes, such as the poverty gap and the Foster-Greer-Thorbecke index.

3.5 Evaluating the data

On the basis of the previous two sections we can construct a whole new inequality dataset. Unlike the DS series, we have several inequality series for each country (gross-household, net-household and expenditure-person), five inequality measures (Gini coefficient and four Atkinson indexes) and two types of series based on their reliability (basic and extended).[29]

To assess the implementation of our comparability assumptions, we begin by comparing our dataset with that of DS. We also check how our estimated Gini coefficients fare with respect to the source information provided in the WIID. Finally, we analyze if the Gini coefficient and the Atkinson index yield similar results.

3.5.1 Differences between our series and Deininger and Squire

In general, we want to know if inequality levels and time trends are significantly different when moving from one dataset to the other. Inequality levels are important for cross-

[28] The formulas to integrate the poverty ratios are complicated by the non-linearity of most of the Lorenz curve models. However, with a thousand units we have a three digit approximation of the real value.

[29] The summary statistics for the Atkinson indexes and our own Gini estimates are given in Figure 3-5, under the column: share data.

country comparisons. Time trends provide information on specific country inequality behavior and are relevant when analyzing pooled data.

3.5.1.1 Level differences

AB find serious level differences when comparing the DS dataset with the LIS information. Since the LIS was conducted explicitly to render OECD inequality data comparable, it is a reliable indicator on which to compare the inequality levels for OECD countries. Thus, we follow AB and compare our dataset with the LIS information for a single year (1991 or the closest available). The results are shown in Table 3-7.

Table 3-7: OECD country rankings, Gini Gross-Household series for one year

Year		LIS		Basic series		DS-accept	
1991	1	Finland	29.61	Finland	29.61	Finland	26.11
1991	2	Netherlands	30.59	Netherlands	30.59	Belgium	26.92
1992	3	Sweden	31.11	Sweden	31.11	Canada	27.65
1983	4	Germany	31.37	Germany	31.37	Great Britain	27.80
1991	5	Norway	31.81	Norway	31.81	Netherlands	29.38
1992	6	Belgium	31.95	Belgium	31.95	Germany	31.37
1992	7	Denmark	33.20	Denmark	33.20	Sweden	32.44
1984	8	France	34.91	France	34.91	Denmark	33.20
1991	9	Canada	35.08	Canada	35.08	Norway	33.31
1986	10	Great Britain	36.18	Great Britain	36.18	France	34.91
1991	11	USA	39.15	USA	39.15	USA	37.94

For this specific year our observations are exactly the same as the LIS and thus we have the same country ranking. However, the ranking provided by the DS series is very different. This is evident in the low inequality reported for Canada and Great Britain, and the higher inequality in Sweden and Norway.

The results are similar when evaluating net income rankings. In Table 3-8 our basic series is almost identical to that of LIS, the only difference being the observation for Italy, which is very low in the LIS series. On the other hand, the DS accept series once more provides a completely different ranking. Spain has a very low inequality (e.g. DS use expenditure information for this country). Once more, Great Britain and Canada have unexpected low positions and Sweden and Norway very high ones.

One can argue that a one-year ranking is not adequate, since a single uncharacteristic observation can alter the ranking. Thus, we also rank the OECD countries by the average Gini coefficient for a five year period: 1983-1987. We choose this period since it provides

Table 3-8: OECD country rankings, Gini Net-Household series for one year

Year		LIS		Basic series		DS-accept	
1991	1	Finland	26.11	Finland	26.11	Spain	25.91
1992	2	Belgium	26.92	Belgium	26.92	Finland	26.11
1991	3	Italy	27.12	Norway	28.80	Belgium	26.92
1991	4	Norway	28.80	Sweden	29.16	Canada	27.65
1992	5	Sweden	29.16	Germany	29.36	Great Britain	27.80
1983	6	Germany	29.36	Netherlands	29.38	Netherlands	29.38
1991	7	Netherlands	29.38	Denmark	29.96	Germany	31.37
1992	8	Denmark	29.96	Spain	*30.60*	Italy	32.19
1990	9	Spain	*30.60*	Canada	31.47	Sweden	32.44
1991	10	Canada	31.47	France	31.94	Denmark	33.20
1984	11	France	31.94	Italy	32.19	Norway	33.31
1989	12	Australia	32.85	Australia	32.85	France	34.91
1986	13	Great Britain	33.29	Great Britain	33.29	Australia	37.32
1991	14	USA	35.24	USA	35.24	USA	37.94

Note: For Spain the LIS observation refers to family equivalent and we only have an extended series.

the most LIS observations possible. The results for the Gross-Household series are given in Table 3-9 and those for Net-Household in Table 3-10.

Table 3-9: OECD Country rankings, Gini Gross-Household average for 1983-1987

	LIS		Basic series		DS-accept	
1	Belgium	26.22	Belgium	26.22	Belgium	26.22
2	Finland	30.10	Norway	29.44	Great Britain	27.14
3	Norway	30.36	Finland	*30.30*	Netherlands	28.94
4	Sweden	30.77	Sweden	30.77	Finland	29.34
5	Germany	31.78	Netherlands	31.69	Sweden	31.30
6	Netherlands	32.94	Germany	31.78	Norway	31.69
7	Denmark	33.15	Denmark	33.15	Germany	31.78
8	Canada	34.28	Canada	34.67	Canada	32.67
9	France	34.91	France	34.91	Denmark	33.15
10	New Zealand	*35.00*	New Zealand	35.48	France	34.91
11	Great Britain	36.18	Great Britain	36.18	New Zealand	35.48
12	Australia	36.50	USA	38.93	USA	37.20
13	USA	39.23	Australia	39.09	Australia	39.09

Notes: The New Zealand LIS observation refers to family.
We only have an extended series for Finland.

Again our basic series for gross income ranks the OECD in a very similar way as the LIS. Nevertheless, the DS series once more yields an unsuitable ranking. Great Britain is again very highly positioned and Norway too low. When turning to Net income, some differences appear between our data and the LIS with respect to Finland, Great Britain and Australia. The difference is justified in these cases by the existence of more observations in our basic series than those in the LIS dataset.

Table 3-10: OECD Country rankings, Gini Net-Household, average for 1983-1987

	LIS		Basic series		DS-accept	
1	Finland	26.19	Belgium	26.22	Belgium	26.22
2	Belgium	26.22	Norway	26.87	Great Britain	27.14
3	Netherlands	28.35	Sweden	28.64	Netherlands	28.94
4	Norway	28.35	Netherlands	28.94	Finland	29.34
5	Sweden	28.64	Finland	29.34	Sweden	31.30
6	Germany	29.36	Germany	29.36	Norway	31.69
7	Canada	31.21	Canada	29.87	Germany	31.78
8	Denmark	31.30	Denmark	31.30	Canada	32.67
9	Australia	31.49	Great Britain	31.84	Denmark	33.15
10	France	31.94	France	31.94	Italy	33.80
11	Italy	32.78	Italy	33.80	France	34.91
12	Great Britain	33.29	USA	34.93	USA	37.20
13	USA	35.24	Australia	36.24	Australia	39.09

We can draw two conclusions from this example of OECD countries. First, our series are compatible with the LIS and the few divergences are justified by more observations present in our series. The second conclusion is that the DS accept series yields some rankings that are very hard to justify and can only be explained by the inconsistent use of different income concepts.

In the case of developing countries, there is no equivalent to the LIS that we can use as a benchmark to compare datasets. However, the differences in levels between our basic series and the DS accept series persist. When ranking OECD countries some differences were produced by the loose interchange of net and gross income data present in the DS series. For developing countries the source of divergence in levels is produced by mixing expenditure and gross income data. Since expenditure data is significantly lower than gross income data, this alters the inequality levels between countries.

Thus, these results support the idea that mixing different income concepts can lead to misguiding conclusions and is an important limitation for inequality cross-country studies.

3.5.1.2 Time trend differences

The use of fixed adjustments can help to render the DS series comparable to the LIS. If provisions are made to adjust for income differences some of the divergences shown before disappear. However, there is still an element of arbitrariness in the procedure. How much should we adjust the series? The countries average difference or the overall average? The decision of which fixed adjustment to apply affects the outcome and the resulting country ranking.

Nevertheless, the fixed adjustment of series can mainly solve for level-differences. For instance, when the time trend of series with different definitions is not the same, mixing the series creates a whole new trend. In many cases, the use of a fixed adjustment does not correct for this problem

Therefore, we proceed to compare the time-trend differences between our basic and extended series, with the DS accept series. To do so, we regress the Gini coefficient against time for those series that have five or more observations. We use two equations:[30]

$$G_i = \alpha_1 + \alpha_2 t_i + \varepsilon_i \tag{3.16}$$

$$G_i = \beta_1 + \beta_2 t_i + \beta_3 t_i^2 + \varepsilon_i \tag{3.17}$$

The first regression tests for any linear time trends and the second for quadratic trends. To do a valid comparison of our series with the DS-accept series we do not take into consideration information that was not available to DS (data after 1996 and studies published after this year). We limit the analysis to the three main series: gross-household, net-household and expenditure person. In total we study 44 basic series, 39 extended series and 43 DS-accept series with five or more observations, where 45% of the series come from OECD countries.

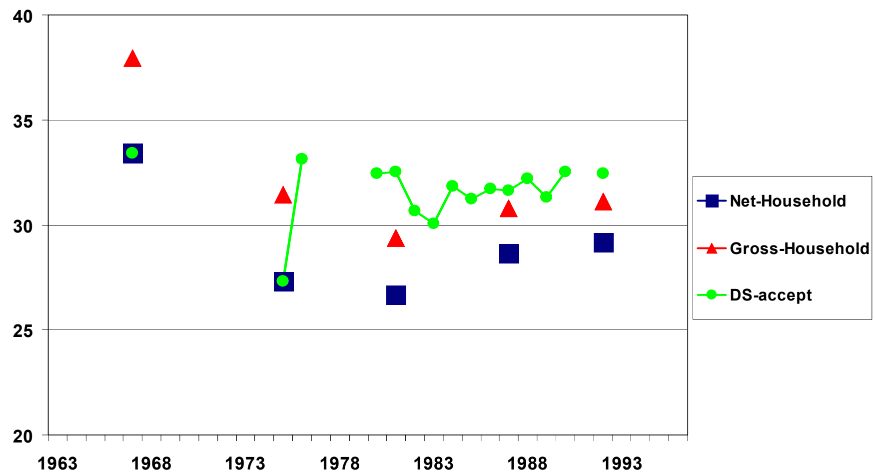
To have a statistically significant trend, we must find that α_2 is significantly different from zero and/or that β_1 and β_2 are jointly different from zero. In total, 27% of the basic series have a different time trend than the DS-accept series. The figure is over 66% in the case of the extended series.

We can also analyze only those series with ten or more observations. Although the number of series decreases, the results remain the same. Now we have 30 basic and 16 extended series. 23% of the basic series and 73% of the extended series have a different time trend than the DS-accept series. The OECD proportion of observations remains close to half (47%). These comparisons show significant time trend differences between both datasets, which suggest dissimilar conclusions regarding inequality variation.

[30] Given the small number of observation per country, we do not pursue a time series analysis.

As an example we can analyze the case of Sweden. There are two main sources: LIS and SAS for the net and gross household series. However, when conducting our comparability analysis both sources are rendered not comparable and thus, we do not mix them. Instead, we choose the LIS series, which has fewer observations than the SAS series, but has a larger time span and provides better cross-country comparisons, as explained above. In Figure 3-4 we plot our two basic series and the DS-accept series. The first two DS observations are taken from the LIS series but the remaining are from the SAS series.

Figure 3-4: Sweden, Gini coefficient, basic and DS-accept series



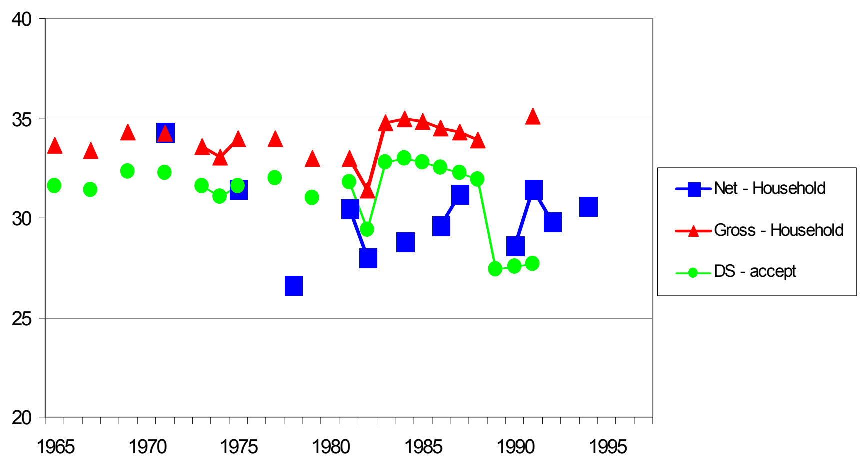
The first conclusion is that the levels are different. This was already analyzed in the previous section. More interesting, the DS-accept series fails to show any time trend at all.[31] From our basic series it is clear that there is a trend, in particular an U-shape pattern. This also is corroborated by significant coefficients in the quadratic regression.

Another compelling example is Canada. In Figure 3-5 we present two of our basic series and the DS-accept. The gross-household series has been adjusted by the LIS data and thus, has a different level from the DS-accept series until the 1988 observation. After this year the series diverge, DS take some observations which fail the WIID quality criteria. Instead, we use the LIS observation for 1991. Although the last part of the two series

[31] This could be corrected by a fixed adjustment of the Net series. However, DS do not find any significant difference between Net and Gross, and thus, do not apply any adjustment.

show different time trends, there is no overall time trend in both series. However, when using our basic net-household series, we have significant coefficients for the quadratic form regression, i.e., we spot again the U-shape pattern. This is a clear example of the limitations of the dummy variable solution, since a fixed adjustment does not solve the inconsistency and the time trend does not change.

Figure 3-5: Canada, Gini coefficient, basic and DS-accept series



There are many other cases that show different levels and time trends between our series and the DS-accept. Most of the differences are given by the grouping approach used; not only by the income concepts and reference units, but also by incompatible sources.

On the other hand, there are some differences that are provided by the quality labels used. In the example of Canada we rejected some observations used in DS. Sometimes it is the other way around and we include observations not labeled as high quality by DS, but accepted as reliable data by the WIID.[32]

Yet, the differences are significant and it is clear that the decision regarding how to group different definitions is essential when analyzing inequality data. Furthermore, the use of dummy variables or fixed adjustments does not solve the problem satisfactorily.

[32] A noteworthy example is the data presented by Paukert (1973). Most of his data is not accepted by DS because it lacks a clear reference to the primary source. However, the WIID accepts all his observations. Since this sources provide information from the 1960s it expands many country series and thus, may alter the time trends for some countries. For instance, Barro (2000) also uses observations that do not pass this "primary source" quality test.

However, in a recent paper Deininger and Squire (2002) continue to justify this practice. In our view, the empirical work they cite to back their claims is not compelling and does not tackle the main limitations pointed in this paper. Therefore, we agree with Atkinson and Brandolini and conclude that researches employing inequality data should avoid the use of the DS-accept series.

On the other hand, our dataset does comply with the recommendations of AB:

> "A secondary dataset should be a consolidation of earlier work, with multiple observations for the same country and the same date being justified by differences in source, in definition, or in methods of calculation."

3.5.2 Compatibility of the source and our estimated Gini coefficients

We turn our attention to our own estimated Lorenz curves. It is a straightforward exercise to obtain Gini coefficients and Atkinson indexes once the Lorenz curve is estimated.[33] With this information we first test how our estimated Gini coefficients compare to the source information. Afterwards, we analyze if the Atkinson indexes provide different information on inequality than the Gini coefficient.

In order to compare our Gini estimates and the source coefficients we conduct the same comparability tests done in section 3.2.1. Again we limit our analysis to the three main series: gross-household, net-household and expenditure-person. We obtain 87 comparable series. 85% are fully comparable with c (an average difference) of zero; 92% are fully comparable with $c \leq |1|$ and 95% with $c \leq |2|$. The average difference between series is of -0.15 and the average difference weighted by the number of observations is -0.17. These results show a very good estimation of the Gini coefficients from our constructed Lorenz curves.

Moreover, of the three non-comparable estimations, two of them can be explained by inconsistent source information for one particular year. For example, the expenditure-

[33] Using the chosen parametric equation we can construct the whole income distribution and the Lorenz curve. In our case, we use centuples to do so. The resulting inequality indexes do not change significantly if a lower unit is used.

person series of Estonia has a Gini estimate of 39.47 for 1993 while the source value is of 31.52 (see Table 3-11). This single observation renders the series incompatible.

Table 3-11: Estonia, Expenditure-person series, accumulated share data and Gini coefficients

Year	Gini (source)	Gini est.	Quintile 1	Quintile 2	Quintile 3	Quintile 4
1992	35.82	35.79	0.0702	0.1879	0.3490	0.5691
1993	**31.52**	**39.47**	0.0624	0.1717	0.3245	0.5386
1995	36.63	36.60	0.0691	0.1835	0.3423	0.5623

Source: WIID and DS datasets.

Yet, a closer inspection of the accumulated share data reveals a source inconsistency. All the quintile accumulated income shares are higher for 1992 and 1995, but the Gini coefficients are higher in both years. This result is very contradictory and can be a typo in the source data.[34] On the other hand, our Gini estimate is consistent with the share information.

In general, our estimates are very close to the source information and we can be assured of the quality of our estimated Gini coefficients. This also provides confidence in the values of the Atkinson indexes that also use the constructed Lorenz curves. Li *et al.* (1998) state that the estimation methods vary across different sources and therefore, the use of one standard technique can minimize this problem. Consequently, our main series include our own estimations when there is income share information, and the source data when there is no way to estimate the Gini coefficient and the Atkinson indexes. Nevertheless, we retain both observations: the source and our estimations and hence, we can test the robustness of our results.

[34] From equations 4 and 5 we know that the Lorenz curve for 1992 and 1995 dominates that of 1993 for the whole middle section of the distribution. Thus, the lower Gini value given by the source could only be justified by significant differences at the extremes of the distribution in 1993. However, given that the Gini attaches more weight to the middle of the distribution this possibility seems very unlikely.

3.5.3 Comparing inequality measured by the Gini coefficient and the Atkinson indexes

We analyze now the inequality results provided by the two measures we have. First, we check the behavior of the Atkinson index for each value of θ. Afterwards we compare the Atkinson and the Gini data. Throughout the section, we only use our estimated Gini coefficient. In this way we have the same sample as the Atkinson index and both indexes are derived from the same estimation technique.

3.5.3.1 Atkinson index results

As expected, the level and variance of the Atkinson index increases with θ. The higher the inequality aversion, the index gives higher values and also is more sensitive to changes in the distribution (see Table 3-12). The overall values of the Net series are lower than those of the Expenditure series because the first series has more OECD countries in its sample and the later more developing countries. In fact, from the Gini coefficient values we know that the level of inequality and the variance between OECD countries is smaller than for developing countries. The gross-household series has intermediate values, reflecting the more balanced sample.

Table 3-12: Basic series, averages for different indexes

	Gross-household		Net-household		Expenditure-person	
	Mean	Variance	Mean	Variance	Mean	Variance
Atkinson θ = 0.5	14.27	6.92	8.59	2.64	12.72	10.82
Atkinson θ = 1	25.96	17.31	16.90	8.00	22.63	29.66
Atkinson θ = 1.5	36.90	35.57	25.53	22.68	30.64	50.23
Atkinson θ = 2	46.63	65.71	34.91	69.50	37.27	69.23
Gini coefficient	39.90	10.39	31.38	5.81	38.53	18.96

3.5.3.2 Differences between the Atkinson and the Gini

In the following sections we analyze only the two middle Atkinson indexes. Furthermore, we only consider series with five or more observations. First we explore how the two inequality indexes rank a sample of countries and then we analyze the time-trend information that both indexes provide.

Level differences. In Table 3-13 we rank a sample of 13 countries by the three inequality indexes. We use the average for the whole 1980s decade. In general, the Atkinson ranking is very similar to that given by the Gini coefficient. The only significant difference is in the ranking of Canada and Bangladesh for the last index. Thus, although the Atkinson provides lower levels, this does not change much the relative position of each country.

Table 3-13: Basic Gross-Household series, 1980s average ranking

	Gini coefficient		Atkinson θ=1		Atkinson θ=1.5	
1	Bulgaria	22.66	Bulgaria	8.79	Bulgaria	12.59
2	Germany	31.38	Germany	16.78	Germany	25.62
3	Canada	33.94	Spain	18.84	Spain	27.57
4	Spain	34.27	Canada	19.40	Bangladesh	29.09
5	Japan	34.49	Japan	19.61	Japan	29.28
6	Korea	36.19	Korea	21.20	Korea	30.13
7	Bangladesh	37.52	Bangladesh	21.71	Canada	32.79
8	Australia	38.40	Australia	23.75	Australia	34.64
9	United States	38.45	United States	25.76	Hong Kong	36.17
10	Hong Kong	42.40	Hong Kong	26.71	United States	40.19
11	Bahamas	43.15	Bahamas	32.89	Colombia	50.48
12	Colombia	51.16	Colombia	37.89	Bahamas	50.81
13	Brazil	56.98	Brazil	45.45	Brazil	57.96

However, it does provide more information about the income distribution. In Table 3-14 we rank the 12 countries that have a Basic Net-Household series with more than five observations. The general ranking does not change much from index to index, but it does have some interesting cases. For example, Sweden has a relatively high inequality for the Atkinson indexes and Italy a relatively low one. A closer inspection of the share information shows that the difference is provided by the lowest quintile, where Italy has an average share income of 8.2% and Sweden of 7.4%. Therefore, Italy has a lower Atkinson level when $\theta = 1.5$. Yet, the middle quintiles are very similar and Sweden's highest quintile has around four points less. This explains the lower inequality when $\theta = 1$ or when we use the Gini coefficient.

Table 3-14: Basic Net-Houshold series, 1980s average ranking

	Gini coefficient		Atkinson θ=1		Atkinson θ=1.5	
1	Romania	23.37	Romania	9.04	Romania	13.74
2	Poland	25.06	Norway	9.74	Poland	14.29
3	Norway	26.84	Poland	9.87	Norway	17.20
4	Sweden	27.44	Netherlands	12.50	Netherlands	18.89
5	Netherlands	27.44	Germany	13.26	Germany	19.86
6	Germany	28.58	Sweden	13.67	Italy	20.29
7	Taiwan	29.09	Taiwan	13.82	Taiwan	21.05
8	Italy	30.18	Italy	14.11	Sweden	21.93
9	Finland	30.94	Finland	16.88	Finland	25.24
10	Great Britain	33.76	Great Britain	17.01	Great Britain	25.39
11	United States	34.43	United States	21.67	United States	34.08
12	Mexico	46.45	Mexico	31.42	Mexico	41.79

Time trend differences. To analyze the time trend information provided by both inequality measures, again we only use the three main series for those countries with five or more observations, and we regress once more equations (3.16) and (3.17).

We have a sample of 60 observations, 23 of those representing OECD countries. For $\theta = 1$ the Atkinson index yields a different time trend in 27% of the cases from that of the Gini coefficient. When $\theta = 1.5$ the difference increases to 31%. If we expand the series length to ten or more observations the results are similar: 31% for $\theta = 1$ and 28% for $\theta = 1.5$. Therefore, we can conclude that the Atkinson index gives a different time trend in roughly a third of all cases. This is a significant divergence and confirms that the changes in different parts of the income distribution can be responsible for both indexes reporting different inequality results.

In summary, although there is not much variation in the ranking given by both inequality indexes, there are important differences when we analyze the time trend information. Thus, the Atkinson index *does* provide additional information about the extremes of the distribution and as such, it is a useful resource when analyzing inequality data.

3.5.4 International and intertemporal patterns of inequality

As before, in this section we use our three main series and in addition, we use our own Gini estimates when share data was available. First we summarize the characteristics of the basic and extended series and we end the section by analyzing how inequality varies across and within countries.

3.5.4.1 Descriptive information

In the Appendix we present the summary statistics for our six series: three basic and three extended. It is important to remember that the Net series are more representative of OECD countries, the Expenditure series of developing countries and the Gross series has a balanced sampled between both groups. The disparities in the sampled countries of each series makes it difficult to compare the results when different income concepts are used. Nonetheless, a simple examination of the tables shows that the Gross series has a higher mean and standard deviation in average. The standard deviation is very similar for the Net and Expenditure series, while the last has a higher mean in average.

We also present for each series the results of a simple ANOVA analysis, which shows the percentage of variation represented by between and within country changes. The results are homogeneous and for each series between-country variation represents between 80% and 90% of total variation. This suggests that inequality levels are more important than inequality trends, a conclusion also reached by Li, *et al.* (1998). However, in our case the within-country variation is also significant and thus, we find evidence for the weaker hypothesis they test, i.e., that intertemporal shifts in inequality are modest compared with international differences.

3.5.4.2 International patterns

Using again the two time trend equations, we run random-effect regressions on the six series. Here we want to find if there are any inequality patterns that are common to countries and groups. The results are presented in Table 3-15.

First there is no linear pattern in the Gross and Net series. The Expenditure series has a significantly decreasing trend. However, all six series do present significant quadratic

Table 3-15: Random-effect regressions of Gini series

	Linear trend			Quadratic trend							
	α_2	std. error	p-value	β_2	std. error	p-value	β_3	std. error	p-value	obs.	countries
Basic series											
G - HH	0.005	0.016	0.75	**-0.312**	0.101	0.002	**0.005**	0.002	0.002	412	38
N - HH	-0.013	0.018	0.46	**-0.316**	0.074	0.000	**0.005**	0.001	0.000	279	20
E - P	**-0.057**	0.027	0.03	**-0.311**	0.128	0.017	**0.004**	0.002	0.045	123	21
Extended series											
G - HH	-0.017	0.015	0.26	**-0.430**	0.086	0.000	**0.006**	0.001	0.000	573	52 a/
N - HH	-0.004	0.016	0.80	**-0.344**	0.093	0.000	**0.005**	0.001	0.001	376	31 b/
E - P	**-0.063**	0.023	0.01	**-0.252**	0.118	0.034	**0.003**	0.002	0.105 c/	173	28

Notes: a/ Uses the joint series SUN/RUS. The results do not change if the series are separated.
b/ Uses the joint series CSK/CZE and CSK/SVK. The results do not change if the series are separated.
c/ We reject the null hypothesis that both coefficients are zero at a 99% confidence level.

trends. In particular, all the series reveal a U-shape pattern with a significantly negative β_2 and a positive β_3. These results support the idea of what Atkinson (2003) labeled the "U-turn" pattern. This is represented by a decrease of inequality after the Second World War and a turning point around the 1980s when inequality began to increase again.[35] In particular, for the gross and net series the turning point is around the late 1970s and early 1980s. Although Atkinson finds this pattern for OECD countries, our results suggest that it can represent a broader phenomenon.[36]

Additionally, the values of the quadratic coefficients produce a different U-pattern for the extended expenditure series. It has a prominent decrease in inequality and the turning point is in the mid-1990s. These divergent results suggest that the sample of countries may be behaving differently, since expenditure series represent mainly developing countries. Therefore, to analyze this point further we divide the series by OECD and developing countries (see Table 3-16).[37]

For the case of OECD countries the results are very robust. For the four analyzed series there is no linear trend but a quadratic U-pattern. On the other hand, the series for developing countries present different results. The gross-household basic series does

[35] Li, *et al.* (1998) fail to find any significant time trend. However, they only use the linear approach.
[36] He explains this inequality behavior by a decrease in governmental intervention and the increase of more liberal economic policies.
[37] In the following tables and figures, LDCs refers to developing countries.

Table 3-16: Random effects regressions of Gini series for OECD and developing countries separated

	Linear trend			Quadratic trend								
	α_2	std. error	p-value	β_2	std. error	p-value	β_3	std. error	p-value	obs.	countries	
LDCs sample												
G - HH - basic	-0.022	0.027	0.40	-1.012	0.536	0.060	0.015	0.008	0.050 a/	224	22	
E - P - basic	**-0.056**	0.027	0.04	**-0.316**	0.130	0.016	**0.004**	0.002	0.043	120	20	
G - HH - ext	-0.025	0.023	0.26	**-1.358**	0.541	0.013	**0.018**	0.008	0.021	340	35	b/
E - P - ext	**-0.061**	0.024	0.01	-0.219	0.144	0.132	0.002	0.002	0.231 a/	151	24	
OECD sample												
G - HH - basic	0.030	0.018	0.10	**-0.265**	0.074	0.000	**0.005**	0.001	0.000	188	16	
N - HH - basic	0.014	0.018	0.43	**-0.267**	0.074	0.000	**0.005**	0.001	0.000	200	15	
G - HH - ext	-0.008	0.019	0.69	**-0.256**	0.083	0.002	**0.004**	0.001	0.002	233	17	
N - HH - ext	0.015	0.018	0.41	**-0.261**	0.071	0.000	**0.004**	0.001	0.000	220	18	

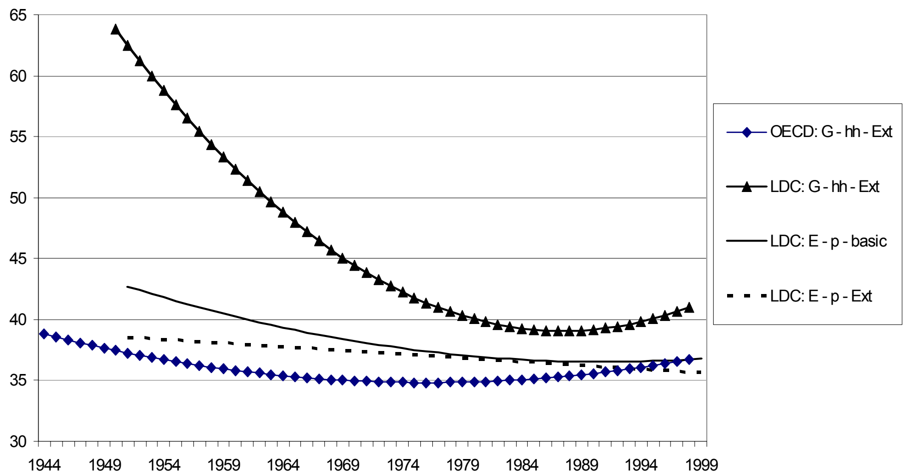
Notes: a/ We cannot reject the null hypothesis that both coefficients are zero.
b/ Uses the joint series SUN/RUS. The results do not change if the series are separated.

not have any significant trend, while the extended series presents the familiar U-pattern. Moreover, both expenditure series have decreasing linear trends and the expenditure basic series has a significant U-pattern. When the estimated regression curves are plotted[38] (Figure 3-6), we observe that the series for the developing countries are mainly decreasing in the period. The two series with a significant quadratic trend have their turning point late in the period: for the Gross extended series it is 1987 and for the Expenditure basic series, 1991. This generates a trend that is decreasing trough the period and increases slightly at the end.

Therefore, we can conclude that the OECD countries present a clear U-pattern time trend, with a turning point around the late 1970s. In addition, for developing countries inequality has been mainly decreasing in the period, with a slight increase in the 1990s. Subsequently, although inter-country inequality is more variable, within-country trends are also significant.

[38] We only plot the OECD Gross-extended series, since the other series are very similar. The linear trend of the developing countries Expenditure-extended series is very similar to that of the basic series and thus, it is not plotted either.

Figure 3-6: Estimated random-effects regression curves



3.5.4.3 Country-specific intertemporal patterns

Inequality changes over time have important policy implications. For instance, if inequality is stable over time, then economic growth has a direct impact on poverty reduction. Moreover, the particular level of inequality determines how much poor households benefit from countrywide growth. Conversely, significant shifts in inequality can offset pro-poor growth, signal important socioeconomic changes and strengthen the importance of redistributive policies.

We focus now in country specific inequality trends. Again we use a linear and a quadratic time trend to test for intertemporal variations in inequality, but now we use fixed-effects estimations to capture the individual coefficients for each country. Moreover, as some countries have more series with different reference units (i.e. household equivalent), then we incorporate these series into the pooling to obtain more country observations.[39]

In the Appendix (Tables 3-23 and 3-24) we present the fixed-effects regressions for the basic and extended gross-household series. For the basic series 24% had a significant linear trend, 24% a quadratic trend and 13% both. In total, 61% of the countries had some kind of time trend. For the extended series the numbers increase to 32%, 29% and

[39]Note that we are not mixing different reference units in a same series, but using distinctive series, each with a different reference unit.

26%, leaving the total to 87%. The differences between both series can be accounted by the increase in the individual country observations provided by the extended series.

Using panel data provides more observations per country and thus, a better approximation of the inequality trend. For each country we perform pool regressions using all its basic series and then asses if there is any significant trend at a 5% confidence level. Together with the results of the basic gross-household series, the results of the panel data estimations for all the series are presented in the appendix (Table 3-25). In some cases the panel data regressions confirm the results of the gross-household series, in others it provides other trends or produces a trend that was not present in the gross-household series. Additionally, 71% of the countries of the countries for which we regressed the pooled data have some kind of significant time trend.

The different results can be a consequence of several factors: the number of series in each country, the increased number of observations provided by the panel data analysis and the difference in definitions (income concept and/or recipient unit). Thus, with our inequality dataset there are many ways to analyze an individual country. The presence of series with different definitions, the basic series and the extended series and the Atkinson indexes provide a richer source of information from where to draw inequality conclusions. What seems to be evident is that many countries present inequality time series with some significant trends and thus, within-country inequality variations are important.

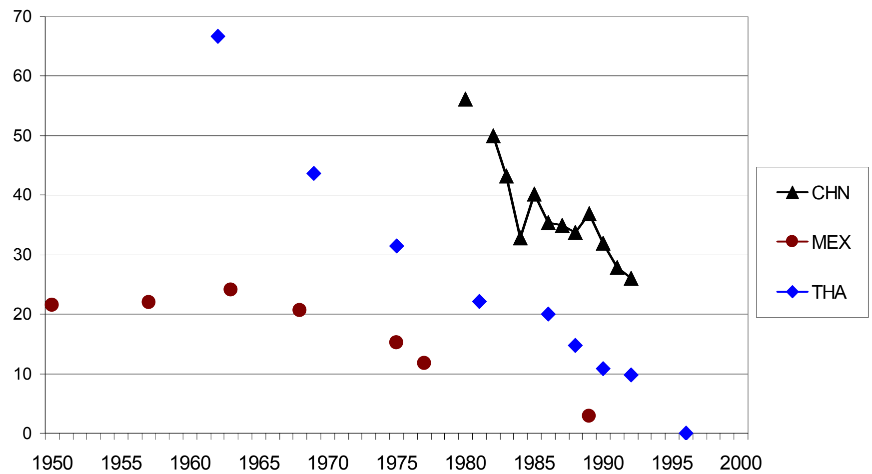
3.5.5 Poverty results

We estimate poverty ratios for the three basic extended series.[40] In general, poverty has been declining in most of the countries and as expected, there is no poverty in OECD countries. In many Asian and Latin American countries the ratios have been declining (i.e. China and India) and have become zero for some (e.g. Indonesia and Thailand). The exception is the African continent, were the ratios continue to remain high. These results

[40] Since some of the observations in these series have been adjusted for comparability reasons, we have some share data that does not correspond with the associated Gini values. Thus, we have some observations that present a lower or higher inequality than the non-adjusted data. There is no easy way to correct for this problem. However, since most of the adjustments were performed in OECD countries, the poverty results are mainly not affected, since there is usually no poverty in these countries. The three exemptions are the G-hh series for Brazil, Chile and Mexico. For Brazil we did not use the adjusted data. In the later two cases Chile has a higher poverty and Mexico a lower poverty for the adjusted data.

are consistent with Sala-i-Martin (2002a,b), who used a similar estimation technique. In Figure 3-7 we show the poverty ratios for the \$2/day poverty line for China, Mexico and Thailand.

Figure 3-7: \$2/day Poverty Ratios: Selected Countries



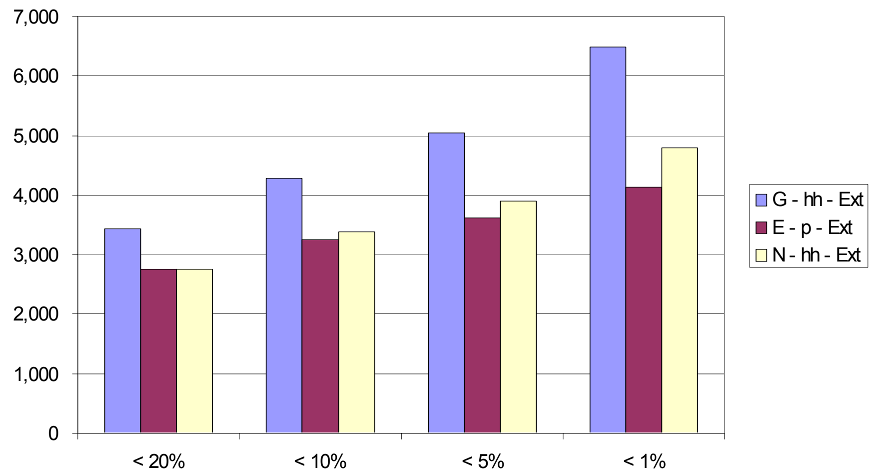
These poverty results may seen surprising, especially the lack of poverty in some Southeast Asian countries. However, one must keep in mind that the poverty lines are analytical constructs that show minimum living standards in poor countries and do not reflect any relative poverty. For some countries there can be poverty as defined in national standards, but it disappears when an international absolute poverty line is used.

Furthermore, the interrelation between growth and inequality is crucial to understand poverty reduction. To clarify this point, we have done some poverty numerics that show the relationship between income shares, GDP per capita and poverty ratios. Using equation (3.15), for any given income share we can establish the minimum GDP per capita needed to cross the poverty line. The income shares are determined by the underlying income distribution of each data point.

For illustration purposes we obtain the average income shares for the three extended series and estimate the minimum GDP per capita that assures that the poverty line is crossed.[41] Figure 3-8 shows that the required GDP per capita is higher for the gross-

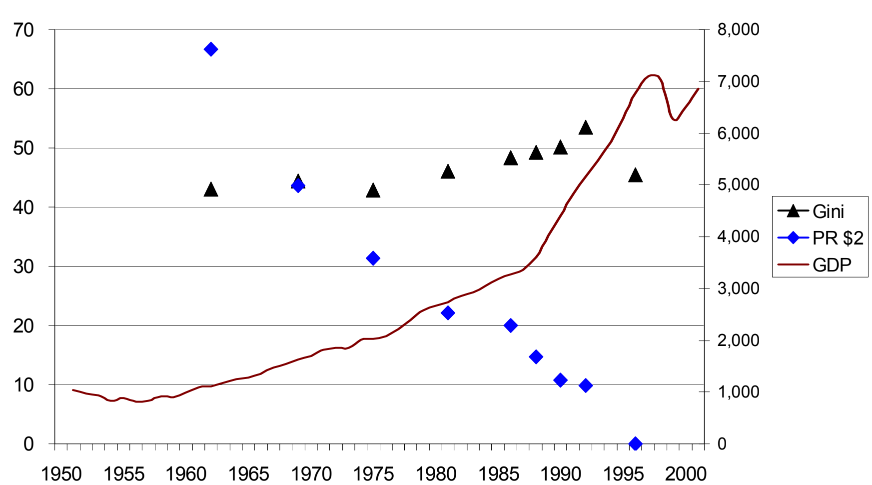
[41] We are implicitly assuming that growth does not change income inequality, although this is a controversial point (Dollar and Kraay, 2002).

Figure 3-8: Minimum GDPpc to cross $2/day poverty line



household series. This is consistent with the fact that this series has higher inequality than the net and expenditure series. Moreover, the figure shows the percentage of the population that lives below the poverty line of $2/day. For example, with a GDP per capita of at least $3000, the poverty line is below 20% when inequality is measured by expenditure or net income. Equivalently, with a GDP per capita of at least $5000 there is less than 1% of poverty. However, these minimum total income requirements can vary when the country has extreme income distributions. Additionally, for a poverty line of $1/day, the minimum GDP per capita is exactly half of the values shown in the figure.

Figure 3-9: Thailand: Poverty ratios, Gini coefficient and GDP per capita



Finally, we present the specific case of Thailand. In Figure 3-9 we plot the poverty ratio for $2/day, the GDP per capita levels and the Gini coefficient for the gross-household series. The Gini coefficient has been relatively stable over the period, with an increase in the late 1980s and beginning of the 1990s. Nevertheless, there have been high rates of GDP per capita growth in the same period and this has allowed a sharp decrease in the poverty ratio. For 1996 the poverty ratio is zero, when the GDP per capita was above $7000. The data is consistent with our poverty numerics: with a GDP of around $3000 the poverty ratio was above 20% and in 1996 the poverty ratio is zero, when the GDP per capita was above $6000.

3.6 Conclusions

The empirical study of cross-country inequality is limited by the heterogeneity and vast amount of available data. Some of these limitations can be overtaken by filtering the data with quality criteria. A complementary step is to make comparability assumptions and group data with different definitions. However, we have convincingly shown that it is not a sound practice to collapse the whole available information of a particular country into a single income inequality time series. We have proven that using different income concepts in a same series may seriously affect inequality levels and time trends. Likewise, mixing some recipient units may also alter significantly the series. There are already important measurement errors implicit in most of the inequality data and freely mixing different concepts and reference units only adds more noise to the data. Furthermore, using dummy variable adjustments does not always correct for this problem and in some cases, increases the data distortions. Thus, the single time series approach followed by Deininger and Squire (1996) must not be continued and those inequality studies that used their dataset must be reassessed.

As an alternative, we propose the use of more than one series per country, where each series is characterized by a different income concept and/or reference unit. Although an individual country may have several series, there are three series for which there is a considerable world-wide coverage: gross-household, net-household and expenditure-person. Moreover, we created two main set of series, based on the reliability of the comparability assumptions used. The "basic" series uses only strong comparability assumptions, while the "extended" series allows for less reliable assumptions, but has longer time series. In sum, our approach yields six main inequality series that can readily be used in empirical tests and within these series we have reduced the measurement error implicit in inequality data.

In any case, if the researcher is only interested in a single series to analyze world inequality, then we recommend the use of our gross-household extended series. This series offers significant country coverage and has a balanced sample of OECD and developing countries. Alternatively, if the interest is centered on OECD countries, then the net-household extended series is preferable. For developing countries the expenditure-person extended series presents a better coverage for these countries.

Additionally, we improve existing methods to estimate Lorenz curves from grouped data. Our approach is more reliable and extensive than the often used POVCAL software. The Gini estimations obtained from the income share data are highly satisfactory and statistically comparable to the coefficients reported by the primary sources. Furthermore, using our estimated Lorenz curves we can estimate Atkinson indexes, which are a conveniently complement to the information provided by the Gini coefficient. We have proven that in roughly a third of the cases both indexes report different inequality trends and thus, the use of both indexes is advisable in order to obtain robust conclusions about income inequality.

Finally, we used our dataset to test for international patterns of inequality and poverty. A first conclusion is that between-country inequality variation is more significant than within-country. This suggests that country specific characteristics have a bigger role in explaining inequality levels than time trends. However, we also find that within-country inequality is still important and there are significant time trends in our series. Therefore, we reject the "glacial change" hypothesis (Li, *et al.,* 1998) that inequality does not vary significantly over time. For the specific case of OECD countries, we clearly detect a U-shape pattern that confirms the "U-turn" hypothesis of Atkinson (2003). For developing countries the cross-country pattern is less clear, but it suggests a decrease in inequality for most of the analyzed period, with a slight increase in the 1990s. Country-specific time trends are diverse and it is difficult to spot precise trends. The choice of income concept, basic or extended series and the use of pool data may produce different results. Nevertheless, this variety of choice emphasizes the richness of our inequality dataset, which is not limited by a single series and provides wider information from where to draw conclusions. With respect to poverty, we can observe a decline in the poverty ratios in most of the countries covered by our sample. The only exception is the poverty experience in the African continent.

3.7 Appendix to Chapter 3

Table 3-17: Summary statistics, Gross-Household Basic series

Country	Obs.	Mean	St. dev.	Max	Min	Max-Min	Coverage
AUS	12	37.88	3.75	44.11	31.82	12.29	68 ~ 96
BEL	3	28.22	3.13	31.81	26.11	5.71	85 ~ 92
BGD	10	35.20	2.77	38.50	29.00	9.50	63 ~ 86
BGR	29	23.18	3.67	34.41	17.83	16.58	63 ~ 96
BHS	11	44.34	4.71	53.61	38.74	14.88	70 ~ 93
BRA	17	57.85	2.68	65.05	53.46	11.59	60 ~ 87
CAN	18	33.60	1.09	35.04	31.39	3.66	65 ~ 91
CHL	18	55.01	3.55	59.63	46.40	13.23	68 ~ 96
CHN	4	35.20	13.83	55.80	26.60	29.20	53 ~ 75
COL	7	51.69	1.90	54.52	49.24	5.28	70 ~ 92
CRI	5	47.52	3.02	51.40	44.69	6.72	61 ~ 83
DEW	8	32.07	2.16	35.56	29.40	6.16	73 ~ 84
DNK	6	36.46	4.69	41.27	30.98	10.30	78 ~ 92
ESP	5	33.19	3.24	36.30	27.77	8.53	65 ~ 91
FRA	7	43.10	6.08	49.00	34.72	14.28	56 ~ 84
GBR	4	33.87	5.44	40.38	28.40	11.97	69 ~ 95
HKG	8	44.24	3.61	52.00	39.68	12.32	71 ~ 96
IND	4	41.17	5.51	47.75	34.34	13.41	56 ~ 75
JPN	23	36.38	1.90	41.49	33.27	8.22	62 ~ 90
KOR	9	35.94	2.14	39.85	33.98	5.87	65 ~ 88
LKA	8	42.70	4.96	47.95	35.65	12.30	53 ~ 87
MEX	11	52.29	6.42	62.28	42.90	19.38	50 ~ 96
MYS	6	50.63	1.87	52.83	48.30	4.53	67 ~ 84
NLD	10	31.16	1.61	33.37	28.40	4.97	77 ~ 97
NOR	10	31.80	3.20	37.50	27.22	10.29	62 ~ 96
NZL	13	34.12	3.19	40.11	29.23	10.88	73 ~ 97
PAK	10	34.54	1.96	38.65	32.38	6.27	63 ~ 88
PER	4	50.08	3.69	55.00	46.43	8.57	71 ~ 97
PHL	11	47.61	2.64	51.45	43.61	7.83	56 ~ 97
POL	5	30.96	0.90	32.20	30.07	2.13	86 ~ 92
PRI	3	50.30	1.50	51.98	49.12	2.86	69 ~ 89
SGP	7	40.76	1.88	43.23	37.88	5.35	73 ~ 93
SWE	5	31.76	3.03	36.96	29.02	7.93	67 ~ 92
THA	11	46.88	3.57	53.53	42.90	10.63	62 ~ 96
TTO	4	46.04	4.22	51.64	41.49	10.15	58 ~ 81
TUR	3	50.81	6.12	56.26	44.20	12.06	68 ~ 87
TWN	31	31.15	1.71	34.60	28.82	5.78	64 ~ 97
USA	53	38.21	1.79	42.72	35.34	7.38	44 ~ 97
Average	10.87	40.21	3.50	45.52	35.81	9.71	
Overall	413	38.94	9.39	65.05	17.83	47.22	44 ~ 97

Between Country variation	*89%*	*Countries*	*38*	
Within Country variation	*11%*	*OECD*	*14*	*37%*

Table 3-18: Summary statistics, Net-Household Basic series

Country	*Obs.*	*Mean*	*St. dev.*	*Max*	*Min*	*Max-Min*	*Coverage*
AUS	8	37.23	6.01	44.00	31.04	12.96	68 ~ 96
BEL	4	27.08	0.96	28.39	26.11	2.29	79 ~ 92
CAN	12	30.03	1.97	34.30	26.60	7.70	71 ~ 94
DEW	7	29.98	2.35	33.56	27.40	6.16	63 ~ 83
DNK	3	30.79	1.01	31.48	29.63	1.85	76 ~ 92
FIN	10	30.43	2.35	33.93	26.37	7.55	77 ~ 91
FRA	3	30.42	2.23	31.85	27.85	4.00	79 ~ 84
GBR	32	32.79	2.57	38.38	26.23	12.15	61 ~ 95
IRL	3	38.70	1.28	39.86	37.32	2.54	73 ~ 87
ITA	25	35.19	4.58	42.00	28.78	13.22	48 ~ 95
MEX	13	50.94	5.95	58.06	40.90	17.16	50 ~ 96
NLD	16	29.17	2.14	32.40	24.66	7.74	75 ~ 94
NOR	15	28.39	2.75	34.50	24.22	10.29	62 ~ 96
POL	16	26.97	4.40	34.19	18.85	15.34	76 ~ 97
POR	4	37.09	2.51	40.36	34.25	6.11	73 ~ 91
ROM	9	26.76	3.21	31.26	22.88	8.38	89 ~ 97
SVK	11	20.96	3.91	30.60	17.73	12.87	58 ~ 97
SWE	5	28.73	2.41	32.61	26.44	6.17	67 ~ 92
TWN	30	30.11	1.66	33.60	27.82	5.78	64 ~ 97
USA	53	34.22	1.78	38.72	31.34	7.38	44 ~ 97
Average	13.95	31.80	2.80	36.20	27.82	8.38	
Overall	279	32.17	6.25	58.06	17.73	40.33	44 ~ 97

Between Country variation	*77%*	*Countries*	*20*	
Within Country variation	*23%*	*OECD*	*15*	*75%*

Table 3-19: Summary Statistics, Expenditure-Person Basic series

Country	*Obs.*	*Mean*	*St. dev.*	*Max*	*Min*	*Max-Min*	*Coverage*
BGD	3	30.22	2.98	33.64	28.23	5.41	89 ~ 96
CIV	5	38.67	1.97	41.20	36.64	4.56	85 ~ 95
ESP	3	33.41	1.37	34.90	32.18	2.71	74 ~ 91
EST	5	36.72	1.82	39.47	34.57	4.90	92 ~ 98
GHA	5	34.64	1.63	36.73	32.73	4.00	88 ~ 97
GIN	3	42.54	3.73	46.84	40.36	6.48	91 ~ 95
HUN	3	21.86	4.41	26.96	19.24	7.71	93 ~ 97
IDN	13	34.37	1.72	37.71	31.68	6.03	64 ~ 96
IND	33	32.56	2.17	37.48	29.10	8.38	51 ~ 97
IRN	5	43.23	1.41	45.45	41.88	3.57	69 ~ 84
JAM	9	41.07	2.96	45.58	36.47	9.11	71 ~ 96
JOR	4	39.38	3.80	44.21	36.33	7.87	80 ~ 97
LKA	3	34.46	4.32	38.80	30.15	8.65	87 ~ 95
MRC	3	39.27	0.24	39.53	39.09	0.44	84 ~ 99
NGA	5	43.71	5.07	50.60	36.93	13.67	86 ~ 97
PER	3	43.70	1.22	45.11	43.00	2.11	86 ~ 94
POL	3	30.10	2.75	32.66	27.20	5.47	92 ~ 96
PHL	4	42.66	2.46	46.06	40.68	5.38	85 ~ 97
THA	3	30.10	2.75	32.66	27.20	5.47	89 ~ 98
TUN	3	43.77	2.45	46.20	41.31	4.90	65 ~ 90
TZA	5	42.61	1.45	44.00	40.15	3.85	69 ~ 93
ZMB	4	45.04	9.66	59.01	38.15	20.86	91 ~ 96
Average	5.77	37.46	2.83	41.13	34.69	6.43	
Overall	127	36.97	6.00	59.01	19.24	39.77	51 ~ 99

Between Country variation	*79%*	*Countries*	*22*	
Within Country variation	*21%*	*OECD*	*1*	*5%*

Table 3-20: Summary statistics, Gross-Household Extended series

Country	Obs.	Mean	St. dev.	Max	Min	Max-Min	Coverage
AUS	15	39.14	4.22	44.22	31.82	12.40	68 ~ 98
BEL	3	28.22	3.13	31.81	26.11	5.71	85 ~ 92
BGD	10	35.20	2.77	38.50	29.00	9.50	63 ~ 86
BGR	29	23.18	3.67	34.41	17.83	16.58	63 ~ 96
BHS	11	44.34	4.71	53.61	38.74	14.88	70 ~ 93
BRA	21	58.02	2.46	65.05	53.46	11.59	60 ~ 96
BRB	3	47.76	0.79	48.27	46.85	1.42	51 ~ 79
CAN	18	33.60	1.09	35.04	31.39	3.66	65 ~ 91
CHL	18	55.01	3.55	59.63	46.40	13.23	68 ~ 96
CHN	16	32.42	7.12	55.80	24.36	31.44	53 ~ 92
COL	11	53.95	5.20	64.53	47.83	16.70	64 ~ 94
CRI	12	47.17	2.95	53.54	43.90	9.64	61 ~ 95
DEW	8	32.07	2.16	35.56	29.40	6.16	73 ~ 84
DNK	15	33.71	3.96	41.27	28.29	12.98	63 ~ 95
DOM	4	47.07	3.55	51.00	43.29	7.71	76 ~ 92
ECU	3	47.00	7.94	53.00	38.00	15.00	68 ~ 94
ESP	5	33.19	3.24	36.30	27.77	8.53	65 ~ 91
EST	6	31.97	7.23	37.75	21.00	16.75	88 ~ 96
FIN	3	35.61	10.17	47.35	29.47	17.88	62 ~ 98
FRA	7	42.98	6.97	52.09	34.72	17.36	56 ~ 84
GBR	15	30.22	3.63	40.38	27.20	13.18	64 ~ 95
GTM	4	56.10	5.18	59.56	48.40	11.16	79 ~ 89
HKG	9	44.74	3.69	52.00	39.68	12.32	65 ~ 96
HND	7	55.04	3.98	61.88	50.00	11.88	68 ~ 93
IND	5	39.54	6.01	47.75	33.00	14.75	56 ~ 75
JPN	23	36.38	1.90	41.49	33.27	8.22	62 ~ 90
KOR	9	35.94	2.14	39.85	33.98	5.87	65 ~ 88
LKA	8	42.70	4.96	47.95	35.65	12.30	53 ~ 87
MEX	11	52.29	6.42	62.28	42.90	19.38	50 ~ 96
MYS	7	50.33	1.89	52.83	48.30	4.53	67 ~ 89
NGA	3	40.77	8.87	51.00	35.18	15.82	59 ~ 82
NLD	11	32.41	4.41	44.89	28.40	16.49	62 ~ 91
NOR	11	32.25	3.37	37.50	27.22	10.29	63 ~ 91
NZL	13	34.12	3.19	40.11	29.23	10.88	73 ~ 97
PAK	11	34.77	2.00	38.65	32.38	6.27	63 ~ 88
PAN	7	54.54	4.50	58.92	47.46	11.46	70 ~ 97
PER	6	52.77	6.25	63.95	46.43	17.52	61 ~ 97
PHL	11	47.61	2.64	51.45	43.61	7.83	56 ~ 97
POL	5	30.96	0.90	32.20	30.07	2.13	86 ~ 92
PRI	4	49.15	2.61	51.98	45.68	6.30	63 ~ 89
ROM	4	28.52	1.94	31.20	27.10	4.10	89 ~ 94
RUS	7	32.16	6.58	40.01	25.90	14.11	88 ~ 98
SUN/RUS	10	30.27	6.21	40.01	24.52	15.49	80 ~ 98
SGP	7	40.76	1.88	43.23	37.88	5.35	73 ~ 93
SLV	4	51.10	2.29	53.00	48.40	4.60	65 ~ 95
SUN	4	26.00	1.24	27.54	24.52	3.02	80 ~ 89
SVK	5	22.60	1.28	24.50	21.50	3.00	89 ~ 93
SWE	6	33.14	4.34	40.06	29.02	11.04	63 ~ 92
THA	11	46.88	3.57	53.53	42.90	10.63	62 ~ 96
TTO	5	45.80	3.69	51.64	41.49	10.15	57 ~ 81
TUR	3	50.81	6.12	56.26	44.20	12.06	68 ~ 87
TWN	31	31.15	1.71	34.60	28.82	5.78	64 ~ 97
UKR	8	27.96	5.08	34.43	21.82	12.61	80 ~ 97
USA	53	38.21	1.79	42.72	35.34	7.38	44 ~ 97
VEN	12	43.88	3.26	49.63	37.68	11.95	62 ~ 97
YUF	9	33.40	1.82	37.68	31.84	5.84	63 ~ 90
ZAF	3	61.34	2.09	63.00	59.00	4.00	90 ~ 95
Average	10.09	40.34	3.86	46.02	35.25	10.77	
Overall	580	38.97	9.96	65.05	17.83	47.22	44 ~ 98

Between Country variation	*87%*	*Countries*	*57*	
Within Country variation	*13%*	*OECD*	*15*	*26%*

Table 3-21: Summary statistics, Net-Household Extended series

Country	Obs.	Mean	St. dev.	Max	Min	Max-Min	Coverage
AUS	8	37.23	6.01	44.00	31.04	12.96	68 ~ 96
BEL	4	27.08	0.96	28.39	26.11	2.29	79 ~ 92
BGR	6	34.10	2.34	37.10	30.98	6.13	92 ~ 97
CAN	12	30.03	1.97	34.30	26.60	7.70	71 ~ 94
CHN	4	35.33	6.54	43.00	28.40	14.60	78 ~ 95
CSK	10	21.73	2.46	26.99	18.49	8.51	58 ~ 88
CZE	11	22.03	3.15	27.93	18.84	9.09	58 ~ 97
CSK/CZE	19	21.89	2.85	27.93	18.49	9.45	58 ~ 97
DEW	7	29.98	2.35	33.56	27.40	6.16	63 ~ 83
DNK	6	32.07	1.59	33.79	29.63	4.16	76 ~ 95
EST	7	37.94	2.31	41.02	33.80	7.21	92 ~ 98
FIN	10	30.43	2.35	33.93	26.37	7.55	77 ~ 91
FRA	3	30.42	2.23	31.85	27.85	4.00	79 ~ 84
GBR	32	32.79	2.57	38.38	26.23	12.15	61 ~ 95
GRC	3	35.89	4.82	41.30	32.06	9.24	74 ~ 88
HUN	14	23.31	1.52	25.79	20.36	5.43	62 ~ 98
IRL	3	38.70	1.28	39.86	37.32	2.54	73 ~ 87
ITA	25	35.19	4.58	42.00	28.78	13.22	48 ~ 95
MEX	13	50.94	5.95	58.06	40.90	17.16	50 ~ 96
NLD	16	29.17	2.14	32.40	24.66	7.74	75 ~ 94
NOR	15	28.39	2.75	34.50	24.22	10.29	62 ~ 96
POL	16	26.97	4.40	34.19	18.85	15.34	76 ~ 97
POR	4	37.09	2.51	40.36	34.25	6.11	73 ~ 91
ROM	9	26.76	3.21	31.26	22.88	8.38	89 ~ 97
SVK	11	20.96	3.91	30.60	17.73	12.87	58 ~ 97
CSK/SVK	19	20.71	2.15	24.81	17.73	7.08	58 ~ 97
SWE	12	29.41	1.92	32.70	26.44	6.26	67 ~ 96
TWN	30	30.11	1.66	33.60	27.82	5.78	64 ~ 97
UKR	8	27.46	5.08	33.93	21.32	12.61	80 ~ 97
USA	53	34.22	1.78	38.72	31.34	7.38	44 ~ 97
YUG	8	33.48	6.54	45.57	27.32	18.24	90 ~ 97
YUG/YUF	9	32.21	7.22	45.57	22.00	23.57	78 ~ 97
Average	12.72	30.75	3.22	35.86	26.44	9.41	
Overall	407	30.41	6.94	58.06	17.73	40.33	44 ~ 98

Between Country variation	*79%*	*Countries*	*32*	
Within Country variation	*21%*	*OECD*	*16*	*50%*

Table 3-22: Summary statistics, Expenditure-Person Extended series

Country	Obs.	Mean	St. dev.	Max	Min	Max-Min	Coverage
BGD	6	33.46	4.37	39.19	28.23	10.96	73 ~ 96
BGR	6	25.10	2.34	28.10	21.98	6.13	92 ~ 97
CAN	6	22.10	1.15	23.60	20.70	2.90	78 ~ 92
CIV	5	38.67	1.97	41.20	36.64	4.56	85 ~ 95
DEW	3	23.25	0.41	23.68	22.88	0.80	73 ~ 83
EGY	5	36.18	5.52	42.00	28.94	13.06	59 ~ 95
ESP	10	25.34	1.41	26.98	22.59	4.39	74 ~ 96
EST	5	36.72	1.82	39.47	34.57	4.90	92 ~ 98
GHA	5	34.64	1.63	36.73	32.73	4.00	88 ~ 97
GIN	3	42.54	3.73	46.84	40.36	6.48	91 ~ 95
GRC	3	34.60	1.15	35.35	33.28	2.07	74 ~ 88
HUN	3	21.86	4.41	26.96	19.24	7.71	93 ~ 97
IDN	13	34.37	1.72	37.71	31.68	6.03	64 ~ 96
IND	33	32.56	2.17	37.48	29.10	8.38	51 ~ 97
IRN	5	43.23	1.41	45.45	41.88	3.57	69 ~ 84
JAM	9	41.07	2.96	45.58	36.47	9.11	71 ~ 96
JOR	4	39.38	3.80	44.21	36.33	7.87	80 ~ 97
LKA	6	32.34	3.92	38.80	27.38	11.42	63 ~ 95
MRC	3	39.27	0.24	39.53	39.09	0.44	84 ~ 99
NGA	5	43.71	5.07	50.60	36.93	13.67	86 ~ 97
PAK	10	31.46	0.82	32.43	29.89	2.55	69 ~ 96
PER	3	43.70	1.22	45.11	43.00	2.11	86 ~ 94
PHL	4	42.66	2.46	46.06	40.68	5.38	85 ~ 97
POL	4	28.82	3.41	32.66	24.96	7.70	86 ~ 96
SGP	4	37.55	2.99	40.95	33.70	7.25	78 ~ 93
THA	4	42.97	2.57	46.20	40.56	5.65	89 ~ 98
TUN	5	42.61	1.45	44.00	40.15	3.85	65 ~ 90
TZA	4	45.04	9.66	59.01	38.15	20.86	69 ~ 93
UGA	3	37.67	4.14	40.87	33.00	7.87	89 ~ 93
ZMB	3	46.54	3.06	49.75	43.65	6.10	91 ~ 96
Average	6.07	35.98	2.77	39.55	32.96	6.59	
Overall	182	34.78	6.89	59.01	19.24	39.77	51 ~ 99

Between Country variation	85%	*Countries*	*30*	
Within Country variation	15%	*OECD*	*4*	*13%*

Table 3-23: Fixed-effects regressions on Gini, Gross-Household Basic series

	Linear trend			Quadratic trend					
	α_1	α_2	p-value	β_1	β_2	p-value	β_3	p-value	
AUS	23.97	**0.356**	0.000	21.89	0.470	0.518	-0.001	0.874	
BEL	-9.80	0.839	0.120	302.82	-12.933	0.548	0.151	0.522	
BGD	30.44	0.152	0.180	56.32	-1.598	0.118	0.028	0.086	
BGR	16.56	**0.180**	0.002	33.69	**-0.817**	0.042	**0.014**	0.013	
BHS	58.22	**-0.347**	0.001	34.06	0.946	0.462	-0.017	0.314	
BRA	57.75	0.003	0.976	53.40	0.283	0.580	-0.004	0.580	
CAN	31.45	0.060	0.474	41.29	-0.533	0.465	0.009	0.414	
CHL	43.58	**0.269**	0.002	-6.58	**2.928**	0.000	**-0.034**	0.000	
CHN	65.73	**-1.357**	0.000	95.00	**-4.761**	0.000	**0.083**	0.000	
COL	57.06	-0.144	0.196	16.03	2.147	0.219	-0.030	0.189	
CRI	51.11	-0.115	0.462	58.71	-0.693	0.567	0.010	0.631	
DEW	29.16	0.082	0.398	34.00	-0.202	0.698	0.004	0.580	
DNK	58.77	**-0.560**	0.010	239.21	-9.306	0.051	0.104	0.066	
ESP	28.53	0.138	0.314	-9.09	**2.406**	0.015	**-0.032**	0.021	
FRA	58.77	**-0.577**	0.000	49.46	0.198	0.756	-0.014	0.218	
GBR	16.28	**0.463**	0.001	17.21	0.411	0.761	0.001	0.969	
HKG	35.70	0.220	0.057	90.89	**-2.641**	0.018	**0.036**	0.010	
IND	33.01	0.363	0.066	-6.17	**4.195**	0.001	**-0.085**	0.001	
JPN	38.45	-0.066	0.355	39.37	-0.126	0.791	0.001	0.899	
KOR	34.70	0.038	0.740	5.96	1.878	0.079	-0.028	0.084	
LKA	43.96	-0.043	0.644	63.58	**-1.731**	0.000	**0.031**	0.000	
MEX	61.03	**-0.258**	0.000	51.09	**0.575**	0.015	**-0.013**	0.000	
MYS	51.30	-0.021	0.914	-5.49	**3.587**	0.082	**-0.056**	0.080	a/
NLD	25.38	0.128	0.332	1.40	1.225	0.486	-0.012	0.532	
NOR	38.90	**-0.193**	0.022	43.66	-0.482	0.338	0.004	0.561	
NZL	18.86	**0.379**	0.001	3.63	1.136	0.286	-0.009	0.475	
PAK	37.65	-0.099	0.299	46.45	-0.670	0.493	0.009	0.557	
PER	61.09	-0.243	0.065	124.15	**-3.663**	0.062	**0.043**	0.081	a/
PHL	48.63	-0.031	0.580	50.18	-0.146	0.647	0.002	0.714	
POL	28.64	0.050	0.932	-365.52	17.268	0.488	-0.188	0.489	
PRI	54.22	-0.109	0.564	75.98	-1.383	0.497	0.018	0.530	
SGP	37.59	0.078	0.616	44.82	-0.292	0.874	0.005	0.840	
SWE	39.99	-0.220	0.106	74.55	**-2.234**	0.052	**0.028**	0.078	a/
THA	39.67	**0.196**	0.009	33.23	0.587	0.286	-0.005	0.474	
TTO	48.93	-0.101	0.516	12.34	**3.103**	0.004	**-0.062**	0.003	
TUR	71.20	**-0.618**	0.001	89.36	-1.715	0.559	0.016	0.708	
TWN	30.19	0.025	0.635	49.33	**-1.037**	0.007	**0.014**	0.006	
USA	36.84	**0.049**	0.041	40.29	**-0.307**	0.000	**0.006**	0.000	
	R-square						**0.96**		
	F-statistic						**87.55**		

Note: a/ We reject the null hypothesis that both coefficients are zero at a 95% confidence level.

Table 3-24: Fixed-effects regressions on Gini, Gross-Household Extended series

	Linear trend				Quadratic trend							
	α_1	α_2	SE	p-value	β_1	β_2	SE	p-value	β_3	SE	p-value	
AUS	**23.30**	0.381	0.084	0.000	22.72	0.411	0.576	0.477	0.000	0.007	0.959	
BEL	-9.80	0.839	0.613	0.172	302.82	-12.933	21.865	0.555	0.151	0.240	0.529	
BGD	30.44	0.152	0.129	0.239	56.32	-1.598	1.037	0.124	0.028	0.016	0.091	
BGR	**16.56**	0.180	0.064	0.005	33.69	**-0.817**	0.408	0.046	**0.014**	0.006	0.014	
BHS	**58.22**	-0.347	0.115	0.003	34.06	0.946	1.308	0.470	-0.017	0.017	0.322	
BRA	56.87	0.029	0.077	0.710	57.52	-0.009	0.375	0.980	0.001	0.005	0.918	
CAN	31.45	0.060	0.096	0.530	41.29	-0.533	0.741	0.472	0.009	0.011	0.422	
CHL	43.58	0.269	0.099	0.007	-6.58	**2.928**	0.672	0.000	**-0.034**	0.008	0.000	
CHN	**42.93**	-0.275	0.073	0.000	82.95	**-3.378**	0.312	0.000	**0.050**	0.005	0.000	
COL	56.07	-0.056	0.087	0.517	122.65	**-3.918**	0.757	0.000	**0.052**	0.010	0.000	
CRI	53.49	-0.165	0.088	0.061	55.89	-0.311	0.481	0.519	0.002	0.007	0.760	
DEW	29.16	0.082	0.110	0.458	34.00	-0.202	0.529	0.703	0.004	0.007	0.587	
DNK	**44.44**	-0.253	0.097	0.010	11.97	**1.597**	0.512	0.002	**-0.025**	0.007	0.000	
DOM	32.21	0.352	0.251	0.162	94.29	-2.766	3.530	0.434	0.038	0.043	0.377	
ECU	**24.65**	0.532	0.146	0.000	147.15	-6.789	9.820	0.490	0.097	0.130	0.456	
ESP	28.53	0.138	0.156	0.377	-9.09	**2.406**	1.003	0.017	**-0.032**	0.014	0.023	
EST	**-64.50**	1.942	0.442	0.000	-478.19	18.907	15.869	0.234	-0.173	0.162	0.286	
FIN	**45.03**	-0.281	0.068	0.000	83.28	**-2.527**	0.420	0.000	**0.030**	0.006	0.000	
FRA	**60.48**	-0.621	0.126	0.000	49.34	0.307	0.646	0.635	-0.017	0.012	0.147	
GBR	**17.76**	0.419	0.099	0.000	26.33	-0.100	0.571	0.861	0.007	0.008	0.359	
HKG	42.31	0.066	0.109	0.545	86.00	**-2.399**	0.669	0.000	**0.033**	0.009	0.000	
HND	**71.37**	-0.384	0.143	0.008	87.04	-1.326	2.404	0.582	0.013	0.033	0.695	
IND	**29.03**	0.505	0.195	0.010	-10.87	**4.523**	1.158	0.000	**-0.091**	0.026	0.001	
JPN	38.45	-0.066	0.082	0.417	39.37	-0.126	0.484	0.794	0.001	0.007	0.901	
KOR	34.70	0.038	0.129	0.770	5.96	1.878	1.083	0.084	-0.028	0.016	0.089	
LKA	43.96	-0.043	0.107	0.685	63.58	**-1.731**	0.405	0.000	**0.031**	0.007	0.000	
MEX	**61.03**	-0.258	0.061	0.000	51.09	**0.575**	0.239	0.016	**-0.013**	0.004	0.000	
MYS	53.10	-0.082	0.160	0.609	21.58	1.801	1.320	0.173	-0.027	0.019	0.153	
NGA	**61.83**	-0.679	0.166	0.000	106.89	-4.661	8.039	0.562	0.073	0.147	0.621	
NLD	**43.75**	-0.265	0.095	0.005	74.78	**-2.073**	0.459	0.000	**0.024**	0.006	0.000	
NOR	**39.76**	-0.213	0.085	0.013	45.20	-0.557	0.466	0.233	0.005	0.007	0.456	
NZL	**18.86**	0.379	0.130	0.004	3.63	1.136	1.083	0.295	-0.009	0.013	0.483	
PAK	38.24	-0.114	0.102	0.264	48.40	-0.785	0.915	0.392	0.010	0.014	0.462	
PAN	49.02	0.129	0.127	0.310	116.42	**-3.341**	1.046	0.002	**0.042**	0.013	0.001	
PER	**68.50**	-0.398	0.096	0.000	87.21	**-1.586**	0.577	0.006	**0.016**	0.008	0.039	
PHL	48.63	-0.031	0.063	0.627	50.18	-0.146	0.325	0.653	0.002	0.005	0.719	
POL	28.64	0.050	0.661	0.940	-365.52	17.268	25.311	0.496	-0.188	0.276	0.497	
PRI	46.40	0.086	0.154	0.577	29.63	1.200	1.092	0.273	-0.017	0.016	0.305	
ROM	52.21	-0.488	0.844	0.563	1160.15	-46.240	37.861	0.223	0.472	0.390	0.228	
SGP	**-0.31**	0.612	0.174	0.001	100.66	**-3.791**	2.321	0.103	**0.047**	0.025	0.058	a/
SLV	37.59	0.078	0.178	0.659	44.82	-0.292	1.866	0.876	0.005	0.023	0.843	
SVK	14.44	0.170	0.962	0.860	-889.92	37.884	60.082	0.529	-0.393	0.626	0.531	
SWE	**44.23**	-0.321	0.120	0.008	71.53	**-2.073**	0.802	0.010	0.026	**0.012**	0.029	
THA	**39.67**	0.196	0.085	0.021	33.23	0.587	0.559	0.294	-0.005	0.008	0.482	
TTO	46.70	-0.035	0.142	0.804	14.78	**2.940**	0.985	0.003	-0.059	**0.019**	0.003	
TUR	**71.20**	-0.618	0.218	0.005	89.36	-1.715	2.981	0.565	0.016	0.042	0.713	
TWN	30.19	0.025	0.059	0.677	49.33	**-1.037**	0.391	0.008	0.014	**0.005**	0.007	
UKR	45.79	-0.388	0.229	0.092	273.83	**-10.567**	2.579	0.000	0.112	**0.028**	0.000	
USA	36.84	0.049	0.027	0.072	40.29	**-0.307**	0.087	0.000	0.006	**0.002**	0.000	
VEN	38.15	0.147	0.090	0.103	59.10	**-1.046**	0.466	0.025	0.016	**0.006**	0.010	
YUF	41.62	-0.207	0.132	0.118	45.51	-0.467	0.902	0.605	0.004	0.014	0.772	
	R-square		**0.92**				**R-square**		**0.96**			
	F-statistic		**106.79**				**F-statistic**		**92.96**			

Note: a/ We reject the null hypothesis that both coefficients are zero at a 95% confidence level.

Table 3-25: Time trends, Basic series, Gross-Household and panel data of all definitions

	Gross-Household		**Fixed-effects**				**Gross-Household**		**Fixed-effects**		
	Linear	**Quad.**	**Linear**	**Quad.**	**Series**		**Linear**	**Quad.**	**Linear**	**Quad.**	**Series**
AUS	0.36		0.30		5	**JPN**				U	1
BEL					n.a.	**KOR**					1
BGD					3	**LKA**		U			3
BGR	0.18	U	0.21		2	**MEX**	-0.26	inv U	-0.19		3
BHS	-0.35		-0.30		1	**MYS**		inv U		inv U	1
BRA			0.11		2	NGA	n.a.	n.a.	1.19		1
CAN					4	**NLD**			0.17		4
CHL	0.27	inv U	0.28	inv U	2	**NOR**	-0.19		-0.22		4
CHN	-1.36	U		U	3	**NZL**	0.38		0.37		1
CIV	n.a.	n.a.			1	**PAK**					2
COL					2	PAN	n.a.	n.a.			1
CRI			-0.18		2	**PER**		U		U	2
CSK	n.a.	n.a.	-0.19		1	**PHL**					3
CZE	n.a.	n.a.		U	1	**POL**			0.40	U	3
DEW				U	7	**PRI**					n.a.
DNK	-0.56			inv U	6	ROM	n.a.	n.a.	1.86		1
EGY	n.a.	n.a.	-0.42		1	RUS	n.a.	n.a.			1
ESP		inv U	-0.15		4	**SGP**					2
EST	n.a.	n.a.			3	SVK	n.a.	n.a.	-0.24	U	1
FIN	n.a.	n.a.	-0.18	U	3	**SWE**		U		U	4
FRA	-0.58		-0.49	U	3	**THA**	0.20		0.19		3
GBR	0.46		0.23	U	4	TUN	n.a.	n.a.			1
GHA	n.a.	n.a.	-0.41		1	**TTO**		inv U			n.a.
HKG		U	0.40		1	**TUR**	-0.62				n.a.
HND	n.a.	n.a.			1	**TWN**		U		U	2
HUN	n.a.	n.a.		U	2	UKR	n.a.	n.a.		U	2
IND		inv U	-0.68		2	**USA**	0.05	U	0.50	U	4
IRN	n.a.	n.a.			1	VEN	n.a.	n.a.			1
ITA	n.a.	n.a.	-0.21		2	YUF	n.a.	n.a.			1
JAM	n.a.	n.a.		inv U	1	YUG	n.a.	n.a.			1

Notes: n.a. stands for countries without a series with more than five observations.
"n.a." is for countries without a Gross-HH basic series.
"U" refers to significant negative and positive quadratic coefficients for β2 and β3 respectively.
"inv U" is assigned when β2 and β3 are significant, while positive and negative respectively.

Chapter 4

Household Inequality, Welfare, and the Setting of Trade Policy[1]

4.1 Introduction

There has been widespread interest in the policy community in the distributional effects of trade liberalization. Recent anti-globalization rhetoric has focused on the potential negative impact of trade and outsourcing for unskilled wages in developed countries, while arguing that it may also hurt workers in poor countries. On the other hand, many economists regard trade liberalization as an instrument for increasing growth, but are less certain about the distributional effects of increased openness.

Driven in large part by continued policy interest, the relationship between trade openness and wages has also been an active topic of debate in the research community.[2] However, this literature focuses on the functional distribution of income, with less emphasis on household distribution issues.[3] The current literature has stressed theoretical linkages based on the Stolper-Samuelson theorem. In this context, tariff reductions in poor labor-abundant countries are expected to increase the real income of workers and hurt capital

[1] **This chapter is based on "Household Inequality, Welfare, and the Setting of Trade Policy" by Francois and Rojas-Romagosa (2004b).**

[2] Comprehensive surveys are provided by Richardson (1995), Cline (1997) and Feenstra and Hanson (2004).

[3] Recent papers include Edwards (1997), Higgins and Williamson (1999), Barro (2000) and Spilimbergo *et al.* (2001).

owners (or skilled labor). In developed countries the opposite effect is expected. The empirical evidence remains mixed and somewhat contradictory.

In this paper, we follow Bourguignon and Morrisson (1989, 1990) and Spilimbergo *et al.* (1999) and use an ownership matrix that allows us to move from functional to household income. We then obtain a general function of the personal income distribution in terms of endowments, tariffs and the ownership structure. Using this analytical framework, we analyze the impact of trade and tariffs on the household distribution of income in general equilibrium.

Treating equity issues as relevant, we work with Sen-type social welfare functions (Sen, 1974) and embed inequality indexes in social welfare indexes. In particular, we work with the widely used Gini coefficient and with the Atkinson family of inequality indexes (Atkinson, 1970), although other indexes may be employed. Using this framework we are able to decompose the general equilibrium import protection effects into real income level and dispersion changes.

The efficiency properties of tariffs are well developed in the literature. What we highlight here is how distributional aspects of social welfare related to import protection may be examined alongside efficiency aspects. For government, this is manifested not only in special interest politics, but also through the direct impact of inequality on a government's objective function. We find that equity considerations may serve to counter lobbying interests in both capital-rich and capital-poor countries, though with opposite marginal impact on the final policy outcome. Although we focus our attention only on import tariffs, the main message that follows from this approach can be applied in a more general context of trade policy instruments. The precise distributional and efficiency components may change, but in essence the trade-off and interrelation between both economic outcomes is still present.

The paper is organized as follows. Section 2 develops a formal representation of social welfare inclusive of income inequality. In Section 3, we embed this social welfare function into a general equilibrium trade model. We also develop the equilibrium representation of inequality, based on the general equilibrium system fundamentals. Section 4 then explores linkages between trade policy, inequality, and welfare. It also examines theoretical linkages between country size, development, policy, and inequality. In Section 5, we explore the implications of the addition of inequality to the social welfare component of a government's objective function for political support function models of tariff formation. We conclude in Section 6.

4.2 Defining social welfare with respect to inequality

Our goal in this section is to develop a functional linkage between inequality and aggregate (social) welfare. This will then be integrated in the next section into a general equilibrium trade model. A critical condition for inequality to have a meaningful link to aggregate (social) welfare is that the utility function be strictly concave with respect to income. Additionally, for tractability we prefer to work with a social welfare function that is additively separable and symmetric in individual incomes.[4]

The existence of social welfare functions depends crucially on the possibility to compare interpersonal utility levels. One such possibility is offered by the "veil of ignorance" approach first proposed by Harsanyi (1953, 1955) and further developed by Rawls (1971), where we rank different individual situations not knowing which would be the actual situation. As stated by Sen (1997) this interpersonal comparison can be defined as those situations where we make judgements of the type:

> "I would prefer to be person A rather than person B in this situation" and "while we do not really have the opportunity (or perhaps the misfortune, as the case may be) of in fact becoming A or B, we can think quite systematically about such a choice, and indeed we seem to make such comparisons frequently".

Because GDP per capita is the most common indicator of social welfare, the "veil of ignorance" approach supports the use of an inequality measure to complement GDP per capita comparisons. If we do not know which individual household we are in a specific country, then the expected utility becomes a function of mean income and the

[4]Note that the assumption of symmetric individual incomes implies that each individual is treated anonymously. Thus, it is the individual's income which provides the concavity of the function, not his identity.

personal distribution of income. How we evaluate the probability of receiving any given income is then determined by the functional representation of the utility function and more specifically by the degree of concavity of this function. In this context, a natural extension of cross-country welfare comparisons is to complement GDP per capita levels with some measure of inequality.[5]

Under the social welfare approach to income distribution measurement, inequality is associated with variance in the distribution of income. This raises two measurement problems. The first is that we cannot generally rely on first moment-based indicators. The second is that even though the concepts of Lorenz-dominance and General Lorenz-dominance (Shorrocks, 1983) are accepted as ways to impartially rank two different distributions[6], in many cases the Lorenz-curves intersect at least once, so that we obtain incomplete ranking of distributions. To solve both these problems, inequality indexes are usually used to rank distributions in indeterminate cases and to provide a summary variable that can be used in empirical models. While the most commonly used is the Gini coefficient, most inequality measures are implicitly based on a social welfare function (Dalton, 1920; Kolm, 1969; Atkinson,1970). As such, there is no perfect index, and any index has built in social preferences.

In this paper, we employ two representations of household utility and social welfare. Both reflect Sen's (1974) preferred definition of social welfare as:[7]

$$SW = \overline{y}\,(1 - I) \tag{4.1}$$

where SW is the social welfare, $\overline{y}$ is mean income, and I is an index of inequality.

Starting with CRRA preferences yields the well-known Atkinson inequality index directly as a natural metric for a mapping from income distribution to social welfare (see Atkinson 1970). In this sense, Atkinson's index fits naturally into Sen's proposed social welfare function.

Sen actually offered equation (4.1) as defined with respect to the Gini coefficient. In this case, the social welfare function is axiomatic, in that we do not have an obvious

[5] This approach was formally treated by Sen (1976).

[6] See Lambert (1993) for details.

[7] A summary with the definition of the variables is presented in the Appendix.

mapping –through aggregation– from individual preferences to an aggregate social welfare function. This follows because the social welfare function is then rank sensitive. We work with both the Atkinson index and Gini coefficient in this paper.

4.2.1 The Atkinson index-based social welfare function

Formally, we define a composite consumer good over the range of all consumption goods, which follows from a linear homothetic aggregation function. As such, cost minimization yields a composite consumer price index. This is defined over all consumer prices p_c.

$$p_c = f(p) \tag{4.2}$$

Household utility u^h is defined as a function of household consumption of the composite consumer good c^h:

$$u^h = \psi\left(c^h\right) \tag{4.3}$$

We next map aggregate individual utility to aggregate welfare ϕ, which is defined as the sum of household utility,

$$\phi = \sum_h u^h \tag{4.4}$$

while aggregate consumption c is the sum of household consumption.

$$c = \sum_h c^h \tag{4.5}$$

We will assume that the function ψ is CRRA:[8]

$$\psi\left(c^h\right) = \begin{cases} \frac{\left(c^h\right)^{1-\theta}}{1-\theta} & \text{if } \theta \neq 1 \\ \ln c^h & \text{if } \theta = 1 \end{cases} \tag{4.6}$$

[8] In the present context, constant relative inequality aversion (CRIA) is a better label and acronym.

In general, we assume that $\theta > 0$, and in this paper we focus on the case where $\theta \neq 1$.[9] We employ a simple linear transformation, and are then able to define a social welfare index in per-capita terms.

$$SW_A = \frac{\phi}{n\left(1-\phi\right)} = \frac{1}{n}\sum_h \left(c^h\right)^{1-\theta} \tag{4.7}$$

Simple manipulation then yields social welfare as a function of per-capita income $\overline{y}$, consumer prices, and income equality.

$$SW_A = \left(\frac{\overline{y}}{p_c}\right)^{1-\theta} E_A^{1-\theta} \tag{4.8}$$

With some further manipulation, our equality measure E_A can be mapped directly to the Atkinson index of income inequality, yielding a Sen-type social welfare function. In particular, taking the definition of the Atkinson index (Atkinson, 1970), we have the following relationships between the Atkinson index I_A, E_A, and social welfare.

$$I_A = 1 - \left(\frac{1}{n}\sum_h \left(\frac{y^h}{\overline{y}}\right)^{1-\theta}\right)^{\frac{1}{1-\theta}} = 1 - E_A \tag{4.9}$$

$$SW_A = \left[\left(\frac{\overline{y}}{p_c}\right)\left(1 - I_A\right)\right]^{1-\theta} \tag{4.10}$$

Note that as $\theta \rightarrow 0$ only average income matters, rather than income inequality. Alternatively, when $\theta \rightarrow \infty$, then $SW_A = \min\left(y^h\right)$ and we have the extreme Rawlsian maximin social welfare function, were the income of the poorest individual is the only relevant variable and average income is unimportant. Moreover, for a given distribution (measured as shares of total income) we have declining marginal utility of income.

4.2.2 The Gini index-based social welfare function

The Gini coefficient is defined as twice the area between the Lorenz curve and the 45-degree line. As such, $(1 - G)$ is then twice the area below the Lorenz curve. Formally,

[9] One gets the same basic results with log preferences. Estimates in the macro literature are that θ is less than 1.

this index is defined as follows:

$$I_G = 1 + \frac{1}{n} - \frac{2}{n^2\overline{y}}\left(y^1 + 2y^2 + ... + ny^n\right) = 1 + \frac{1}{n} - \left[\left(\frac{2}{n^2}\sum_h \frac{hy^h}{\overline{y}}\right)\right] \tag{4.11}$$

$$SW_G = \left[\left(\frac{\overline{y}}{p_c}\right)(1 - I_G)\right] \tag{4.12}$$

where we have arranged households so that $y^1 \geq y^2 \geq ... \geq y^n$. Unlike the Atkinson-based social welfare function, the Gini-based social welfare function embodies asymmetry not on specific individuals, but rather on relative income rankings. This ranking provides the concavity of the utility function with respect to income. The higher the income in the ranking, the less social weight it has. At the same time, equation (4.12) is linear in average income. As such, SW_G is relatively more sensitive to mean income than SW_A and less sensitive to inequality.

4.3 Inequality and trade in general equilibrium

To explore the interaction between production, trade and trade policy, and inequality, we work with a modified dual representation of trade in general equilibrium (Dixit and Norman, 1980). To do so, we first adopt the following additional set of assumptions:

- Rational behavior by households and firms.
- Complete and perfectly competitive markets.
- Convex technology, with neoclassical production functions.
- Goods are tradable and factors are not.
- Every household has the same neoclassical technology for producing the composite consumption good.

Given these assumptions, we are able to define the core general equilibrium system for demand and production in terms of expenditure and revenue functions, with expenditure defined in terms of the composite consumption good. Social welfare then follows as a set of side equations from the core general equilibrium system.

4.3.1 The core general equilibrium system

Because we assume that all households have the same consumption technology defined with respect to the composite consumption good, we can drop the household index from consumption and represent aggregate expenditure as a function of aggregate consumption and prices:

$$e(p, c) = c \cdot f(p) \tag{4.13}$$

On the production side, we assume standard neoclassical production functions with constant returns to scale: $x_i = g_i(v_{ij})$, where $g_i(\cdot)$ is the production function for good i and v_{ji} is the use of factor j in the production of good i. Defining unit input coefficients as a_{ji} we also obtain: $1 \leq g_i(a_{ji})$. Endowment constraints are then $\sum a_{ji}x_i \leq \overline{v}_j$. From these conditions, we can define the economy-wide revenue function with respect to goods prices and endowments. This is represented in equation (4.14).

$$r(p, v) = \max_{x_i, a_{ji}} \left\{ \sum_i p_i x_i \mid \sum_i a_{ji} x_i \leq \overline{v}_j \text{ and } 1 \leq g_i(a_{ji}) \;\; \forall i, j \right\} \tag{4.14}$$

From the envelope theorem and the properties of the revenue function r, factor incomes and goods production can be expressed in terms of the value of the partial derivatives of the revenue function, evaluated at the equilibrium set of prices:

$$\frac{\partial r(p, v)}{\partial v_j} = w_j = w_j(p, v) \qquad \forall j \tag{4.15}$$

$$\frac{\partial r(p, v)}{\partial p_i} = x_i = x_i(p, v) \qquad \forall i \tag{4.16}$$

Taking equations (4.15) and (4.16) in conjunction with equations (4.13) and (4.14), we can write the general equilibrium system for production, consumption, and trade as follows:[10]

$$
\begin{aligned}
c^h f(p) &= \left(\sum_j w_j(p,v)\cdot v_j^h\right) + \omega_\tau^h \tau m \qquad \forall h && (4.17)\\
m &= \sum_h c^h \cdot f(p) - x(p,v) && (4.18)\\
e(p,c) &= \sum_h \left[\left(\sum_j w_j(p,v)\cdot v_j^h\right) + \omega_\tau^h \tau m\right] && (4.19)\\
p &= P^* + \tau = 1+\tau && (4.20)
\end{aligned}
$$

In equations (4.17) – (4.20), we have assumed the home country imposes a tariff of τ on imports from the rest of the world, while world prices are normalized to one. In addition, ω_t^h is the household share of the tariff revenue and v_j^h is the household ownership share of factor j. In the first equation, household consumption is equal to the household budget. Equation (4.18) defines imports on which tariff revenue is generated and equation (4.19) sets economy wide expenditure equal to national income. Together, the system of four equations has an equally dimensioned set of unknowns: c^h, m, e and p.

4.3.2 Household inequality

As explained earlier, the recent literature on trade and the distribution of income has focused on the functional distribution of income. The functional distribution of income is also an important building block here for the representation of the household distribution of income. In equation (4.21) we define factor incomes s, which follow directly from the endowment stock and the properties of the revenue function.

$$s_j = r_{vj}(p,v)\cdot v_j = w_j v_j \qquad (4.21)$$

[10] A two-country general equilibrium system can readily be formalized using the same framework.

Thus, the functional distribution of income is a function of equilibrium prices, preferences, the production technology and the endowment set. In reduced form, the functional distribution of income $F(s)$ is then an artefact of the equilibrium matching of preference and the technology set, given our endowment vector.

$$F(s) = F(p, v) \tag{4.22}$$

Using factor incomes w_j and the household ownership share of production factors, ω_j^h we can readily obtain household income. In addition, we include the assignment of import tariff revenue, again represented by a household share parameter. Equation (4.23) presents the basic definition of household income in terms of its primary components.

$$y^h = \left(\sum_j w_j \cdot v_j \cdot \omega_j^h \right) + \omega_\tau^h \tau m \tag{4.23}$$

$$c^h = \frac{y^h}{p_c} \tag{4.24}$$

where $1 \geq \omega^h \geq 0$ and $\sum \omega_j^h = \sum \omega_\tau^h = 1$. In reduced form, the personal distribution of income $F(y)$ is a consequence of the elements affecting the functional distribution and the $h \times j$ ownership matrix of coefficients ω_j^h, represented by S:

$$F(y) = F(p, v, S) \tag{4.25}$$

Note that social welfare is ultimately a function of the ownership matrix in the economy, while the impact of trade policy will then depend on the interaction of the underlying economic structure and the ownership matrix.

4.3.3 Inequality indexes with system fundamentals

We can write our social metrics of the distribution of income –the Atkinson and Gini indexes– in terms of system fundamentals. Making a substitution from (4.23) into (4.9) and (4.11), we obtain the following equations:

$$
\begin{aligned}
I_A &= 1-\left\{\frac{1}{n}\sum_h\left[\frac{n\left(\sum_j w_j v_j \omega_j^h\right)+n\omega_\tau^h \tau m}{y}\right]^{1-\theta}\right\}^{\frac{1}{1-\theta}} \\
I_A &= 1-\left\{n^{-\theta}\sum_h\left[n^{-1}+\sum_j \beta_j\left(\omega_j^h-n^{-1}\right)\right]^{1-\theta}\right\}^{\frac{1}{1-\theta}} \qquad (4.26)
\end{aligned}
$$

$$
\begin{aligned}
I_G &= 1+\frac{1}{n}-\frac{2}{n^2}\sum_h h\left[\frac{n\left(\sum_j w_j v_j \omega_j^h\right)+n\omega_\tau^h \tau m}{y}\right] \\
I_G &= 1+\frac{1}{n}-\frac{2}{n}\sum_h h\left[n^{-1}+\sum_j \beta_j\left(\omega_j^h-n^{-1}\right)\right] \qquad (4.27)
\end{aligned}
$$

where $\beta_j = \frac{w_j v_j}{y}$ represents the national income share of factor and $\sum_j \beta_j + \frac{\tau m}{y} = 1$. In what follows, we apply the additional normalization that each household receives an equal share of the tariff revenues, so that $\omega_\tau^h = n^{-1}$.[11]

The ratio of the household's income to per capita income, which accounts for income dispersion, is given by the sum of the differences between the actual ownership share of factors and equal shares for each household. From equations (4.26) and (4.27), we can make a substitution back into equations (4.10) and (4.12), yielding social welfare itself as a function of system fundamentals.

$$
\begin{aligned}
SW_A &= \left[\frac{\overline{y}}{p_c}\left(1-I_A\right)\right]^{1-\theta} \\
SW_A &= \left(\frac{\overline{y}}{p_c}\right)^{1-\theta} n^{-\theta}\sum_h\left[n^{-1}+\sum_j \beta_j\left(\omega_j^h-n^{-1}\right)\right]^{1-\theta} \qquad (4.28)
\end{aligned}
$$

[11] The distributional impact of tariff revenues can be substantial. This is the emphasis of the paper by Galor (1994), which includes tariffs in his general equilibrium Overlapping-Generations model.

$$SW_G = \frac{\overline{y}}{p_c}(1-I_G)$$
$$SW_G = \left(\frac{\overline{y}}{p_c}\right)\left\{\frac{2}{n}\sum_h h\left[n^{-1}+\sum_j \beta_j\left(\omega_j^h - n^{-1}\right)\right]-n^{-1}\right\} \tag{4.29}$$

4.4 Trade policy, distribution, and welfare

4.4.1 Generalized effects

From equations (4.28) and (4.29) above, social welfare is a function of the first two moments of the household distribution of income. This is especially obvious with the Atkinson index, as it is actually the weighted variance of income, with inverse income weights, that provides the variance component of the social welfare function.[12] Because the contributions of the mean and variance components of income to social welfare are separable in equations (4.28) and (4.29), we can decompose the impact of trade policy as well into its impact on per-capita income (an efficiency effect), and its impact on the variance of income (a distributional effect). Together, they determine the overall social welfare impact. Formally, differentiating equations (4.28) and (4.29) with respect to tariffs, we obtain the following equations:

$$\begin{aligned}\frac{\partial SW_A}{\partial \tau} &= (1-\theta)\left[\frac{\overline{y}}{p_c}(1-I_A)\right]^{-\theta}(1-I_A)^{1-\theta}\left(\frac{\partial \overline{y}}{\partial \tau}-\frac{\partial p_c}{\partial \tau}\frac{\overline{y}}{p_c^2}\right)\\ &\quad -(1-\theta)\left(\frac{\overline{y}}{p_c}\right)^{1-\theta} I_A^{-\theta}\frac{\partial I_A}{\partial \tau}\end{aligned} \tag{4.30}$$
$$\frac{\partial SW_G}{\partial \tau} = (1-I_G)\left[\frac{\partial \overline{y}}{\partial \tau}-\frac{\partial p_c}{\partial \tau}\frac{\overline{y}}{p_c^2}\right]-\left(\frac{\overline{y}}{p_c}\right)\frac{\partial I_G}{\partial \tau} \tag{4.31}$$

How do we interpret equations (4.30) and (4.31)? The efficiency component is well known (see for example Dixit and Norman, 1980), and is shown here in equation (4.32). Basically, the impact of the tariff on per-capita income will depend on the combination of

[12] While the functional form is different, the social welfare function underlying other income distribution indexes yields a similar result, though with different weights in the variance component of the welfare function. The CRRA function yields a particularly clear and parsimonious reduced form.

terms-of-trade and allocation effects (the first set of terms in square brackets in equation (4.32)), and tariff revenue (the second set of terms).

$$\frac{\partial \overline{y}}{\partial \tau} = \frac{1}{n} \sum_h \frac{\partial y^h}{\partial \tau} = \frac{1}{n} \left[m \left(1 - \frac{\partial p}{\partial \tau} \right) + \tau \frac{\partial m}{\partial \tau} \right] \tag{4.32}$$

For a small country, negative allocation effects outweigh the terms-of-trade effects, so that the impact of the tariff on mean income is strictly negative. Also, for the small country, the impact on the cost of living will be to raise prices. As such, the real mean-income effect will be strictly negative for a small country. With a large country, the combined income and cost-of-living effect, or in other words the real income effect of the tariff change as represented by the term in square brackets in the equations (4.30) and (4.31) may be positive or negative depending on the magnitude of terms-of-trade effects.

The impact on household income distribution, the other part of equations (4.30) and (4.31), follows from differentiation of equations (4.26) and (4.27). This is shown below:

$$\begin{aligned} \frac{\partial I_A}{\partial \tau} &= -n^{\frac{-\theta}{1-\theta}} \left\{ \sum_h \left[\sum_j \beta_j \left(\omega_j^h - n^{-1} \right) + n^{-1} \right]^{1-\theta} \right\}^{\frac{\theta}{1-\theta}} \\ &\quad \sum_h \left\{ \left[\sum_j \beta_j \left(\omega_j^h - n^{-1} \right) + n^{-1} \right]^{-\theta} \left[\sum_j \frac{\partial \beta_j}{\partial \tau} \left(\omega_j^h - n^{-1} \right) \right] \right\} && (4.33) \\ \frac{\partial I_G}{\partial \tau} &= -\frac{2}{n} \sum_h h \left[\sum_j \frac{\partial \beta_j}{\partial \tau} \left(\omega_j^h - n^{-1} \right) \right] && (4.34) \end{aligned}$$

Note that we also have an inverse income weighting, by a factor of θ, in equation (4.33) applied to induced changes in income. The weighting of induced changes in income for the Gini index depends on the ranking of individual households on the relative income scale. In both cases, the changes in income in turn depend on Stolper-Samuelson derivatives, and induced price changes that follow from tariff changes. This is expressed in equation (4.35), where the term $\frac{\partial \beta_j}{\partial \tau}$ depends on system fundamentals and Stolper-Samuelson relationships.

$$\frac{\partial \beta_j}{\partial \tau} = \frac{\partial w_j}{\partial p} \frac{\partial p}{\partial \tau} \frac{v_j}{y} - \frac{\partial y}{\partial \tau} \frac{w_j v_j}{y^2} \tag{4.35}$$

We can also represent the relationship in elasticity terms: $\varepsilon_{\beta_j,\tau} = \varepsilon_{w_j,p}\varepsilon_{p,\tau} - \varepsilon_{y,\tau}$.

4.4.2 Heckscher-Ohlin inequality effects

Assuming that all households hold a claim on one unit of labor, we can apply equations (4.33) and (4.34) to a two-factor, two-good Heckscher-Ohlin model. In this framework, equations (4.36) and (4.37) determine the impact of tariff changes on household inequality.

$$
\begin{aligned}
\frac{\partial I_A}{\partial \tau} &= -n^{\frac{-\theta}{1-\theta}} \left\{ \sum_h \left[\beta_k \left(\omega_k^h - n^{-1}\right) + n^{-1}\right]^{1-\theta} \right\}^{\frac{\theta}{1-\theta}} \\
&\quad \sum_h \left\{ \left[\beta_k \left(\omega_k^h - n^{-1}\right) + n^{-1}\right]^{-\theta} \left[\frac{\partial \beta_k}{\partial \tau} \left(\omega_k^h - n^{-1}\right)\right] \right\} \qquad (4.36) \\
\frac{\partial I_G}{\partial \tau} &= -\frac{2}{n} \sum_h h \left[\frac{\partial \beta_k}{\partial \tau} \left(\omega_k^h - n^{-1}\right)\right] \qquad (4.37)
\end{aligned}
$$

Inequality is purely a function of the allocation of capital in this model. At the same time, the impact of the tariff is then a function of which sector is protected. If protection leads to an increase in wages and a drop in capital income, inequality is reduced. Alternatively, if capital income is protected, we will see a rise in inequality.

The social welfare effect, however, will depend on the trade-off between real income effects following from import protection, and the impact on inequality. In other words, the trade-off between equity and efficiency. From equations (4.30) and (4.31), this is ultimately a function of the degree of inequality aversion, combined with the structural features of the economy and its market power on world markets. For a small country, real income effects will be strictly negative, while inequality effects may be positive or negative, depending on the relative endowment structure of the economy. For a large country, it is possible for both effects to work in the same direction. However, in this case, note that positive terms-of-trade gains will slow any rise (or slow any fall) in capital income shares, from equation (4.35). This in turn means that terms of trade effects will tend to mitigate the inequality effects of protection.

On the basis of equation (4.35), (4.36) and (4.37), we can summarize our discussion above with the following observations about import protection and inequality in the 2x2 Heckscher-Ohlin model.

Observation 4.1 *In a small labor-abundant Heckscher-Ohlin economy where the poor receive only or mostly labor income, and where the mean real-income effects of import protection are negative, the effect of import protection on welfare through mean income is magnified by the impact through inequality. Because of this magnification effect, net effects remain unambiguous and negative.*

Observation 4.2 *In a small, capital-abundant Heckscher-Ohlin economy, where the poor receive only or mostly labor income, and where the mean real-income effects of import protection are negative, the effect of import protection on welfare through mean income is at least partially offset by the impact through inequality. Hence the net welfare effect is ambiguous. It depends on the specification and parameterization of the underlying social welfare function.*

Observation 4.3 *The impact of protection on inequality as measured by the Atkinson and Gini indexes will be weaker, in a Heckscher-Ohlin economy, for large countries. This is because of terms of trade effects from equation* (4.35)*, which will dampen the goods-price to factor-price transmission mechanisms at play.*

4.4.3 Inequality effects in the 2x3 Ricardo-Viner model

Next, consider the specific factors model. We can make a similar manipulation of equations (4.33) and (4.34), like that yielding equations (4.36) and (4.37), for the standard 2-good, 3-factor model. This yields equations (4.38) and (4.39). Again, if we assume that inequality follows from the ownership pattern of (specific) capital, then in this case a shift in income shares through protection from more to less concentrated factors (in terms of the concentration of factor ownership) yields a reduction in inequality. The same points then follow, as before, with regard to country size and inequality effects in the Heckscher-Ohlin model. Otherwise, the impact of protection on inequality depends on the pattern of relative wage and ownership effects.

$$
\begin{aligned}
\frac{\partial I_A}{\partial \tau} &= -n^{\frac{-\theta}{1-\theta}} \left\{ \sum_h \left[\beta_{k1} \left(\omega_{k1}^h - \frac{1}{n} \right) + \beta_{k2} \left(\omega_{k2}^h - \frac{1}{n} \right) + \frac{1}{n} \right]^{1-\theta} \right\}^{\frac{\theta}{1-\theta}} \\
& \sum_h \left\{ \left[\beta_{k1} \left(\omega_{k1}^h - \frac{1}{n} \right) + \beta_{k2} \left(\omega_{k2}^h - \frac{1}{n} \right) + \frac{1}{n} \right]^{-\theta} \right. \\
& \left. \left[\frac{\partial \beta_{k1}}{\partial \tau} \left(\omega_{k1}^h - \frac{1}{n} \right) + \frac{\partial \beta_{k2}}{\partial \tau} \left(\omega_{k2}^h - \frac{1}{n} \right) \right] \right\} \qquad (4.38) \\
\frac{\partial I_G}{\partial \tau} &= -\frac{2}{n} \sum_h h \left[\frac{\partial \beta_{k1}}{\partial \tau} \left(\omega_{k1}^h - n^{-1} \right) + \frac{\partial \beta_{k2}}{\partial \tau} \left(\omega_{k2}^h - n^{-1} \right) \right] \qquad (4.39)
\end{aligned}
$$

We can summarize our results with respect to the Ricardo-Viner model as follows:

Observation 4.4 : *In a small Ricardo-Viner country, where any income effects from tariffs will be negative, protection may still be welfare improving if the induced change in inequality is large enough. This depends on the assumed functional form and parameterization of the social welfare function. If inequality is not improved, then the worsening inequality magnifies the negative efficiency effects.*

Observation 4.5 : *Unlike the Heckscher-Ohlin model, the impact of protection on inequality as measured by the Gini or Atkinson index is ambiguous in the Ricardo-Viner model when capital ownership patterns are the source of inequality. This is because the degree of concentration of specific-factor ownership (the weights applies in equations* (4.38) *and* (4.39) *to the induced changes in the specific-factor share of income) may vary between the two sectors, and because the impact on capital income will have opposite sign depending on the sector protected.*

This last observation associates the impact of protection on inequality to differences in ownership patterns for specific factors. If the ownership of the import-competing specific factor is sufficiently less concentrated than that for the export-sector, import protection will reduce inequality. However, if capital in the import-competing sector has relatively concentrated ownership, it will make the situation worse. For example, in a developing country where the poor have labor and land, and the rich labor and capital, protection will make the concentration of income worse, assuming the sector using capital is an import-competing sector. On the other hand, if ownership of land is very highly concentrated relative to capital, import protection may improve the distribution of income.

4.5 The Setting of Trade Policy

In this section, we examine the impact of inequality issues on the determination of import protection in an endogenous tariff model. We focus on the Heckscher-Ohlin model, where we obtain unambiguous results with respect to the impact of inequality considerations, in the region of a political equilibrium, on the observed tariff rate. We also discuss, briefly, the meaning of our results for endogenous tariff formation in a specific-factors model.

4.5.1 Capital rents

In our 2x2 Hecksher-Ohlin economy, we now assume consumers can be divided between workers and investors. The latter differentiate themselves by earning capital rents in addition to their labor earnings. Moreover, we assume that both groups are internally homogeneous and index them by $h = l, k$. Thus, we have n_k investors and n_l workers, each with real consumption defined as:

$$c^l = \frac{w_l + \omega_\tau^l \tau m}{f(p)} \tag{4.40}$$

$$c^k = \frac{w_l + w_k n_k^{-1} + \omega_\tau^k \tau m}{f(p)} \tag{4.41}$$

Assuming a CRRA utility function as in equation (4.6) with $\theta \neq 1$, we get:

$$u^l = \frac{1}{1-\theta}\left(\frac{w_l + \omega_\tau^l \tau m}{f(p)}\right)^{1-\theta} \tag{4.42}$$

$$u^k = \frac{1}{1-\theta}\left(\frac{w_l + w_k n_k^{-1} + \omega_\tau^k \tau m}{f(p)}\right)^{1-\theta} \tag{4.43}$$

Introducing a tariff change, the utility of each representative consumer is modified according to equations (4.44) and (4.45). We assume that each household is entitled to the same share of tariff revenue: $\omega_\tau^l = \omega_\tau^k = n^{-1}$.

$$
\begin{aligned}
\frac{\partial u^l}{\partial \tau} &= \left(\frac{w_l + \tau m n^{-1}}{f(p)}\right)^{-\theta} \frac{1}{f(p)} \\
&\left[\frac{\partial w_l}{\partial \tau} + m n^{-1} + \frac{\partial m}{\partial \tau} \tau n^{-1} - \frac{\left(w_l + \tau m n^{-1}\right)}{f(p)} \frac{\partial f(p)}{\partial \tau}\right]
\end{aligned} \tag{4.44}
$$

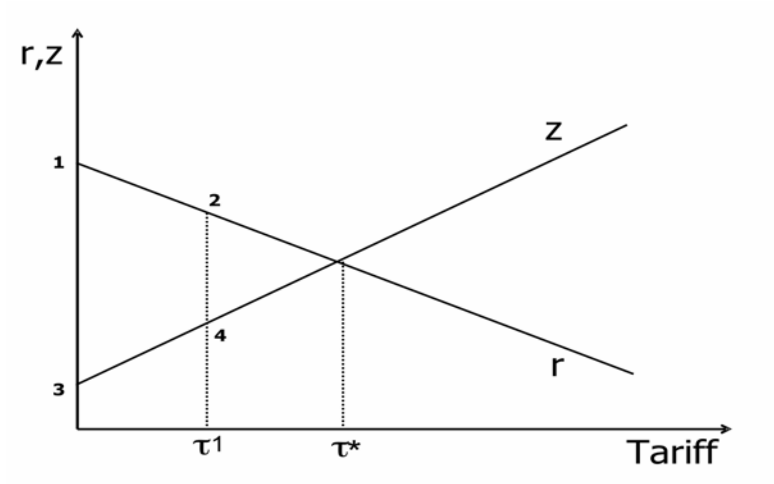
$$
\begin{aligned}
\frac{\partial u^k}{\partial \tau} &= \left(\frac{w_l + w_k n_k^{-1} + \tau m n^{-1}}{f(p)}\right)^{-\theta} \frac{1}{f(p)} \\
&\left[\frac{\partial w_l}{\partial \tau} + \frac{\partial w_k}{\partial \tau} n_k^{-1} + m n^{-1} + \frac{\partial m}{\partial \tau} \tau n^{-1} - \frac{\left(w_l + w_k n_k^{-1} + \tau m n^{-1}\right)}{f(p)} \frac{\partial f(p)}{\partial \tau}\right]
\end{aligned} \tag{4.45}
$$

In a small open labor-abundant economy, we know from the Stolper-Samuelson theorem that: $\frac{\partial w_l}{\partial \tau} < 0$, $\frac{\partial w_k}{\partial \tau} > 0$ and by construction $\frac{\partial f(p)}{\partial \tau} > 0$. Therefore, the workers are hurt by the tariff, both because of the efficiency and the distributional impact of the tariff. We obtain the optimum tariff for each group by equating (4.44) and (4.45) to zero. The workers' optimum tariff is zero, since tariffs reduce their share of total income (itself decreasing) and increase prices. For the investors, equation (4.46) shows the first order equilibrium condition.

$$
\begin{aligned}
\frac{\partial w_k}{\partial \tau} n_k^{-1} + \left(n^{-1} m + \frac{\partial m}{\partial \tau} \tau n^{-1} - \frac{\tau m n^{-1}}{f(p)}\right) &= \frac{\left(w_l + w_k n_k^{-1}\right)}{f(p)} \frac{\partial f(p)}{\partial \tau} - \frac{\partial w_l}{\partial \tau} \\
r &= z
\end{aligned} \tag{4.46}
$$

The left hand side term represents investor tariff gains (r) that compensate for the tariff cost (z), which are shown in the right hand side. Note that these costs are the same as the net welfare effect (folding in tariff revenues) on workers. While r is a combination of increased capital earnings and the tariff revenue, z includes the rise in the consumer goods price and the reduction in labor earnings. Moreover, both components are not monotonically related to tariffs. In particular, $\frac{\partial r}{\partial \tau} < 0$ and $\frac{\partial z}{\partial \tau} > 0$. We depict the investors' optimal tariff in Figure 4-1 as $\tau = \tau*$.

Figure 4-1: Investor rents with respect to tariff levels



For any given tariff, the net rents are equivalent to the area between both curves. In the figure, this is equal to the trapezoid connected by points 1, 2, 3 and 4, when $\tau = \tau^1$. From equation (4.46) the optimum tariff τ^* for investors is the intersection of r and z. Beyond this point, the increase in capital rents is offset by the increase in prices and the decrease in wage earnings.

As a group, investor rents R are given by:

$$R(\tau) = n_k \int (r - z)\, d\tau \tag{4.47}$$

where r and z are defined in equation (4.46). Moreover, $R(\tau) > 0$, $\frac{\partial R(\tau)}{\partial \tau} > 0$, and $\frac{\partial^2 R(\tau)}{\partial \tau^2} < 0$ over the relevant range.

4.5.2 The government's maximization problem

At this point, we could invoke a variety of different political economy models to generate political underpinnings for the setting of an equilibrium tariff in the political marketplace. These models have been extensively analyzed in the recent literature[13], and following Helpman (1995) we note that many of these can be represented, in reduced form, by the now standard political support function.

[13] See for example Helpman (1995) and Grossman and Helpman (2002).

Direct democracy is a rare political mechanism and generally public policies are decided by representative governments that balance conflicting interests. From Hillman (1989) we know that when one of the factors is sufficiently concentrated across only a few households, these individuals can organize to form pressure groups and overcome the free-rider problem. In such cases, Grossman and Helpman (2002) have demonstrated that in the reduced form the policy maker has two arguments to consider: the general interest and the interest of special groups (for example capital owners and unions). The interest of the government can follow from electoral support when social welfare is increasing and electoral contributions go with lobbying. For example, in a poor labor-abundant country investors can offer a contribution ρ (where $\rho \leq R$) to induce the policy-maker to increase import protection.

The precise weight the policy maker assigns to each group is established by her political support function, as in equation (4.48).

$$U(\tau) = \lambda_1 SW(\tau) + \lambda_2 \rho(\tau) \tag{4.48}$$

where U is the policy-makers utility, where we assumed that the tariff level is the only policy instrument of the government. The weights λ characterize the political system (how important are the contributions for the electoral campaign) and the policy-makers preferences (how she values reelection against more contributions).

Conditional on the particular values of these weights, she maximizes her utility by the first order condition shown in equation (4.49).

$$\frac{\partial U(\tau)}{\partial \tau} = \lambda_1 \frac{\partial SW(\tau)}{\partial \tau} + \lambda_2 \frac{\partial \rho(\tau)}{\partial \tau} = 0 \tag{4.49}$$

where $\frac{\partial SW(\tau)}{\partial \tau}$ has been already defined in equations (4.30) and (4.31). Since ρ is the fraction of the capital rents R that are assigned to political contributions, the additional term $\frac{\partial \rho(\tau)}{\partial \tau}$ is directly derived from equation (4.46). Furthermore, $\frac{\partial \rho(\tau)}{\partial \tau} > 0$ until the optimum tariff for investors is reached.

The additional element in the political mixture here is the effect of the tariff on social welfare $\left(\frac{\partial SW(\tau)}{\partial \tau}\right)$, which is not known forehand. The tariff impact on an inequality-adjusted social welfare function can induce changes to the standard results of the political support model. Thus, there is not necessarily a trade-off between both right hand side terms in equation (4.48) and in some cases they can reinforce each other.

If we analyze small open economies and consider only the social welfare implications, when the distributional effect of the tariff compensates for the efficiency losses we observe a positive optimum tariff rate. This can be the case only for capital-abundant countries. On the other hand, in poor countries the distributional and efficiency effects reinforce each other and the socially optimum tariff is zero, though the equilibrium rate may be positive.

When the influence of special interest groups is introduced, the previous partial results can change. In a rich capital-abundant country, the capital-owners have an incentive to lower tariffs, and if the workers can organize, they lobby to increase tariffs. The final outcome depends on the specific rents each group obtains and its political influence. In labor-abundant countries positive tariffs can be explained by the presence of an effective lobby, and in capital rich countries they can be explained by equity concerns that partially overcome free trade lobbying.

These multiple outcomes are summarized in the following observations:

Observation 4.6 *: In a Hecksher-Ohlin world, with homogeneous labor owners, concentrated capital and a policy-maker that cares about equity and assigns no weight to political contributions, the government's optimum tariff is higher in small capital-abundant countries than in small labor-abundant countries.*

Observation 4.7 *: In a Hecksher-Ohlin capital-abundant economy, with greater inequality aversion, capital owners will lobby for lower tariffs, while the government will be more protectionist because of equity reasons than otherwise. Equity concerns then offset to some extent pressure for lower tariffs.*

Observation 4.8 *: In a Hecksher-Ohlin labor-abundant economy, with greater inequality aversion, the government will favor low tariffs for equity and efficiency reasons, but will be lobbied by capitalists for higher tariffs. Equity concerns then offset to some extent pressure for higher tariffs.*

Basically, when the distributional effects are not significant enough to upset the efficiency losses imposed by the tariff, the common results of the literature emerge and higher tariffs are directly associated to the weight and the contributions of special interest groups. At the same time though, in the presence of distributional concerns rich countries tend to impose higher tariffs than otherwise. Relatively high average tariffs across a subset of capital-rich countries can then be seen as a consequence of inequality considerations by the policy-maker, as well as the presence of influential unions. In poor countries positive tariff rates are a direct consequence of the investor lobby overcoming both equity and efficiency concerns. In developing countries with a political system that assigns a significantly higher weight to social welfare than average, tariffs should remain lower than otherwise.

A similar analytical exercise can be carried out with a 2x3 specific-factors model. In particular, ignoring equity concerns, we have an equilibrium tariff that balances the efficiency effects of the tariff against the interests of owners of sector 1 and sector 2 capital. However, unlike the results for the Heckscher-Ohlin model we developed here, we will not then have unambiguous results when we add inequality to the policy objective mix. This is because, from Observations 4.4 and 4.5, the inequality impact of a tariff may itself be ambiguous. If a tariff reduces inequality in the region of the political equilibrium, we would again expect the equity-conscious government to be more protectionist than otherwise. If not, we expect the opposite to hold.

4.6 Conclusions

We have explored theoretical linkages between import protection and the household distribution of income. This complements the existing literature that links trade policy to factor incomes and the functional distribution of income, which is well developed in the literature. The main insight of this literature is provided by the Stolper-Samuelson theorem and constitutes a first step in our analysis. In a general equilibrium context, tariff changes ultimately affect the household distribution through variations in ownership patters in conjunction with Stolper-Samuelson effects. To model ownership structures, we used the Heckscher-Ohlin and Ricardo-Viner trade models. Within both frameworks, we are then able to produce theoretical predictions between trade protection, country size, level of development, and personal income inequality.

Another contribution of this paper is that we examine the formal link between social welfare and the equilibrium determinants of the distribution of income. Using Sen-type social welfare functions, we decompose the general equilibrium welfare effects of import protection into real income level and distribution components. The dispersion component is estimated using Gini and Atkinson inequality indexes. With these explicit inequality derivatives we map import protection to inequality-adjusted welfare. In addition, when standard trade models are employed this framework also yields predictions relating social welfare with protection, country size and levels of development. In conjunction with the relevant inequality index, the general form of the decomposition of welfare and inequality we develop here may also be useful in applied general equilibrium applications focused on inequality.

Once the distributional effects of trade liberalization are determined, we can apply endogenous tariff formation models to assess how the optimum tariff is affected by equity concerns. In representative democratic systems, we find that positive optimum tariffs can be sustained in capital-abundant countries even when the policy-maker assigns a low or zero weight to the contributions of special interests groups. In this case, the positive distributional effect of import protection can offset or compensate the efficiency losses of reduced trade. In poor countries, characterized by the relative abundance of labor, positive tariffs are explained by the influence of special interest groups (i.e. capitalists) that heavily lobby for higher tariffs. Thus, import protection in developing countries not only diminishes social welfare through efficiency and equity considerations, but also signals the economic and political weight of the capital-owners.

4.7 Appendix to Chapter 4

Summary of the variable's definitions

Variable	Definition	Variable	Definition
n	Number of households	x	Final goods supply function
h	Household index	g	Production function
y	Individual income	v_j^h	Endowment of factor j owned by h
$\overline{y}$	Mean income	a	Unit input coefficients
θ	Inequality aversion parameter	$\overline{v}$	Fixed factor endowment
SW	Social welfare function	r	Revenue function
I	Inequality index	w	Factor price
E	Equality index	m	Imports
p_c	Consumer price	τ	Import tariff
p	Composite price	ω_j^h	Share of factor j owned by h
c^h	Composite consumer good for h	ω_τ^h	Share of tariff revenue owned by h
u^h	Household utility	P^*	World prices
c	Aggregate consumption	s	Factor income
ϕ	Aggregate welfare	F	Functional distribution of income
ψ	Household utility function	S	Ownership matrix
e	Expenditure function	β_j	National income share of factor j
i	Final goods index	k	Capital index
j	Production factors index	l	Labor index
v	Factor endowments		

Chapter 5

Trade Policy and the Household Distribution of Income[1]

5.1 Introduction

In the early modern literature on economic development, Kuznets (1955, 1963) first noted the apparent relationship between the household distribution of income and growth. He stressed that rising income inequality seemed to be a normal process –a "stylized fact." Distributions seemed to first worsen and then improve with rising per-capita incomes. There has since been a sizeable literature in development economics, starting with Kuznets, Champernowne (1953), and Tinbergen (1956).

Taken together with the more recent literature on openness and growth, a logical development in the recent empirical literature has been the search for three-way linkages between openness, growth, and the personal/household distribution of income.[2] This includes general empirical evidence on the impact of international trade on income distribution and poverty in developing countries (McCulloch, Winters and Cirera, 2001; Winters 2000b). It also includes the growing computational literature on the household impact of policy reform. Along these lines, Devarajan and van der Mensbrugghe (2000)

[1] **This chapter is based on "Trade Policy and the Household Distribution of Income" by Francois and Rojas-Romagosa (2004a).**

[2] While the emphasis of this paper is on the personal and household distribution of income, there has also been a large recent literature on the role of trade in the evolution of the functional distribution of income. See Richardson (1995), and Feenstra and Hanson (2004).

have examined the household impact of trade policy reform in South Africa, Hertel *et al.* (2000) have explored the impact of multilateral trade liberalization on poverty reduction, Ianchovichina *et al.* (2000) have examined reform and distribution in Mexico, Robilliard *et al.* (2001) have examined the recent crisis in Indonesia, and Khan has focused on tax reform in Bangladesh (1997).

The recent econometric work in this area draws on data developed by Deininger and Squire (1996). Higgins and Williamson (1999) find evidence that demographical sources are the most important factor for explaining the distribution of income. They also find that, once one controls for demographic structure and openness to trade, there is strong evidence for Kuznets' inverted-u curve. However, the evidence of a link between income distribution and openness is mixed. Dollar and Kraay (2000) conclude that the income of the poor tends to grow at the same rate as economy-wide income. In other words, in contrast to the Kuznets-U reported elsewhere in the literature (like Higgins and Williamson), their results suggest that economic growth does not cause a relative deterioration in the mean income of the poor. In addition, they find that the share of growth following from trade does not significantly affect the income share of the poor either. Barro (2000) finds that inequality in developing countries is negatively correlated with economic growth, while Adams (2003) uses a new dataset to argue that growth is important for poverty reduction.

In this paper we analyze the impact of trade and tariffs on the observed cross-country variation in the household distribution of income. We build on results of the literature on production and the functional distribution of income in the 2x2 and higher-dimensional manifestations of the 2x2 trade model (Kemp 1956, Mussa 1979, Ethier 1982, Deardorff 1979, 1982) to link import protection to variations in the household distribution of income.[3] Working with a new dataset on inequality, we then examine cross-country variation in inequality with respect to import protection. Results are fully consistent with the predictions of the factor-intensity model of trade. Our regression results suggest that import

[3]The theoretical literature closest to the present paper includes Bourguignon and Morrisson (1990) and Spilimbergo et al. (1999). Both studies use an ownership matrix to write a general function of the personal income distribution in terms of endowments, tariffs and the ownership structure. A related paper is Galor (1994). He includes tariffs in his general equilibrium Overlapping-Generations model, but centers his analysis on the redistributive effects of the tariff revenues. He does not consider the efficiency and distributional effects caused by the tariffs.

protection makes income distribution worse for countries in labor-intensive diversification cones. This relationship shifts to one of falling inequality as incomes rise and we move to capital-intensive diversification cones. Results also suggest that OLS-based inequality regressions might best be viewed with some suspicion.

The paper is organized as follows. The next section develops a formal representation of inequality in general equilibrium. We then explore linkages between trade policies, the pattern of production, and inequality. This is followed by an econometric assessment of the linkages between trade policy and the observed pattern of inequality. The last section offers a summary and conclusions.

5.2 Production and trade in general equilibrium

We start with a dual representation of trade in general equilibrium for a single country. This involves adopting the following set of assumptions:

- Rational behavior by households and firms;
- Complete and perfectly competitive markets;
- Convex preferences, with neoclassical production functions;
- Convex technology for a composite consumption good;
- Identical and strictly quasi-concave composite good aggregation technologies across households.

Given these assumptions, we can define the core general equilibrium system for demand and production on the basis of expenditure and revenue functions, with expenditure defined in terms of the composite consumption good. Income distribution indexes (and social welfare) are then developed as a set of side equations from the core general equilibrium system.

5.2.1 The core general equilibrium system

We assume that all households consume the composite good c. This means we can represent aggregate expenditure e as a function of aggregate consumption c and prices p. This is represented by equation (5.1)

$$e(p, c) = c \cdot f(p) \tag{5.1}$$

In equation (5.1), $f(p)$ is the homothetic price index for the composite good.

On the production side, we assume standard neoclassical production functions with constant returns to scale: $x_i = g_i(v_{ij})$, where $g_i(\cdot)$ is the production function for good i and v_{ij} is the use of factor j in the production of good i. If we define unit input coefficients as a_{ji} then we also have: $1 \leq g_i(a_{ji})$. Endowment constraints are then $\sum a_{ji} x_i \leq \overline{v}_j$. From these, we can then define the economy-wide revenue function with respect to goods prices and endowments, as in equation (5.2).

$$r(p, v) = \max_{x_i, a_{ji}} \left\{ \sum_i p_i x_i \mid \sum_i a_{ji} x_i \leq \overline{v}_j \text{ and } 1 \leq g_i(a_{ji}) \;\; \forall i, j \right\} \tag{5.2}$$

From the envelope theorem and the properties of the revenue function r, factor incomes and goods production can be expressed in terms of the value of the partial derivatives of the revenue function, evaluated at the equilibrium set of prices:

$$\frac{\partial r(p, v)}{\partial v_j} = w_j = w_j(p, v) \qquad \forall j \tag{5.3}$$

$$\frac{\partial r(p, v)}{\partial p_i} = x_i = x_i(p, v) \qquad \forall i \tag{5.4}$$

We are ultimately interested in making empirical comparisons across countries. Empirically, this implies taking actual world prices as given by the data, and characterizing individual countries within this set of world prices. Taking equations (5.3) and (5.4) above in conjunction with equations (5.1) and (5.2), we can write the open-economy general equilibrium system for production, consumption, and trade for an individual country as follows:

$$
\begin{aligned}
c^h f(p) &= \left(\sum_j w_j(p,v)\cdot v_j^h\right) + \omega_\tau^h \tau m \qquad \forall h && (5.5)\\
m &= \sum_h c^h \cdot f(p) - x(p,v) && (5.6)\\
e(p,c) &= \sum_h \left[\left(\sum_j w_j(p,v)\cdot v_j^h\right) + \omega_\tau^h \tau m\right] && (5.7)\\
p &= P^* + \tau = 1 + \tau && (5.8)
\end{aligned}
$$

In equations (5.5) – (5.8), we have assumed that a tariff of τ is imposed on imports from the world, while world prices are normalized to one. The term ω_τ^h is the household share of tariff revenue, while v^h is the household ownership share of factors and c^h is household consumption. We will generally assume that the household share of tariff income is n^{-1} where n indexes the number of households. Equation (5.5) sets the value of household consumption equal to the household budget, equation (5.6) defines imports on which import tariff revenue is generated, equation (5.7) sets economy wide expenditure equal to receipts, and equation (5.8) defines the wedge between world and domestic prices. Together, equations (5.5) – (5.8) define an equally dimensioned set of relationships and unknowns: c^h, c, m and p.

5.2.2 Household Inequality

As noted earlier, both the older and the more recent literature on trade and the distribution of income have been focused on the functional distribution of income. The functional distribution of income is also an important building block here for the representation of the household or personal distribution of income. Starting with factor incomes s, they follow directly from the endowment stock and the properties of the revenue function, as represented by equation (5.9).

$$
s_j = w_j v_j = r_{vj}(p,v)\cdot v_j \tag{5.9}
$$

In reduced form, the functional distribution of income $F(s)$ is then an artifact of the equilibrium matching of preferences and the technology set, given our endowment vector.

$$F(s) = F(p, v) \tag{5.10}$$

The household distribution of income follows from the combination of factor incomes w_j, the vector of endowments, and the household ownership share in factors of production, ω_j^h. It will also depend on the distribution of tax revenue (tariffs in the present context), again represented by a household share parameter, this one applied to import tax revenues. This is shown in equation (5.11), which gives the basic definition of household income in terms of its primary components. By substitution from equation (5.5), this is also shown as a function of equilibrium prices, the production technology set, and the endowment set.

$$\begin{aligned} y^h &= \left(\sum_j w_j \cdot v_j \cdot \omega_j^h\right) + \omega_\tau^h \tau m \\ &= \left(\sum_j r_{v_j}(p, v) \cdot v_j \cdot \omega_j^h\right) + \omega_\tau^h \tau m \qquad (5.11) \\ c^h &= \frac{y^h}{p_c} \end{aligned}$$

where $1 \geq \omega_j^h \geq 0$ and $\sum \omega_j^h = \sum \omega_\tau^h = 1$. In reduced form, the household distribution of income $F(y)$ is a consequence of endowments, the technology set, preferences, the endowment vector, and the ownership matrix of coefficients $\omega_j^h(S)$. From equation (5.12) we thus have:

$$F(y) = F(p, v, S) \tag{5.12}$$

We introduce two inequality indexes to our analytical framework: the Gini coefficient I_G and the Atkinson index I_A. Constant relative risk aversion (CRRA) preferences yield the Atkinson inequality index directly as a natural metric for the income distribution component of social welfare (see Atkinson 1970). In this sense, Atkinson's index maps naturally into a social welfare function and in particular, to Sen's (1974) preferred definition of social welfare. In contrast, the Gini coefficient implied social welfare function is axiomatic, in that we do not have an obvious mapping, through aggregation, from individual preferences to an aggregate social welfare function. This follows because the Gini coefficient (and hence the implied social welfare function) is rank sensitive. The formal definitions of these inequality indexes are provided in equations (5.13) and (5.14).

$$I_A = 1 - \left[\frac{1}{n} \sum_h \left(\frac{y^h}{\overline{y}} \right)^{1-\theta} \right]^{\frac{1}{1-\theta}} \tag{5.13}$$

$$I_G = 1 + \frac{1}{n} - \frac{2}{n^2 \overline{y}} \left(y^1 + 2y^2 + ... + ny^n \right) \tag{5.14}$$

In equation (5.13), θ corresponds to the coefficient of relative *inequality* aversion assuming CRRA preferences. In equation (5.14) the Gini coefficient is defined as twice the area between the Lorenz curve and the 45-degree line, where we have arranged households so that $y^1 \geq y^2 \geq ... \geq y^n$.

We can also write the Atkinson and Gini indexes in terms of system fundamentals. Making a substitution from (5.9) into (5.13) and (5.14), we obtain equations (5.15) and (5.16).

$$I_A = 1 - \left\{ \frac{1}{n} \sum_h \left[\frac{n \left(\sum_j r_{v_j} (p, v) \, v_j \omega_j^h \right) + n \omega_\tau^h \tau m}{y} \right]^{1-\theta} \right\}^{\frac{1}{1-\theta}}$$

$$I_A = 1 - \left\{ n^{-\theta} \sum_h \left[n^{-1} + \sum_j \beta_j \left(\omega_j^h - n^{-1} \right) \right]^{1-\theta} \right\}^{\frac{1}{1-\theta}} \tag{5.15}$$

$$I_G = 1 + \frac{1}{n} - \frac{2}{n} \sum_h h \left[\frac{\left(\sum_j r_{v_j} (p, v) \, v_j \omega_j^h \right) + \omega_\tau^h \tau m}{y} \right]$$

$$I_G = 1 + \frac{1}{n} - \frac{2}{n} \sum_h h \left[n^{-1} + \sum_j \beta_j \left(\omega_j^h - n^{-1} \right) \right] \tag{5.16}$$

where $\beta_j = \frac{w_j v_j}{y}$, which is the national income share accruing to factor j.

5.3 Tariffs and household inequality

5.3.1 Generalized effects

From equations (5.15) and (5.16) above, inequality is a function of the first two moments of the household distribution of income. This is especially obvious with the Atkinson index.[4]

The impact of a tariff on the household income distribution follows from differentiation of equations (5.15) and (5.16). This is shown in equations (5.17) and (5.18) below.

$$\begin{aligned}
\frac{\partial I_A}{\partial \tau} &= -\frac{1}{1-\theta}\left\{n^{-\theta}\sum_h\left[\left(\sum_j \beta_j\left(\omega_j^h - n^{-1}\right)\right) + n^{-1}\right]^{1-\theta}\right\}^{\frac{\theta}{1-\theta}} \\
&\quad \left(n^{-\theta}\sum_h (1-\theta)\left\{\left[\left(\sum_j \beta_j\left(\omega_j^h - n^{-1}\right)\right) + n^{-1}\right]^{-\theta}\left(\sum_j \frac{\partial \beta_j}{\partial \tau}\left(\omega_j^h - n^{-1}\right)\right)\right\}\right) \\
&= -n^{\frac{-\theta}{1-\theta}}\left\{\sum_h\left[\sum_j \beta_j\left(\omega_j^h - n^{-1}\right) + n^{-1}\right]^{1-\theta}\right\}^{\frac{\theta}{1-\theta}} \\
&\quad \sum_h\left\{\left[\sum_j \beta_j\left(\omega_j^h - n^{-1}\right) + n^{-1}\right]^{-\theta}\left[\sum_j \frac{\partial \beta_j}{\partial \tau}\left(\omega_j^h - n^{-1}\right)\right]\right\} && (5.17) \\
\frac{\partial I_G}{\partial \tau} &= -\frac{2}{n}\sum_h h\left[\sum_j \frac{\partial \beta_j}{\partial \tau}\left(\omega_j^h - n^{-1}\right)\right] && (5.18)
\end{aligned}$$

Recall that the term β_j represents the national income share of factor j. Note that we also have an inverse income weighting, by a factor of θ, in equation (5.17) applied to induced changes in the deviation from the mean component of household income – $\beta_j\left(\omega_j^h - n^{-1}\right)$. The weighting of induced changes in income for the Gini index depends on the ranking of individual households on the relative income scale. In both cases, the changes in income in turn depend on Stolper-Samuelson derivatives, and induced price changes that follow from tariff changes. This set of induced price changes is expressed

[4] While the functional form is different, the social welfare functions underlying other income distribution indexes yield a similar result, though with different weights in the variance component of the welfare function. The CRRA function yields a particularly clear and parsimonious reduced form.

in equation (5.19), where the term $\frac{\partial\beta_j}{\partial\tau}$ depends on system fundamentals and consequent Stolper-Samuelson relationships. The last line of equation (5.19) expresses this relationship in elasticity terms.

$$\begin{aligned} \beta_j &= \frac{r_{v_j} v_j}{y} \\ &\Rightarrow \frac{\partial \beta_j}{\partial \tau} = \frac{\partial r_{v_j}}{\partial p}\frac{\partial p}{\partial \tau}\frac{v_j}{y} - \frac{r_{v_j} v_j}{y^2}\frac{\partial y}{\partial \tau} \\ \varepsilon_{\beta_j,\tau} &= \varepsilon_{r_{v_j},p}\varepsilon_{p,\tau} - \varepsilon_{y,\tau} \end{aligned} \tag{5.19}$$

5.3.2 Tariffs and Household Inequality in the Heckscher-Ohlin Model

Consider the application of equations (5.17) and (5.18) to a standard two-factor, two-good Heckscher-Ohlin model. If we apply the additional normalization that all households hold a claim on one unit of labor, then our inequality indexes can be manipulated to yield equations (5.20) and (5.21). (We have also assumed a neutral redistribution of the tariff revenue).

$$\begin{aligned} \frac{\partial I_A}{\partial \tau} &= -n^{\frac{-\theta}{1-\theta}} \left\{ \sum_h \left[\beta_k \left(\omega_k^h - n^{-1}\right) + n^{-1}\right]^{1-\theta} \right\}^{\frac{\theta}{1-\theta}} \\ &\quad \sum_h \left\{ \left[\beta_k \left(\omega_k^h - n^{-1}\right) + n^{-1}\right]^{-\theta} \left[\frac{\partial \beta_k}{\partial \tau}\left(\omega_k^h - n^{-1}\right)\right] \right\} \end{aligned} \tag{5.20}$$

$$\frac{\partial I_G}{\partial \tau} = -\frac{2}{n}\sum_h h \left[\frac{\partial \beta_k}{\partial \tau}\left(\omega_k^h - n^{-1}\right)\right] \tag{5.21}$$

Inequality is purely a function of the allocation of capital in this model. At the same time, the impact of the tariff is then a function of which sector is protected. If protection leads to an increase in wages and a drop in capital income, inequality is reduced. Alternatively, if capital income is protected, we will see a rise in inequality.

On the basis of equation (5.20) and (5.21) we can summarize our discussion above with the following observations about import protection and inequality in a 2x2 Heckscher-Ohlin economy.

Observation 5.1 *: In a labor-rich 2x2 Heckscher-Ohlin economy under the assumption that poorer households by definition derive income only or mostly from labor, then observed inequality as measured by the Gini or Atkinson indexes will be greater, as import protection in labor rich countries will help capital owners, who receive relatively low weight in equations* (5.20) *and* (5.21).

Observation 5.2 *: In a capital-rich (developed) Heckscher-Ohlin economy, under the assumption that the poor receive only or mostly labor income, import protection within a Heckscher-Ohlin economy means a lower observed inequality as measured by the Gini or Atkinson indexes.*

Observation 5.3 *: When making comparisons across 2x2 countries at a given set of world prices (as will be the case for cross-section inequality regressions), Observations 5.1 and 5.2 can be reinterpreted as saying that tariffs will be linked to greater observed inequality in capital-poor countries and less inequality in capital-rich countries.*

While Observation 5.3 has empirical appeal, the classic 2x2 model is actually problematic when we confront it with real data. Countries present a range of relative endowments, while we might reasonably expect the global trade matrix to include more than two broad classes of goods. Following Deardorff (1979), if we assume Heckscher-Ohlin technologies and a range of goods, we can then rank goods by their relative factor intensities in production. Countries in turn can be ranked by their location in a chain of product diversification cones. More labor intensive countries will be located in diversification cones characterized by more labor intensive goods. We can then generalize Observation 5.3 to allow for a continuum of goods and factor intensity rankings.

Observation 5.4 *: Assuming a Deardorff-type continuum of goods, then when making comparisons across countries within a given diversification cone, at a given set of world prices higher tariffs will be linked to greater observed inequality in cones occupied* *by capital-poor countries and less observed inequality in cones occupied by capital-rich countries.*

Observation 5.5 *: Assuming a Deardorff-type continuum of goods, as we move across diversification cones from more labor-intensive to more capital-intensive goods, the marginal impact of a tariff on inequality will fall.*

5.4 Econometrics

5.4.1 Data

We turn next to an empirical examination of the linkages between import protection and the household distribution of income. Almost exclusively, the current literature on openness and household inequality is built on the Deininger and Squire (1996) dataset. This dataset mixes inequality indexes drawn from income surveys, expenditure surveys; surveys based on net- and gross- incomes; and household and individual reference units. The same is true of the household distribution data reported by the World Bank (World Development Indicators) and the United Nations (Human Development Report). The assumption is that different income concepts and reference units are broadly comparable. Actually though, it turns out that they are not, and in many cases they do not even convey the same broad information set.[5]

In our data we control for comparability between different sources of household distribution data. To avoid measurement error problems, we therefore work with inequality series that combine comparable definitions. To be consistent with our theoretical framework, we use inequality data that we have either estimated directly from gross income household survey data, or that are reported as coming from gross income household surveys (where the underlying share data was not available).[6] We have two indexes of inequality, the Atkinson index[7] and the Gini coefficient. Because the Gini coefficient is commonly reported from household surveys, even when the share data is not supplied,

[5] Atkinson and Brandolini (2001).

[6] We take the inequality data from the World Income Inequality Database (2000).

[7] We use $\theta = 1$.

we have a broader sample of countries with estimated Gini coefficients. These data are shown in Table 5-1. Values are centered on 1994 (though inequality indexes are taken from the nearest available year in a 4-year span before and after 1994). Per-capita income in the sample, in 1995 dollars, ranges from \$532 (China) to \$40,515 (Japan). The mean per-capita income is \$10,774. The trade-weighted tariff for the sample ranges between 0 percent (Hong Kong) and 43.7 percent (Thailand), with a sample average duty of 9.9 percent. The mean Gini coefficient from our gross household data is .40, with a maximum and minimum value of .62 (South Africa) and .22 (Slovakia). The mean Atkinson index from our gross household data is .23, with a maximum and minimum of .53 (South Africa) and .08 (Ukraine).[8]

5.4.2 Regression Model and Results

As a starting point, we specify a reduced-form relationship between income level and inequality, formally known as a Kuznets-type inverted-u relationship. In specifying a reduced form for inequality in this way, we are building on a broad empirical literature.[9] To this, we add an indicator of the trade-weighted import duty. We also include an interaction term between duties and income levels, to reflect the possible role (as reflected in Observations 5.3 and 5.5) that relative development levels may play in the interaction between import tariffs and the distribution of income.

Formally, our estimation equation is as follows:

$$LI_i = a_0 + a_{PCI} \cdot y_i + a_{PCI^2} \cdot y_i^2 + b_T \cdot T_i + b_{T \cdot PCI} \cdot (T_i \cdot y_i) + \varepsilon_i \tag{5.22}$$

[8] These data are part of a larger project aimed at evaluating the impact of household data comparability problems on the current empirical literature on globalization, distribution of income, and poverty. See Chapter 3.

[9] See Kanbur (2000) for a recent survey. Following Higgins and Williamson (1999), we can distinguish two versions of the Kuznets hypothesis in the literature. The original (strong) version presented by Kuznets (1955), states that labor demand steers income inequality during the development process. On the other hand, the weak version stressed in the more recent literature recognizes that other factors can also be involved, and that these can reinforce or offset the basic demand forces at play. Deininger and Squire (1996, 1998) do not find support for the unconditional Kuznets curve, while Higgins and Williamson (1999), also using the Deininger and Squire data set, find support for the weak Kuznets hypothesis.

where LI_i is a logistic transformation of the inequality index I for country i, y_i is the log of per-capita income (PCI), T_i is the weight-average tariff on merchandise imports, a and b are the coefficients to be estimated, and ε_i is the error term.

Table 5-2 first presents basic OLS results. The signs on coefficients are all as expected. However, coefficient estimates are insignificant for the tariff terms in the Gini coefficient regressions, while the Gini and Atkinson regressions both provide relatively poor fits to the inequality data. In this sense, initial results are consistent with the broad literature in this area. In other words, they are disappointing and offers mixed results. However, there is good reason to be suspicious of OLS results in this context, both in columns A and B in Table 5-2, and also in the literature in general. It is pretty well recognized that data from low-income countries tend to be of poorer quality that data from the OECD countries. In other words, we should expect that low-income country data tend to be noisier. Columns A and B report Chi-squared test statistics for the hypothesis of homoskedasticity, weighed against the alternative hypothesis of variance inversely related to level of development as proxied by income level. These results point strongly to heteroskedasticity, and hence to the preferability of weighted regressions. It is also worth recalling that the Gini and Atkinson indexes are based on the same set of income survey results. As such, the error structure of the Gini-based observations may tell us something about the error of the error-based regressions. As such, we may improve the accuracy of the our estimates of coefficient variance by estimating the two equations jointly.

The results of weighted SUR regressions are reported in columns C (Gini) and D (Atkinson) of Table 5-2. These estimates address both the heteroskedasticity issue, and the joint nature of the calculation of the Gini and Atkinson indexes. The results for all coefficients have the expected sign, and are highly significant – in the .001 to .002 range. The import-protection coefficients suggest import protection makes the distribution of income worse at the margin for countries in labor-intensive diversification cones, while this relationship shifts to one of falling inequality as incomes rise and we move to capital-intensive diversification cones. In other words, relative openness helps to improve inequality for the least developed countries. As such, the basic pattern is consistent with our discussion of a Heckscher-Ohlin world with a continuum of goods. It fits the prediction of factor-intensity based trade theory about the linkage between development level, protection, and the observed household income distribution.

To help interpret these results, Figure 5-1 plots the estimated marginal impact of a change in the tariff on inequality for the countries in our sample using the Gini coefficient and Atkinson index estimates from columns C and D of Table 5-2. This involves the combined effects of coefficients b_T and $b_{T \cdot PCI}$. Technically, as the regressions are in logs, this represents an estimate of the elasticity of the logistic transform of the Gini coefficient or Atkinson index with respect to a change in the power of the tariff $T = (1 + t)$. The turning point for the index is at a per-capita income of between \$5,474 (Gini coefficient) and \$8,780 (Atkinson). Recall that the Gini coefficient is more sensitive to redistribution in the middle-income ranges, while the Atkinson index is more sensitive to redistribution at the extremes. The turning points for the different indexes hence relate to inequality more in the middle of the income distribution (Gini) or at the upper vs. lower extremes (Atkinson). For countries below this income range, higher import protection is associated with marginal increases in observed income inequality. For countries above this level, relatively high import protection implies marginal reductions in observed inequality.[10]

5.5 Summary and Conclusions

In this paper we explore the relationship between import protection and the observed cross-country pattern of the household distribution of income. The theoretical linkages between import protection and the functional distribution of income (i.e. factor rewards) are well developed in the literature. Because the functional distribution of income is the first step in mapping import protection to the household/personal distribution of income, the existing functional distribution literature also provides insight into how import protection, through variations in ownership patterns in conjunction with Stolper-Samuelson effects, ultimately impacts the household distribution of income.

[10] Burgess (1976) reaches a similar conclusion for the United States, which is obviously one of the high-income countries in our sample. Working with input-output coefficients and a generalized Leontief structure, Burgess concludes that higher tariffs raise real wages with respect to the United States.

On the theoretical side, our contributions in this paper follow from an explicit general equilibrium formalization of linkages between import protection and standard measures of household income distribution. These inequality derivatives yield predictions for patterns in the observed variation in trade protection, level of development, and the level of inequality across countries. On the empirical side, we then work with a set of comparable inequality indexes (i.e. controlling for income definitions and survey units) to test these predictions. To do this, we add trade policy indicators to a set of standard Kuznets-U type inequality regressions. For the present exercise, which is focused on transmission from functional to household income patterns, gross income survey data has been used. In addition to the basic pattern of results (inequality-tariff linkages), our estimates also point to potential estimation problems with OLS-based unconditional Kuznets inequality regressions.

Regression results point to a highly significant pattern in the data fully consistent with the theoretical impact of tariffs on inequality in an n-good Heckscher-Ohlin world. While consistent with the predictions of theory, the pattern that emerges is not fully consistent with recent anti-globalization rhetoric. We find that in the lowest-income countries, relatively high import protection is associated with observations of greater inequality. It is in higher-income countries that higher import protection appears to improve the gross household distribution of income the most. As such, current import protection patterns impose welfare costs related to inequality exactly in those countries least able to offset distribution losses through a functioning social safety net.[11]

[11] In addition, if one takes distribution as an important component of social welfare, along the lines of Atkinson and Sen, then the implications for numerical analysis of trade policy appear to be first- rather than second-order in developing countries. Crude calculations by the authors not presented here, working with the estimated coefficients from the inequality regressions, suggest that in some cases adjusting trade policy assessments for distribution impacts may yield welfare effects on a par with those linked to efficiency currently stressed in the computational literature on trade policy.

5.6 Appendix to Chapter 5

Table 5-1: Country data

Country		Gini coefficient	Atkinson index	per-capita income 1994 (US$)	Trade weighted import tariff
AUS	Australia	0.440	0.317	20,190	9.7
BEL	Belgium	0.318	0.171	26,705	7.0
BLR	Belarus	0.284	0.128	2,172	10.5
BRA	Brazil	0.595	--	4,299	17.0
CAN	Canada	0.350	0.213	19,366	8.0
CHL	Chile	0.556	0.390	4,212	10.9
CHN	China	0.368	0.213	532	40.6
COL	Columbia	0.580	0.367	2,326	12.4
CRI	Costa Rica	0.455	--	3,408	8.0
DEU	Germany	0.349	0.197	29,645	7.0
DNK	Denmark	0.289	0.162	33,701	7.0
DOM	Dominican Rep.	0.490	--	1,482	17.7
ECU	Ecuador	0.530	--	1,563	8.5
ESP	Spain	0.330	0.169	14,528	7.0
EST	Estonia	0.378	--	3,063	0.5
FIN	Finland	0.295	0.161	24,473	5.0
GBR	Great Britain	0.404	0.298	18,772	7.0
HKG	Hong Kong	0.520	0.288	22,345	0.0
HND	Honduras	0.540	--	693	7.5
JPN	Japan	0.350	--	40,515	1.5
LTU	Lithuania	0.248	--	1,677	1.7
LVA	Latvia	0.240	--	1,940	2.6
MDA	Moldova	0.267	--	721	2.7
MEX	Mexico	0.477	--	3,406	12.2
NLD	Netherlands	0.316	0.172	26,372	7.0
NOR	Norway	0.302	0.120	32,553	4.9
NZL	New Zealand	0.401	0.270	16,056	9.4
PAN	Panama	0.560	--	3,006	11.0
PER	Peru	0.483	--	2,132	16.1
PHL	Philippines	0.462	0.302	1,061	20.4
POL	Poland	0.301	0.150	2,684	10.1
PRY	Paraguay	0.590	--	1,827	8.9
ROM	Romania	0.287	--	1,310	15.2
RUS	Russia	0.272	0.112	2,376	7.4
SGP	Singapore	0.422	0.260	23,017	0.0
SVK	Slovakia	0.215	--	3,220	6.2
SVL	El Salvador	0.530	--	1,612	8.7
SVN	Slovenia	0.281	0.119	9,053	12.1
SWE	Sweden	0.309	0.177	26,380	5.0
THA	Thailand	0.535	0.401	2,647	43.7
TWN	Taiwan	0.327	0.163	12,070	6.5
UKR	Ukraine	0.233	0.079	1,076	6.3
USA	United States	0.424	0.296	27,343	4.4
VEN	Venezuela	0.470	--	3,482	13.4
ZAF	South Africa	0.620	0.538	3,828	16.9

Tariff data and income data are from the World Development Report (various issues).

Income distribution data are from Chapter 3.

"--": Not available. This applies to cases where there is no distribution share information.

Table 5-2: Regression results for the Kuznets curve with tariffs

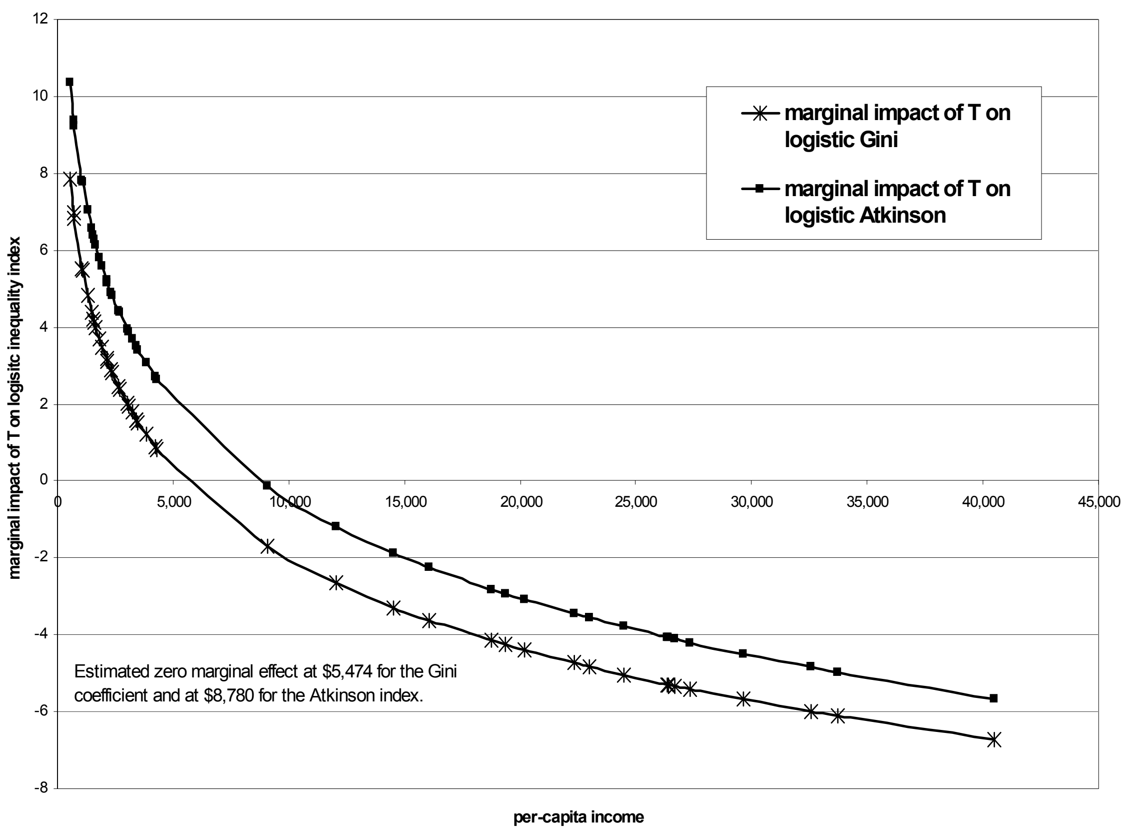
Coefficients		OLS estimates		SUR estimates, weighted †	
		A	B	C	D
		Logistic Gini	Logistic Atkinson	Logistic Gini	Logistic Atkinson
a_0	Intercept	-8.7939	-27.7185	-16.0398	-28.619
		(-1.60)	(-2.58)**	(-3.18)***	(-3.27)***
a_{PCI}	per-capita income	1.9179	5.7372	3.4302	6.0795
		-1.53	(2.39)**	(3.07)***	(2.95)***
a_{PCI^2}	squared per-capita income	-0.11	-0.3088	-0.1855	-0.3229
		(-1.58)	(-2.35)**	(-3.06)***	(-2.78)***
b_T	Tariffs: T= (1+t)	5.533	24.8864	28.9728	33.5697
		-0.61	(1.78)*	(3.80)***	(3.95)***
b_{T*PCI}	T*(per-capita income)	-0.4243	-2.6333	-3.3658	-3.697
		(-0.35)	(-1.47)	(-3.64)***	(-3.55)***
Heteroskedasticity tests					
Breusch-Pagan/Cook-Weisberg ‡		4.77, 0.029	3.92, 0.048		
Szroeter's test ‡		4.43, 0.035	2.68, 0.100		
‡ *Chi2, Prob>Chi2 w.r.t. income*					
Regression statistics					
Observations for each index		45	27	45	27
F, Pr>F		2.31, 0.075	2.78, 0.052	-57.07, 0.000	
R-squared for joint estimate		0.1873	0.3359	0.9092	
Total observations		45	27	72	
Degrees of freedom		40	22	62	

† Regressions are weighted, on the assumption that variance is inversely related to per-capita income. This is supported by heteroskedasticity tests shown in columns A and B. This simply means that data are noisier in lower-income countries, due to a mix of measurement error and more variability at lower levels of income in factors reflected in the reduced form Kuznets curve. According to Szroeter's Chi-squared test statistic for homoskedasticity, the hypothesis of variance monotonic in per-capita income is preferred to homoskedasticity at the .035 level for the Gini-based Kuznets curve and at the .100 level for the Atkinson-based Kuznets curve. The Breusch-Pagan/Cook-Weisberg Chi-squared test statistic also supports heteroskedasticity in per-capita income (at the .029 and .048 significance levels) for both regressions.

Regression results shown are for unweighted OLS, and for weighted iterated SUR regressions with robust standard errors, where t- and z-test statistics are reported in parentheses.

* means significant at the .10 level for a two-tailed test.
** means significant at the .05 level.
*** means significant at the .01 level.

Figure 5-1: Estimated marginal impact of tariffs, (1+t), on logistic inequality indexes



Chapter 6

Income Distribution, Development and Trade: An Empirical Review Using a New Inequality Dataset[1]

6.1 Introduction

In his seminal article of 1955, Kuznets introduced the concept that income inequality first increases and then decreases during the process of economic development. This hypothesis, now well-known as the Kuznets inverted-u hypothesis, has been central to the subsequent literature on income inequality. Following Higgins and Williamson (1999) we can distinguish two versions. The original (strong) version presented by Kuznets argues that labor demand drives income inequality during the development process. In the early stages of development, labor-saving technological change and structural change (urbanization and industrialization) widens inequality. Later on, these forces slow down and inequality is gradually reduced. In this unconditional version inequality is driven solely by labor demand.[2] Alternatively, the conditional (weak) version recognizes that other

[1] **This chapter is based on "Income Distribution, Development and Trade: An Empirical Review Using a New Inequality Dataset" by Francois and Rojas-Romagosa (2004d).**

[2] The influential paper by Lewis (1954) had the same implications. He assumed two sectors, one traditional with labor surplus and low wages and a modern sector with high wages. Growth was achieved by moving labor from one sector to the other. Inequality initially increased and then decreased, as a bigger share of workers received the higher wages of the modern sector.

factors can also be involved. These factors can reinforce or offset the basic labor demand forces at play –e.g. demographic transitions, resource endowments, governmental intervention and trade policy. For example, trade openness can increase the supply of labor-intensive goods in developed countries and thus decrease the income of unskilled workers and increase inequality. Therefore, this version of the Kuznets hypothesis is conditional on alternative factors and provides a better theoretical basis to conduct empirical studies.

There is a substantial literature that tests the continued validity of the original Kuznets' insight. However, the lack of reliable, comprehensible and consistent inequality data has seriously restricted the empirical analysis. Of these shortcomings, the most relevant has been the lack of comprehensive inequality time-series data. The hypothesis is meant to analysis the long-term relationship embodied in any development process and no single country has a long enough time-series that can gauge such a process. For instance, in countries that are already developed, the series do not go further back in time.[3] For developing countries inequality data are even scarcer and since they by definition are not developed yet, one can only partially test the hypothesis. These limitations have given rise to the use of cross-country regressions, which assume that each country's income level can proxy for the development stages of a prototypical economy. The use of cross-country regressions has well known drawbacks and it is an imperfect substitute for time-series analysis. Nonetheless, given the strong inequality data limitations it is the most widely used methodology.[4]

A second drawback of the earlier literature was that many studies used unreliable sources. This includes information driven from regional –not national– country surveys, and surveys not measuring all income sources. An important step toward addressing these issues was taken by Deininger and Squire (1996), who compiled a dataset that considered only inequality observations that fulfilled certain basic quality criteria. In addition, their dataset was larger and more comprehensive than previously available compilations. This has allowed the use of pooled estimation techniques.

[3] The longest series is from the United States and starts in 1944.

[4] The alternative approach is to use case-studies, as in the original paper by Kuznets.

Kanbur (2000) claims that the consensus on the distributional effects of growth has cycled in the postwar period.[5] However, the most recent empirical work in this area has been reshaped by the introduction of the Deininger and Squire inequality dataset (henceforth referred as DS), which has spurred the appearance of new studies that cast some doubt on earlier work. For this reason, one can divide the empirical literature in broad terms into sets dated before and after the appearance of the DS dataset. Using this particular inequality dataset, Deininger and Squire (1996, 1998), Higgins and Williamson (1999) and Barro (2000), among others, have tested for the presence of the Kuznets curve in a cross-section of countries.

Most recently, the DS has itself been found to have important limitations. Atkinson and Brandolini (2001) have criticized the consistency of the DS dataset and observe that the loose use of different inequality concepts (e.g. mixing net income, gross income, and expenditure observations) can lead to important changes in apparent inequality levels and patterns in OECD countries. In Chapter 3 we have conducted a similar analysis with regard to the consistent use of inequality definitions and reference units for a sample expanded to include developing countries in addition to the OECD. That paper also concludes that it is inadequate at best and misleading at worst to mix different inequality concepts. The additional measurement errors introduced by combining these concepts cannot be readily corrected by the use of fixed adjustments or dummy variables, as was proposed by Deininger and Squire.

To avoid these problems, in Chapter 3 we have constructed an alternative inequality dataset that consistently uses inequality concepts and reference units. Moreover, the dataset includes estimates of Atkinson indexes from inequality grouped data. This provides an alternative inequality indicator that can complement the widely used Gini coefficient.

Using this dataset, the aim of this paper is to reassess the empirical results concerning the Kuznets curve found in the recent literature that were based on the DS dataset. In addition, we also reexamine the relation between trade openness and inequality. We follow the empirical specifications used in the literature and we complement this with the

[5] Kanbur surveys the literature until they Deininger and Squire dataset appeared in 1996.

Atkinson index and with weighted linear regressions that help correct for the fact that inequality data measurement errors are associated with the level of development.

In line with the previous findings of the literature, we also conclude that the Kuznets hypothesis is an empirical regularity. Moreover, when income levels (the standard proxy for development level) are complemented with other inequality determinants we are able to explain a substantial portion of the cross-country variation in within-country inequality. At the same time, we find only weak support for the influence of trade openness with respect to inequality, when measured by three different indicators. Finally, the use of the Atkinson index (which is sensitive to income changes at the extremes) yields similar results to those of the Gini coefficient (which is sensitive to income changes around the mean).

Because our data extends into the very recent past, we are also able to examine recent inequality trends in the higher-income range, where we are pushing well beyond the income/development levels that underpinned earlier work. In this regard, we confirm recent findings by Atkinson (2003) and Gottschalk and Smeeding (2000) that inequality has been raising in countries at the highest income levels. Empirically, this appears as a significant cubic trend in inequality data. This outcome can be explained by the recent inequality trends in OECD countries. This does not directly contradict the Kuznets hypothesis, as the pattern persists across the sample range he had to work with. What it does suggest is that inequality does not monotonically decrease after a certain point on the development path. Rather, the industrialized countries have experienced a recent surge in income dispersion. Instead of the simple inverted-u pattern, recent inequality trends suggest a more complex tilde-pattern.

This paper is organized as follows. Section 2 reviews the empirical specifications applied in the literature that used the DS dataset and the suggested inequality determinants used. In the following section we replicate these specifications using an expanded and improved dataset. In addition, we perform some robustness tests, modify the basic specification to deal with the fact that measurement errors are closely associated with income levels in this kind of data, and use the Atkinson index to corroborate the results found with the Gini coefficient. Section 4 analyzes the recent inequality trends in highly industrialized countries and how it affects the Kuznets hypothesis. We conclude in Section 5.

6.2 Empirical review

The first tests using the new DS dataset found no support for the unconditional Kuznets curve (Deininger and Squire, 1996 and 1998). However, these authors used a time-series of four observations (for four decades) to test the within-country evidence. Arguably, this may be too short a time-series to account for development processes. It is not surprising that Li *et al.* (1998) report that the unconditional relationship holds better in the cross-section test than in the within-country regressions.

The conditional version has been tested in Higgins and Williamson (1999). As additional explanatory variables they use cohort size (as a demographic variable) and the Sachs and Warner index (as a trade policy variable). They find strong support for the weak Kuznets hypothesis. The results of Barro (2000) confirm these findings. He adds schooling variables, dummies for Sub-Saharan Africa and Latin America and reports that the curve is a clear empirical regularity, across countries and over time.[6] He also reports that openness is a significant inequality determinant, but surprisingly his results contradict the standard trade theory conclusions. The Stolper-Samuelson theorem states that tariff reductions unambiguously benefit the owners of the country's abundant factor and hurt the scarce factor. Thus, tariff reductions in labor-abundant developing countries should reduce income inequality, with the reverse effect in capital-abundant rich countries (Stolper and Samuelson, 1941).

In the rest of this paper we test both versions of the Kuznets hypothesis, focusing on the distributional effects of trade openness. In the next subsection we start with the empirical specifications used in Higgins and Williamson, and Barro.

[6] Litwin (1998) does not find any evidence for the conditional version when she uses a sample with only developing countries. This can point to the fact that a whole world sample must be used in order to capture the entire curve.

6.2.1 Empirical specifications

There are several ways in which to specify the panel-data regressions and to order the observations. A common feature of cross-country inequality observations is its unbalanced nature, i.e. there are many countries with few scattered observations and some others with complete or almost complete time-series. The later is common for OECD countries, while the former is representative of developing countries. To avoid assigning a higher weight to countries with more inequality observations, we follow Higgins and Williamson (1999) and Barro (2000), and organize the data by decades. However both studies diverge in the way they estimate the decadal observations. Higgins and Williamson use decadal averages, while Barro uses values centered on 1960, 1970, 1980 and 1990. In what follows, we use both approaches.

Moreover, the basic specification has different regressions for each decade, instead of having a regression for each country varying in time, as is the common practice in panel-data estimations. Thus, we estimate a panel of four regressions, with each separate regression having a number of observations equal to the number of countries:

$$\begin{aligned}
G_{i_60} &= \alpha + \beta_1 (\log GDP)_{i_60} + \beta_2 (\log GDP)^2{}_{i_60} + \boldsymbol{\gamma}\mathbf{D}_{i_60} + \varepsilon_{i_60} \\
\vdots \quad & \quad \vdots \qquad \vdots \qquad\qquad\qquad \vdots \qquad\qquad\qquad \vdots \qquad \vdots \\
G_{i_90} &= \alpha + \beta_1 (\log GDP)_{i_90} + \beta_2 (\log GDP)^2{}_{i_90} + \boldsymbol{\gamma}\mathbf{D}_{i_90} + \varepsilon_{i_90}
\end{aligned} \tag{6.1}$$

where G_{i_d} is the Gini coefficient for country i in decade d. GDP is gross domestic product per capita[7] and $\mathbf{D}$ is the vector of inequality determinants, which can include trade openness indicators. Note that in this basic specification, the estimated coefficients are constant over decade and country.

Another divergence between both studies is the estimation technique employed. Higgins and Williamson use random-effects regressions, while Barro uses feasible generalized least squares (FGLS) that account for heteroskedasticity, and temporally and spatial cor-

[7] The use GDP per capita logs is a common practice to capture the differences in development levels and to isolate the fact that growth is a compound rate process. The income data is taken from the Penn World Tables, version 6.1.

related errors.[8] Nevertheless, Beck and Katz (1995) have shown that for samples typically used in social sciences (many sections with few time observations) FGLS yield a small gain in the estimation efficiency, but produce overconfident standard errors. Therefore, we use instead correlated panel corrected standard errors (PCSEs) to take account for the contemporaneous correlation of the standard errors between decades.[9] In addition, we later assume a specific form of heteroskedasticity and model it directly into the PCSEs estimations. Comparing the random-effects and PCSEs results, both generate identical coefficients estimates but different standard errors.

To sum up, we have four basic specifications, which vary by the way data is organized (year centered or decadal average) and the estimation method (random-effects and PC-SEs). In addition, we introduce other tests to the data as in Barro (2000). For instance, we assess if the estimated coefficients are stable over time, and we use fixed country effects to capture country-specific non-observable characteristics.

6.2.2 Inequality determinants

Besides the two income variables that assess the Kuznets hypothesis, we can include additional inequality determinants and check for the conditional Kuznets curve. Unfortunately, we lack a formal model to explain income inequality and to analyze the rich variety of interrelations between different economic variables and income inequality (Atkinson, 1997). Thus, we do not have a theoretical basis on which to choose inequality determinants, and we follow the literature and present some suggested indicators.

To obtain a consistent panel set, we estimate these inequality determinant variables using decadal averages and year centered values.[10]

[8] Barro denotes this estimation technique as Seemingly Unrelated Regressions (SUR), although it is formally not a SUR procedure.

[9] We employ a Breusch-Pagan LM test of the independence of the errors across panels and in all specifications we rejected the null hypothesis.

[10] In particular, we use only values that are not more than four year apart from the initial year of each decade.

6.2.2.1 Educational attainment

As in Barro (2000) we include the average years of schooling in the total population over age 15 at three different levels: primary, secondary and higher education.[11] We expect that the primary and secondary education coefficients are negative and that higher education has a positive impact on inequality.

6.2.2.2 Country dummies

We add three different country-specific dummy variables: socialist or ex-socialist countries, Latin American countries and Sub-Saharan African countries. It is expected that socialist countries have significant less inequality, given the strong redistributional policies practiced in these societies. On the other hand, the other two types of countries have above average inequality, but the specific reasons for this pattern are not easy to explain. Barro tries to explain their significance by introducing colonial heritage and religious affiliation.[12]

6.2.2.3 Demographic factors

Higgins and Williamson argue that cohort size is a fundamental inequality determinant. The basic idea is that fat cohorts are associated with lower earnings and when these cohorts are located in the middle of the demographic distribution it smooths the lifetime pattern of earnings. On the contrary, when fat cohorts are associated with the youngest population, then it tends to exacerbate the differences in earnings between different age groups and thus, increases inequality.

They proxy cohort size by the proportion of the adult population between 40 and 59 years and name this variable “mature”.[13]

[11] This data is taken from Barro and Lee (1994 and 2001).

[12] Nevertheless, he does not present these results and uses both dummies throughout his estimations.

[13] We estimate this variable with data from the United Nations (Population Division of the Department of Economic and Social Affairs of the United Nations Secretariat, World Population Prospects: The 2002 Revision, http://esa.un.org/unpp).

6.2.2.4 Democracy and institutional variables

We use a subjective measure of electoral rights (democracy index) from the Freedom House.[14] Barro also includes a subjective indicator measuring the maintenance of the rule of law, but the earliest value available is for 1982 and thus, we do not use it. The same limitation is present for the corruption index that is also compiled by Political Risk services.

Thus, we only use the democracy index and in general, we expect that more electoral freedom decreases inequality, once we have accounted for the presence of socialist countries.

6.2.2.5 Trade openness indicators

Finally, we introduce three different trade openness indicators. There are a number of important analytical and practical problems involved when measuring trade openness. As explained by Berg and Krueger (2003), the main concern is about policies that distort market allocations and there can be many instruments that can achieve this; among others, tariffs and non-tariffs barriers (NTBs), and discriminatory exchange rates. However, one must also keep in mind that there a possible trade policy substitutions that cannot be monitored by a single variable indicator.

Direct policy measures estimate average tariffs and NTBs and can be directly related to trade policies. However, there is no average tariffs data available for the four decades we are analyzing. In particular, UNCTAD and WTO data on tariffs is only available for the 1990s.[15] Another approach is to take into account the possible substitution between measures. Higgins and Williamson use the Sachs and Warner index to measure openness in this fashion.[16]

[14] Where the index varies from 1 (highest) to 7 (lowest degree of freedom).

[15] In Chapter 5 we use these tariff data to estimate the impact of trade openness on inequality for a cross-country sample centered around 1994.

[16] Sachs and Warner (1995) try to capture any substitution policies by looking at five different trade outcomes. They construct a dummy variable for openness based on those countries that pass each of the five different trade policy tests. If any of the previous criteria is not met, the economy is considered to be closed and the index is assigned a value of zero. This approach has several drawbacks (Rodríguez and Rodrik, 2000). First, being a dummy variable one cannot differentiate degrees of openness. Second, some of the index components do not relate directly to trade liberalization. For instance, Hanson and Harrison (1999) remark that the black market premium measures factors other than trade policy. There

Alternatively, openness can be measured using trade volumes, i.e. $\frac{(X+M)}{GDP}$. This is a useful indicator of trade outcomes, but not necessarily of trade policy (Dollar and Kraay, 2004). In fact, this particular measure can be influenced by a number of country characteristics that are not related to trade policy at all.

Therefore, Barro filters this measure to account for these country-specific characteristics.[17] In particular, the adjusted trade volume variable is filtered by country size, since it is generally accepted that larger countries trade relatively less and thus, an outcome-based openness indicator must take this into account. Country size is represented by the logs of population and land area and formally, we estimate the following panel system:

$$\begin{aligned} TV_{i\_60} &= \alpha + \beta_1 \log Pop_{i\_60} + \beta_2 \log Land_i + \varepsilon_{i\_60} \\ \vdots \quad &\quad \vdots \qquad \vdots \qquad\qquad\qquad \vdots \qquad\qquad \vdots \\ TV_{i\_90} &= \alpha + \beta_1 \log Pop_{i\_90} + \beta_2 \log Land_i + \varepsilon_{i\_90} \end{aligned} \tag{6.2}$$

where TV_{i_d} is the ratio of trade volume $(X+M)$ with respect to GDP in country i for decade d; Pop is population and $Land$ is the country's land area in square kilometers.[18]

In (6.2) we use coefficients for decades and countries, and both estimated coefficients are negative and statistically significant. However, we try several different specifications which do not alter significantly the results.[19] Finally, ε_{i_d} is the adjusted trade volume openness variable to be used in the inequality regressions.

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has also been criticism in the use of the marketing board component and the independence of the tariff and quota measures with regard to the aggregated measure.

[17] However, Berg and Krueger (2003) argue that the empirical models used in these regressions are still not sufficiently accurate to identify the residual with policy.

[18] The trade volume and population data is taken from PWT (version 6.1) and the land data from the WDI.

[19] The system was both estimated using OLS and PCSEs, and with common and decade specific population coefficients, as well as with decade specific constant terms. However, the results do not change significantly and the residuals are almost perfectly correlated for all specifications. Finally, we use the best fitted specifications, i.e., the LS and common coefficients results for the year-centered variables and an iterative GLS regression with correlated errors and heteroskedasticity for decade-specific population coefficients for the decadal-average variables.

Finally, we include an openness indicator that has not been used previously in the literature. In particular, we take the ratio of import duties (MD) to total imports and construct the following variable:

$$LMD = \log\left(1 + \frac{MD}{100}\right) \tag{6.3}$$

Using the World Development Indicators[20] we estimate this ratio for the last three decades.[21] This variable reports collected tariffs and is implicitly weighted by trade. Nevertheless, this is an imperfect indicator since it does not take into consideration that high tariffs may completely deter imports, and thus, no tariff is effectively collected.

Summarizing, there is no ideal openness indicator and each group of measures has its own problems and limitations (Edwards, 1998). A strategy to overcome this has been to test the robustness of the results using different indicators.[22] Nevertheless, the measurement error induced by the openness indicators can be significant and most be considered when conducting empirical research.

6.2.2.6 Distributional effects of trade

Standard trade theory derived from the Hecksher-Ohlin-Samuelson (HOS) model predicts that tariffs and other measures that hinder trade, have distributional effects that are conditional on the factor-abundance of each country. In particular, the Stolper-Samuelson theorem states that price changes produced by tariff reductions unambiguously benefit the owners of the country's abundant factor and hurts the owners of the scarce factor. Thus, tariff reductions in labor-abundant developing countries should reduce income inequality, with the reverse effect in capital-abundant rich countries. Further, if we assume that the relative scarcity of factors in the world is represented by a continuum of countries with different relative factor allocations, then the level of income can proxy for factor-scarcity. For instance, very poor countries are assumed to be mainly unskilled labor abundant and less poor countries have more skilled labor. Thus, the combination of an openness

[20] World Bank, CD version.

[21] There was no data for the 1960s. We use a simple chain calculation: import duties (as a percentage of total revenue) times total revenues (as a percentage of GDP) divided by total imports (as a percentage of GDP).

[22] Among others, Levine and Renelt (1992); Edwards (1997, 1998); and Greenaway *et al.* (1998).

indicator with the level of income of the economy can assess the particular distributional effect of trade for each country. From our panel system, we include a trade openness indicator in the following way:

$$\begin{aligned} G_{i\_d} &= \alpha + \beta_1 (\log GDP)_{i\_d} + \beta_2 (\log GDP)^2{}_{i\_d} + \boldsymbol{\gamma}\mathbf{D}_{i\_d} + \\ &\quad \delta_1 Open_{i\_d} + \delta_2 \left[Open_{i\_d} * (\log GDP)_{i\_60} \right] + \varepsilon_{i\_d} \end{aligned} \tag{6.4}$$

Using the estimated coefficients we can obtain the estimated marginal distributional effect of an increase in openness:

$$\frac{\partial G}{\partial Open} = \delta_1 + \delta_2 (\log GDP) \tag{6.5}$$

To be consistent with the Stolper-Samuelson generalization, we must have $\delta_1 < 0$ and $\delta_2 > 0$. In other words, poor countries that increase openness should experiment a reduction in inequality. Using the empirical specification outlined above and our new inequality dataset, we can test if the theorem is confirmed by the data.

6.3 Reassessment using the new inequality dataset

With the previously specified panel systems and variables, we reassessed the previous inequality results when using our new dataset. We first introduce this dataset and explain the main differences with DS and then we present the results of our estimations.

6.3.1 New inequality data[23]

There are substantial difficulties concerning the empirical measurement of inequality. The most notorious is the lack of an institution that can assure data quality and consistency over time and across countries. In other words, an equivalent to the United Nations System of National Accounts, which provides macroeconomic statistics that are constructed

[23] **This section is based on the work presented in Chapter 3.**

by national agencies and are reasonably consistent over time and countries. In the absence of such an institution, some organizations have constructed "secondary" datasets, of which the most notorious are the World Bank (DS), the World Income Inequality Database (WIID, 2000) and the Luxembourg Income Study (LIS). These datasets compile available national inequality statistics and perform quality assessments to all the data observations. This has been an important first step towards the creation of internationally comparable inequality time series. Nevertheless, important caveats remain, which have to be addressed.

Inequality observations can be separated in three distinctive characteristics: concept measured, reference unit and source. In general, most of the observations measure one of three main concepts (Expenditure, Gross and Net income), there are several reference unit types (e.g. person, household, household equivalent) and a time series can be made up of different sources. The approached used by DS was to freely mix different definitions in order to construct the longest possible time series.[24] The implicit differences are then corrected by the use of fixed adjustments or the inclusion of dummy variables as regressors. This practice has been used in the studies mentioned in the empirical review.

However, Atkinson and Brandolini (2001), using a sample of OECD countries, have forcefully demonstrated that such combination of different concepts and reference units can seriously alter the levels and inequality trends of the data. Moreover, the use of fixed adjustments does not render the series comparable in many cases and furthermore, can introduce additional noise in data already rich in measurement errors. In our own research we have found similar problems when a broader sample of countries is analyzed and also conclude that the use of the DS dataset is not recommendable.

To avoid the additional measurement errors introduced by the combination of different definitions, we use the comprehensive dataset compiled in the WIID[25] and we construct inequality series that use consistently the same concept and reference unit when possible or alternative definitions that have been proven to be comparable for the specific country in question. Strictly speaking, our dataset does not include "new" observations, but

[24] The do try to be consistent in the use of concepts when possible, but in many cases a particular time-series mixes different concepts.

[25] This secondary dataset was based on the DS dataset but has been updated and includes more inequality observations, even those initially rejected by the DS quality criteria.

arranges the existing data in a consistent way in order to reduce the measurement error induced by the combination of diverse inequality definitions. In doing so, we have created three main series, which differentiate themselves by the concept used: Gross-household, Net-household and Expenditure-person. Table 6-1 summarizes the characteristics of each inequality series.

Table 6-1: Characteristics of the three main inequality series

	Gross Income Household		Net Income Household		Expenditure Person		Total for the three series		DS-accept
	Gini	Share data	Gini	Share data	Gini	Share data	Gini	Share data	
BASIC series									
Countries	49	38	27	25	69	63	145	126	115
Observations	427	326	288	241	189	159	904	726	693
Average obs.	8.71	8.58	10.67	9.64	2.74	2.52	6.23	5.76	6.03
OECD countries	29%	39%	70%	72%	1%	2%	23%	27%	17%
EXTENDED series									
Countries	95	70	47	43	85	75	227	188	115
Observations	634	445	433	376	254	205	1321	1026	693
Average obs.	6.67	6.36	9.21	8.74	2.99	2.73	5.82	5.46	6.03
OECD countries	17%	23%	43%	44%	1%	1%	16%	19%	17%

There a two different types of series, regarding the level of comparability criteria used. In the "basic" series, we include observations for which we are confident about their comparability, with respect to each concept and reference unit; whereas the "extended" series only considers definitions over which we are less confident.[26] This approach diminishes the measurement error present in the inequality data. Nevertheless, we do not address all measurement issues and many of the problems embedded in inequality data that have been extensively addressed by Atkinson and Brandolini are still present. Therefore, the data most be considered with some reservations.

This said, we use the inequality series which is more similar to the DS dataset: the Gross-Household-extended series. This data, in relation to the DS-accept series is com-

[26] For example, the extended series for Gross-Household can include Gross-person observations or Monetary Gross Income-Household observations. We are uncertain if these different definitions can be compared, but the inclusion of these observations can increase the number of observations provided by the basic series.

parable in number of observations, average number of observations per country and the representativeness of OECD countries. However, the DS series has more countries and this implies a smaller number of observations in our panel set.[27]

Another important consideration related to inequality measurement, is that there are several indicators that measure inequality and there is no consensus in favor of any particular index.[28] Although the Gini coefficient is the most commonly used indicator, there are many more inequality indexes that can be employed. Alternatively, distribution share analysis (i.e. quintiles and deciles information) can also be used. The main difficulty involved is that an inequality index reduces the information provided by a whole distribution function to a single value. This simplification is helpful, but necessarily disregards information. There are several inequality indexes suggested in the literature, but a convenient choice is to select an index that complements the information provided by the widely used Gini coefficient. In particular, since this coefficient is more sensitive to transfers in the middle of the income distribution and less sensitive to transfers at the extremes, it is convenient to include an index that can solve for this limitation. Such an index is provided by Atkinson (1970), which has an inequality aversion measurement parameter that controls for the sensitivity of transfers at the extremes.

Therefore, in addition to correcting for the combination of different definitions in the DS dataset, we construct Atkinson indexes from the inequality share data. Since most of the sources only report Gini coefficients and grouped income shares, we must first estimate the income distribution from grouped data and then estimate the inequality indexes. Specifically, we use parametric estimation of Lorenz curves to obtain the entire income distribution, and use these estimates to calculate four different Atkinson indexes. With these alternative inequality indexes we conveniently complement the information provided by the Gini coefficient and can run the same panel estimations to check for the robustness of the results.

[27] It is important to note that Barro uses mainly the DS dataset, but also includes some inequality observations (mainly for the 1960s) which are not considered by DS because they lack a clear reference to the primary source. We go along with Barro and consider that the inclusion of these observations is important to enlarge the series, even if it increases the uncertainty about the quality of the data.

[28] A comprehensive survey of the topic can be found in Cowell (2000).

6.3.2 Replication results

We first analyze the results using the year centered and random-effects specification.[29] For this particular case, our sample size is of 82 countries, 15 of which are OECD countries and the number of countries per decade is 36, 40, 42 and 59 for the 1990s. Table 6-2 presents the results.

Table 6-2: Random-effects GLS estimations, years centered on 1960, 1970, 1980 and 1990

Dependent variable: Gini coefficient (Gross-household extended series)	(1)	(2)	(3)	(4)	(5)	(6)
log (GDP)	56.039	35.541	48.911	53.013	53.393	53.186
	[12.378]***	[9.825]***	[14.177]***	[11.915]***	[12.219]***	[12.417]***
log (GDP) squared	-3.529	-2.256	-2.878	-3.168	-3.191	-3.177
	[0.734]***	[0.584]***	[0.845]***	[0.706]***	[0.725]***	[0.740]***
Dummy: Socialist		-12.874	-14.302	-15.726	-15.778	-15.768
		[1.547]***	[2.906]***	[1.727]***	[1.768]***	[1.776]***
Dummy: Latin America		9.137	5.582	6.032	6.040	6.017
		[1.289]***	[1.489]***	[1.433]***	[1.438]***	[1.460]***
Dummy: Sub-Sah. Africa		8.675	8.839	5.945	5.951	5.931
		[2.079]***	[2.937]***	[2.197]***	[2.204]***	[2.220]***
Primary schooling			-0.527			
			[0.606]			
Secondary schooling			-1.472			
			[0.836]*			
Higher schooling			6.130			
			[3.686]*			
Democracy index			0.919	1.015	1.013	1.021
			[0.327]***	[0.319]***	[0.321]***	[0.332]***
Mature			-33.823	-34.252	-33.679	-34.262
			[17.289]*	[15.763]**	[16.269]**	[17.254]**
Adj. Trade vol. (ATV)					0.188	1.360
					[1.252]	[11.323]
ATV*log (GDP)						-0.143
						[1.377]
Observations	177	177	142	160	160	160
R-squared	0.193	0.625	0.678	0.684	0.684	0.684

Notes: Constant terms are not reported. Standard errors in brackets.
* significant at 10% ; ** significant at 5% ; *** significant at 1%

A first conclusion is that the income variables have the sign and significance expected from the Kuznets hypothesis. The coefficients are robust to the different specifications, although the levels vary in some cases. In particular, we find evidence of both the unconditional and conditional hypothesis. However, the unconditional version has a low

[29] Using a Hausman specification test, we cannot reject the hypothesis that the estimated coefficients are the same for the random-effects and fixed-effects models. However, given that our panels are representing decades and not countries, a fixed-effect model controls for differences in decades, not in countries. To check for country-specific fixed-effects we use a different specification below. In what follows, we use the random-effects, but the results do not change if a fixed-effects model is used.

explanatory power, which is much increased by the inclusion of the dummy variables, which have the expected sign and are all significant throughout.

The educational attainment variables are not significant, though they have the expected sign. Only higher schooling is positive and significant in some specifications, but it is not robust to the inclusion of other control variables. The democracy index and the demographic variable are also significant. It is important to note that an increase in electoral freedoms is associated with less inequality. On the other hand, a more mature population is related with less income dispersion, in accordance to the cohort size hypothesis.

Finally, the trade openness variables are not significant in any specification. The table reports only the adjusted trade volume variable, but the other two variables (Sachs and Warner index, and import duties as a ratio of total imports) are not significant either. This is true for both the cases were the variable appears alone or in conjunction with the log of GDP per capita.

Table 6-3: Linear regression, PCSEs, years centered on 1960, 1970, 1980 and 1990

Dependent variable: Gini coefficient (Gross-household extended series)						
	(1)	(2)	(3)	(4)	(5)	(6)
log (GDP)	56.039	35.541	48.911	53.013	53.393	53.186
	[14.920]***	[11.534]***	[15.349]***	[13.397]***	[13.665]***	[13.884]***
log (GDP) squared	-3.529	-2.256	-2.878	-3.168	-3.191	-3.177
	[0.887]***	[0.678]***	[0.902]***	[0.788]***	[0.806]***	[0.822]***
Dummy: Socialist		-12.874	-14.302	-15.726	-15.778	-15.768
		[1.590]***	[2.693]***	[1.543]***	[1.589]***	[1.588]***
Dummy: Latin America		9.137	5.582	6.032	6.040	6.017
		[1.659]***	[1.709]***	[1.618]***	[1.620]***	[1.624]***
Dummy: Sub-Sah. Africa		8.675	8.839	5.945	5.951	5.931
		[2.488]***	[3.167]***	[2.547]**	[2.546]**	[2.568]**
Primary schooling			-0.527			
			[0.685]			
Secondary schooling			-1.472			
			[0.887]*			
Higher schooling			6.130			
			[3.426]*			
Democracy index			0.919	1.015	1.013	1.021
			[0.356]***	[0.340]***	[0.341]***	[0.353]***
Mature			-33.823	-34.252	-33.679	-34.262
			[18.543]*	[16.889]**	[17.523]*	[18.096]*
Adj. Trade vol. (ATV)					0.188	1.360
					[1.449]	[11.951]
ATV*log (GDP)						-0.143
						[1.413]
Observations	177	177	142	160	160	160
R-squared	0.194	0.625	0.678	0.684	0.684	0.684

Notes: Constant terms are not reported. Standard errors in brackets.
* significant at 10% ; ** significant at 5% ; *** significant at 1%

In Table 6-3 we present the results that are fully comparable with the specification of Barro. Here we use again year centered on the beginning of each decade, but now we use PCSEs to take into account contemporaneous correlations of the errors between panels.

Again, we find significant and robust Kuznets variables, the dummy variables are all significant while the schooling variables are not. Furthermore, the openness indicators are once more not significant in any specification. The results do not change when the schooling variables are included with the openness indicators, as originally tested by Barro and where he found significant openness indicators that contradicted the Stolper-Samuelson results. We do not find any such evidence using our consistent inequality dataset. One possible explanation for the contrasting results can be the sample size and composition. We have fewer observations (142 against 214) and in average 15 countries less per decade.[30] Another possibility is that his results were a artifice of the DS inequality dataset he used.

We turn our attention now to the results when decadal averages are used (Table 6-4). Our sample size in this case is of 81 countries, 14 of which are OECD countries and the number of countries per decade varies from 43 (1960s) to 57 (1990s). A first difference from both data specifications is that the number of observations is increased when decadal averages are used.

We find again that the income variables are significant and robust to different specifications. Of the schooling variables, only higher education is significant, but only in certain specifications. The democracy index is also insignificant, while mature is significant in most of the specifications. In this Table we have included the Sachs and Warner index, which was the openness indicator used by Higgins and Williamson. Nevertheless, none of the three openness indicators are significant in any kind of specification. The last two columns use GDP per worker, instead of per capita and also include decadal dummies (not reported). These changes were added to compare directly with the estimations of Higgins and Williamson and it finds the same results as these authors: only the conditional hypothesis is valid, cohort size is an important inequality determinant and trade

[30] This sample differences are the consequence of Barro using Gini coefficients defined for other concepts than gross income (i.e. Net income and expenditure).

Table 6-4: Random-effects GLS estimations, decadal averages

Dependent variable: Gini coefficient (Gross-household extended series)								
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
log (GDP)	56.190	32.354	51.975	39.783	39.528	44.811	*17.300*	*36.172*
	[11.765]***	[8.712]***	[14.479]***	[9.236]***	[10.605]***	[14.540]***	*[13.066]*	*[11.079]****
log (GDP) squared	-3.505	-2.079	-3.112	-2.455	-2.434	-2.784	*-0.948*	*-2.000*
	[0.695]***	[0.517]***	[0.866]***	[0.549]***	[0.630]***	[0.911]***	*[0.706]*	*[0.604]****
Dummy: Socialist		-11.657	-10.232	-12.101	-10.181	-10.037		-12.719
		[1.268]***	[2.551]***	[1.368]***	[2.539]***	[2.560]***		[2.015]***
Dummy: Latin America		9.379	6.624	7.261	6.900	6.947		7.665
		[1.198]***	[1.453]***	[1.306]***	[1.598]***	[1.604]***		[1.892]***
Dummy: Sub-Sah. Africa		11.860	10.168	11.668	11.612	11.563		11.871
		[2.343]***	[2.824]***	[2.297]***	[2.551]***	[2.559]***		[2.713]***
Primary schooling			-0.837					
			[0.579]					
Secondary schooling			-1.046					
			[0.726]					
Higher schooling			8.226					
			[3.423]**					
Democracy index			0.372					
			[0.358]					
Mature			-29.124	-32.926	-41.179	-42.773		-38.450
			[18.691]	[14.594]**	[18.756]**	[19.039]**		[17.567]**
SW index					0.412	1.292		
					[1.595]	[2.426]		
SW * log (GDP)						-10.537		
						[20.621]		
Observations	193	193	154	184	152	152	187	178
R-squared	0.187	0.669	0.698	0.695	0.618	0.619	0.122	0.699

Notes: Constant terms are not reported. Standard errors in brackets.
* significant at 10% ; ** significant at 5% ; *** significant at 1%

openness is not significant (not reported). However, we follow the common practice and prefer per capita levels as an indicator of development levels.

When PCSEs are employed the basic results remain unchanged, and thus, we do not show these results. However, an important difference is that the mature variable is not robust to different specifications in this case. This is not an isolated case and in general, we found that the inequality determinants are not robust to the different specifications. Almost in all regressions they have the expected effect on income inequality, but their levels and significance vary. A possible explanation is the correlation of some of these determinants with the income variables, which can introduce multicollinearity in the equations.

Regressing the equation with the inequality determinants and without the dummy variables, we find that the explanatory power is similar, but lower than the specification where only dummy variables are used. Thus, these dummy variables can be acting as a composite variable that agglomerates different characteristics that affect inequality in different ways for each country. This simplest solution, therefore, is to include only the dummy variables, which increase the explanatory power of the regressions to levels that are not significantly changed by the inclusion of the other inequality determinants.

6.3.3 Time stability tests

In this section we test if the estimated coefficients are stable over time. Hence, we estimate individual coefficients for each decade and run a Wald test to check if the four decadal coefficients are equal. For the particular case where we test the stability of the income variables, the panel-data specification is given by:

$$
\begin{array}{cccccc}
G_{i\_60} = & \alpha + \beta_{1\_60} (\log GDP)_{i\_60} + \beta_{2\_60} (\log GDP)^2{}_{i\_60} + \boldsymbol{\gamma}\mathbf{D}_{i\_60} + \varepsilon_{i\_60} \\
\vdots & \vdots \quad \vdots \qquad \vdots \qquad \vdots \quad \vdots \\
G_{i\_90} = & \alpha + \beta_{1\_90} (\log GDP)_{i\_90} + \beta_{2\_90} (\log GDP)^2{}_{i\_90} + \boldsymbol{\gamma}\mathbf{D}_{i\_90} + \varepsilon_{i\_90}
\end{array} \qquad (6.6)
$$

For the cases where other inequality determinants are tested, then the vector $\boldsymbol{\gamma}$ has decade-specific terms and the β coefficients are stable over time.

Table 6-5 reports the time stability tests for the Kuznets hypothesis variables. For both specifications the Wald test of equal coefficients over decades is not rejected. This is also the case for the unconditional curve and when only the dummy variables are included (not reported).

The following table shows the same test when the Latin America dummy, mature and the democracy index coefficient vary for each decade (Table 6-6). We held constant the other two dummy variables because of sample problems.[31] Again, from the stability test we can conclude that these three variables are relatively stable over different decades.[32]

[31] The Sub-Saharan Africa dummy varies from 8 countries per decade to zero, while the socialist dummy has one or two observations for the first three decades and more than twenty for the 1990s. In both cases it does not make sense to compare the coefficients when the samples are so diverse between decades.

[32] Although in the last column the mature variable fails the Wald test.

Table 6-5: Time stability test, PCSEs: income variables

Dependent variable: Gini coefficient	year cent.	dec. average
log (GDP)		
β1_60	46.331	53.716
	[16.862]***	[13.746]***
β1_70	48.536	52.969
	[16.215]***	[13.298]***
β1_80	47.709	51.146
	[16.007]***	[12.971]***
β1_90	46.877	49.885
	[15.561]***	[12.725]***
log (GDPpw) squared		
β2_60	-2.648	-3.334
	[1.075]**	[0.866]***
β2_70	-2.92	-3.27
	[0.988]***	[0.806]***
β2_80	-2.853	-3.072
	[0.955]***	[0.766]***
β2_90	-2.733	-2.905
	[0.906]***	[0.740]***
Dummy: Socialist	-15.297	-11.765
	[1.796]***	[1.847]***
Dummy: Latin America	5.891	6.436
	[1.629]***	[1.616]***
Dummy: Sub-Sah. Africa	6.103	11.159
	[2.632]**	[2.752]***
Democracy index	0.992	
	[0.340]***	
Mature	-51.393	-45.759
	[20.563]**	[18.124]**
Observations	160	184
Chi2 test: β1's are equal	2.09	5.42
Prob > chi2	0.46	0.12
Chi2 test: β2's are equal	2.56	5.89
Prob > chi2	0.55	0.14

Notes: Constant terms are not reported. Standard errors in brackets.
* significant at 10% ; ** significant at 5% ; *** significant at 1%

6.3.4 Fixed country effects

To allow for unobserved country characteristics, we introduce a country specific constant in the panel system (6.1), so $\boldsymbol{\alpha}$ is then a vector of country-specific terms (α_i), which are constant over time. We use only countries with at least two observations, which do not have to be adjacent.[33] The dummy variables are dropped since they do not change over time. In general, this particular specification captures only the time-series variations within countries.

We do not report the results, but as expected, the income variables are not significant in this specification, although they remain with the same signs. Four decades is not a long

[33] This leaves a sample of 40 countries for the year centered series and 41 when decadal averages are used.

Table 6-6: Time stability test, PCSEs: other inequality determinants

Dependent variable: Gini coefficient		
	year cent.	dec. average
log (GDP)	48.738	40.74
	[14.612]***	[11.328]***
log (GDPpw) squared	-2.863	-2.496
	[0.869]***	[0.676]***
Dummy: Socialist	-15.065	-12.811
	[1.915]***	[1.876]***
Dummy: Sub-Sah. Africa	6.915	11.225
	[2.732]**	[2.738]***
Dummy: Latin America		
γ1_60	7.517	8.542
	[3.233]**	[2.716]***
γ1_70	5.391	6.252
	[2.559]**	[1.906]***
γ1_80	3.781	6.625
	[2.084]*	[1.967]***
γ1_90	5.991	5.558
	[1.881]***	[1.918]***
Democracy index		
γ2_60	0.578	
	[0.653]	
γ2_70	1.328	
	[0.522]**	
γ2_80	1.45	
	[0.525]***	
γ2_90	0.797	
	[0.399]**	
Mature		
γ3_60	-46.655	-40.44
	[18.738]**	[16.252]**
γ3_70	-54.981	-45.314
	[20.225]***	[17.223]***
γ3_80	-62.004	-47.375
	[21.927]***	[18.159]***
γ3_90	-51.049	-37.874
	[21.019]**	[17.858]**
Observations	160	184
Chi2 test: γ1's are equal	2.2	1.37
Prob > chi2	0.53	0.71
Chi2 test: γ2's are equal	2.51	
Prob > chi2	0.47	
Chi2 test: γ3's are equal	5.78	10.53
Prob > chi2	0.12	0.01

Notes: Constant terms are not reported. Standard errors in brackets.
* significant at 10% ; ** significant at 5% ; *** significant at 1%

enough time span to measure changes in development levels. Furthermore, the exclusion of between country effects, which accounts for much of the inequality differences, leaves little variance in the sample to be exploited.

On the other hand, the mature and democracy variables are neither significant. This can also be a direct consequence of the small change these variables experience over time in each country. Nevertheless, in this specification the education variables become significant. Secondary schooling has a negative and significant effect on inequality, while higher schooling is also significant but has a positive effect. Such results suggest that educational variables are important inequality determinants to explain within country inequality over time. Furthermore, when we conducted time stability tests including schooling variables, only higher schooling had significant decadal coefficients, but only for the last two decades.[34] These results are consistent with the explanation that risen wage inequality in the USA and the UK –and increased unskilled unemployment in other OECD countries– is associated with skilled biased technological change, and that such an effect has occur after the 1980s. Nevertheless, this remark is qualified by the empirical use of overall income inequality, instead of wage inequality. Even when the use of gross income as the measured concept, however, is the closest possible definition to wage divergence, we still have to assume that both types of inequality are closely correlated.

6.3.5 Heteroskedasticity corrected regressions

An important characteristic of cross-country data is that richer countries tend to produce better statistics than poorer ones. Hence, measurement error can be associated with the level of development. Using this insight, one can correct for heteroskedasticity by weighting the error term by GDP per capita levels.

Table 6-7 presents the results of such weighting in the PCSEs regressions when the data is centered on the initial year of the decade. A first observation is that the unconditional Kuznets curve disappears. Nevertheless, the conditional version is still present, although is not as robust as before. In addition, schooling variables are now significant and with the expected effects: secondary schooling reduces inequality, while higher education enhances it. The results from the democracy index and the mature variable do not change from previous specifications. Finally, we find once more that trade openness has little explanatory power over inequality. In the table we show the results for the adjusted trade volume indicator, but in general, the other two openness measures yield the same outcome.

In the last column we present the only significant result when trade openness is included. The import duty variable is significant and positive, only when included alone (without the combined effect of GDP per capita). Thus, there is a monotonic relationship between import duties and inequality: higher duties associated with higher income inequality levels. This result can be explained by a strong negative correlation between this indicator and the income variables. For instance, developed countries are less reliant

[34] The low significance in the first two decades is the reason the coefficient is not significant overall.

Table 6-7: Weighted linear regression, PCSEs, year centered on 1960, 1970, 1980 and 1990

Dependent variable: Gini coefficient (Gross-household extended series)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
log (GDP)	22.687	20.398	41.911	32.853	33.326	32.491	20.698
	[20.051]	[13.904]	[17.200]**	[16.301]**	[17.267]*	[17.449]*	[17.544]
log (GDP) squared	-1.617	-1.370	-2.489	-2.000	-2.030	-1.976	-1.184
	[1.134]	[0.785]*	[0.976]**	[0.930]**	[0.997]**	[1.010]*	[0.997]
Dummy: Socialist		-13.199	-12.102	-14.433	-14.446	-14.443	-13.477
		[1.485]***	[2.092]***	[1.666]***	[1.671]***	[1.667]***	[1.684]***
Dummy: Latin America		9.982	6.108	6.755	6.786	6.748	5.739
		[1.558]***	[1.678]***	[1.683]***	[1.709]***	[1.706]***	[1.712]***
Dummy: Sub-Sah. Africa		11.746	12.558	11.790	11.853	11.738	16.534
		[2.893]***	[3.725]***	[3.709]***	[3.785]***	[3.804]***	[3.801]***
Primary schooling			-1.029				
			[0.547]*				
Secondary schooling			-1.394	-1.433	-1.426	-1.449	-1.277
			[0.621]**	[0.637]**	[0.645]**	[0.651]**	[0.631]**
Higher schooling			8.573	6.555	6.602	6.541	4.813
			[2.288]***	[2.140]***	[2.175]***	[2.165]***	[2.086]**
Democracy index			0.699	0.778	0.768	0.788	0.843
			[0.361]*	[0.369]**	[0.387]**	[0.391]**	[0.369]**
Mature			-28.336	-31.576	-30.897	-32.241	-45.959
			[14.296]**	[14.312]**	[15.987]*	[16.266]**	[15.304]***
Adj. Trade vol. (ATV)					0.135	4.856	
					[1.484]	[13.987]	
ATV*log (GDP)						-0.523	
						[1.515]	
Log MD							28.611
							[14.054]**
Observations	177	177	142	143	143	143	105

Notes: Constant terms are not reported. Standard errors in brackets.
* significant at 10% ; ** significant at 5% ; *** significant at 1%

on import duties as a source of governmental revenue. In fact, for most OECD countries the import duties are zero.[35] Under these circumstances, where positive import duties are strongly associated with poorer countries, the results do not contradict the Stolper-Samuelson theorem. These results are in general robust to the use of decadal averages and the Atkinson index, as shown in other tables below.

Nevertheless, since the other two openness indicators do not pick the same effects, and in general, are insignificant for a majority of the specifications, we are skeptical of the direct effect of openness on inequality. In any case, the evidence suggests that the direction of the trade openness effect is the expected one under standard trade theory.

When we analyze the data organized by decadal averages (Table 6-8), the Kuznets hypothesis is less significant than before. The unconditional version is again not significant, and the conditional version disappears in some specifications. For instance, the

[35] This may also reflect that this specific openness measure only considers collected duties and excludes the impact of high tariffs and NTBs on import outcomes.

Table 6-8: Weighted linear regression, PCSEs, decadal averages

Dependent variable: Gini coefficient (Gross-household extended series)									
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
log (GDP)	8.765	2.444	31.963	41.335	41.835	42.398	6.608	3.907	-5.097
	[20.382]	[14.329]	[18.890]*	[18.871]**	[18.808]**	[18.715]**	[14.071]	[14.114]	[17.855]
log (GDP) squared	-0.797	-0.349	-1.953	-2.565	-2.617	-2.646	-0.596	-0.442	0.183
	[1.148]	[0.810]	[1.082]*	[1.072]**	[1.069]**	[1.064]**	[0.796]	[0.800]	[1.007]
Dummy: Socialist		-11.200	-9.527	-10.061	-10.195	-10.220	-11.440	-11.033	-9.425
		[1.735]***	[2.246]***	[2.229]***	[2.216]***	[2.211]***	[1.632]***	[1.633]***	[1.869]***
Dummy: Latin America		10.344	7.690	8.396	8.511	8.566	10.283	10.648	10.209
		[1.653]***	[1.720]***	[1.644]***	[1.635]***	[1.634]***	[1.587]***	[1.697]***	[1.893]***
Dummy: Sub-Sah. Africa		16.873	13.118	13.998	14.167	13.921	16.665	17.657	19.020
		[3.945]***	[4.291]***	[4.403]***	[4.361]***	[4.373]***	[3.789]***	[3.784]***	[4.917]***
Primary schooling			-1.059	-1.147	-1.068	-1.080			
			[0.529]**	[0.539]**	[0.539]**	[0.536]**			
Secondary schooling			-0.956						
			[0.580]*						
Higher schooling			10.097	9.948	10.107	10.158			
			[2.609]***	[2.554]***	[2.536]***	[2.523]***			
Democracy index			0.870	1.020	0.836	0.850			
			[0.402]**	[0.405]**	[0.423]**	[0.422]**			
Mature			-18.675						
			[15.845]						
Adj. Trade vol. (ATV)					1.573	12.082	3.719	-31.119	
					[1.477]	[21.059]	[1.299]***	[17.704]*	
ATV*log (GDP)						-1.098		3.613	
						[2.202]		[1.840]**	
Log MD									51.931
									[17.875]***
Observations	193	193	154	154	154	154	193	193	122

Notes: Constant terms are not reported. Standard errors in brackets.
* significant at 10% ; ** significant at 5% ; *** significant at 1%

conditional version is valid only when the democracy index is included. These results can be the consequence of multicollinearity between the inequality determinants variable. When only the dummy variables are included, the Kuznets curve is not significant.

Besides these results, we also have that the adjusted trade volume variable is significant for the first time. However, the outcome is not robust to different specifications, and neither when we use year centered data or the Atkinson index. If there is any effect at all, this variable has again the expected direction according to standard trade theory and increased openness levels reduces income inequality in poor countries and increases it in richer economies.

In general, the correction for heteroskedasticity changes some of the main outcomes. This contrasts with the results of Barro (2000) who does not report any significant change when the measurement error is corrected to account for income levels.

6.3.6 Atkinson index results

In this last subsection we introduce the Atkinson index, which complements the information provided by the Gini coefficient. Formally, the Atkinson index is a family of indicators that diverge on the value of the relative inequality aversion parameter θ.[36]

The higher θ the more weight is given to the extremes of the income distribution. However, the index is also more volatile to small changes on the extremes (especially when $\theta > 1$). Thus, the Atkinson indicator is usually estimated using values between 0.5 and 1, following the estimation of the associated relative risk aversion parameter from the macro literature. In this paper we use $\theta = 1$ throughout.[37]

Table 6-9: Linear regression, PCSEs, decadal averages

Dependent variable: Atkinson index (Gross-household extended series)										
	Unweighted regressions					**Weighted regressions**				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
log (GDP)	59.774	31.955	68.680	36.801	29.050	10.045	-7.585	30.990	6.608	-53.231
	[19.791]***	[14.300]**	[24.478]***	[15.894]**	[20.501]	[28.868]	[22.015]	[32.233]	[24.684]	[27.228]*
log (GDP) squared	-3.727	-2.023	-4.123	-2.345	-1.717	-0.914	0.233	-1.966	-0.635	2.925
	[1.155]***	[0.833]**	[1.422]***	[0.937]**	[1.187]	[1.608]	[1.233]	[1.817]	[1.402]	[1.522]*
Dummy: Socialist		-10.072	-11.784	-9.785	-10.471		-10.352	-11.119	-11.092	-8.061
		[2.166]***	[3.223]***	[2.417]***	[2.282]***		[2.682]***	[3.148]***	[3.028]***	[2.718]***
Dummy: Latin America		11.990	9.961	12.152	9.827		12.611	10.974	12.579	12.506
		[2.094]***	[2.297]***	[2.274]***	[1.963]***		[2.117]***	[2.509]***	[2.443]***	[2.225]***
Dummy: Sub-Sah. Africa		24.450	23.518	23.649	29.296		27.392	24.933	26.152	33.264
		[4.093]***	[4.749]***	[4.598]***	[5.144]***		[5.640]***	[6.191]***	[6.628]***	[6.971]***
Primary schooling			-1.186					-1.009		
			[0.793]					[0.692]		
Secondary schooling			-0.959					-1.080		
			[0.903]					[0.740]		
Higher schooling			11.798	4.296				11.970	5.203	
			[3.968]***	[3.824]				[3.399]***	[3.250]	
Democracy index			0.399					0.786		
			[0.596]					[0.627]		
Mature			8.335					10.255		
			[26.499]					[21.611]		
Log Md					68.100					77.367
					[20.130]***					[22.176]***
Observations	135	135	113	127	81	135	135	113	127	81
R-squared	0.199	0.644	0.672	0.614	0.737					

Notes: Constant terms are not reported. Standard errors in brackets. * significant at 10% ; ** significant at 5% ; *** significant at 1%

In Table 6-9 we present the results when PCSEs are used with both the unweighted and weighted regressions, and the data is organized by decadal averages. As observed,

[36] See Atkinson (1970).

[37] In a previous study, we found that higher values of theta are too volatile and $\theta = 0.5$ does not have much variance. On the contrary, with $\theta = 1$ we have a comparable index to the Gini with respect to levels and variance (Chapter 3).

the conditional and unconditional Kuznets curve is only significant when unweighted regressions are used. This result is similar to the one found for the Gini coefficient. On the other hand, other inequality determinants are in this case not significant. In particular, the mature variable even changes the direction of its effect on inequality. Higher schooling is the only variable that is significant, but only when combined with the other inequality determinants. On the other hand, we only find a direct relationship between import duties and inequality, while the other openness indicators are not significant (we do not report these results).

Table 6-10: Linear regression, PCSEs, years centered on 1960, 1970, 1980 and 1990

Dependent variable: Atkinson index (Gross-household extended series)										
	Unweighted regressions					Weighted regressions				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
log (GDP)	74.695	51.032	71.037	81.516	39.529	42.664	29.700	44.464	29.116	-9.521
	[19.163]***	[17.030]***	[22.842]***	[20.506]***	[19.644]**	[26.200]	[20.549]	[26.051]*	[22.520]	[24.765]
log (GDP) squared	-4.626	-3.128	-4.165	-4.865	-2.295	-2.812	-1.880	-2.665	-1.834	0.501
	[1.126]***	[0.990]***	[1.331]***	[1.190]***	[1.137]**	[1.473]*	[1.150]	[1.464]*	[1.277]	[1.377]
Dummy: Socialist		-12.291	-16.691	-15.956	-17.117		-13.592	-14.475	-14.735	-14.506
		[2.555]***	[3.735]***	[2.863]***	[2.581]***		[2.511]***	[2.998]***	[2.673]***	[2.418]***
Dummy: Latin America		10.847	7.303	8.310	8.509		12.223	8.751	11.569	11.420
		[2.174]***	[2.261]***	[2.070]***	[1.856]***		[1.979]***	[2.331]***	[2.118]***	[1.882]***
Dummy: Sub-Sah. Africa		12.143	10.814	7.847	23.516		17.890	16.691	19.046	24.770
		[3.957]***	[4.641]**	[3.943]**	[4.496]***		[4.838]***	[5.224]***	[5.325]***	[5.383]***
Primary schooling			-0.708					-0.684		
			[0.907]					[0.728]		
Secondary schooling			-1.543					-1.479	-1.333	
			[1.097]					[0.795]*	[0.855]	
Higher schooling			7.120					8.099	5.620	
			[4.319]*					[3.010]***	[3.032]*	
Democracy index			0.888	0.950	1.148			0.586		1.214
			[0.517]*	[0.540]*	[0.494]**			[0.516]		[0.522]**
Mature			-18.431					-20.564		
			[24.102]					[20.458]		
Log MD					51.127					50.741
					[17.280]***					[19.004]***
Observations	136	136	116	122	86	136	136	116	127	86
R-squared	0.228	0.525	0.624	0.570	0.746					

Notes: Constant terms are not reported. Standard errors in brackets. * significant at 10% ; ** significant at 5% ; *** significant at 1%

Finally, the results for the year centered data are shown in Table 6-10. As in the previous table, we find again that the income variables are not significant when weighted regressions are used and in general, most of the inequality determinants drop their significance. Import duties are again significant when combined with the democracy index (as was the case in Table 6-7) or when only the dummy variables are included. This specification also provides the best fit and is valid even when the errors are weighted by

GDP per capita. Nevertheless, the other openness indicators are again insignificant and the result can be a consequence of the characteristics of the collected duties variable and not of trade openness itself.

6.4 The OECD challenge to the Kuznets hypothesis

In general, we have found strong empirical support for the Kuznets curve. The only specification in which the results were not conclusive was when we weighted the regressions by income, reflecting the likely association of measurement errors and development levels. However, this apparent contradiction can be explained by the recent rise of inequality in OECD countries and the introduction of a cubic function to account for this event.

The inequality pattern of OECD countries in the last two decades has changed. As observed in other studies (Gottschalk and Smeeding, 2000; Atkinson, 2003; Chapter 2), there is an measurable increase in the inequality levels in rich countries that contrasts with the previous decreasing trend for the 1960s and 1970s. This has produced a U-pattern of inequality in OECD countries for the last four decades. When we run the regressions using only OECD countries the signs of the income variables reverse –indicating a U-pattern.

Table 6-11: OECD countries, PCSEs

	Gross Income				Net Income			
	Dec. Average		*Year centered*		*Dec. Average*		*Year centered*	
	Gini	Atkinson	Gini	Atkinson	Gini	Atkinson	Gini	Atkinson
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
log (GDP)	-94.528	-133.554	-20.995	30.439	-371.264	-327.488	-146.542	-118.159
	[126.743]	[181.526]	[69.086]	[97.245]	[86.401]***	[95.936]***	[79.181]*	[98.280]
log (GDP) squared	4.844	6.766	0.850	-1.977	19.094	16.786	7.385	5.913
	[6.598]	[9.440]	[3.679]	[5.163]	[4.529]***	[5.023]***	[4.208]*	[5.192]
Observations	54	51	47	44	55	49	47	42
R-squared	0.031	0.058	0.154	0.157	0.348	0.275	0.285	0.198

Notes: Constant terms are not reported. Standard errors in brackets.
* significant at 10%; ** significant at 5%; *** significant at 1%

In Table 6-11 we present these results. Although the coefficients are not significant when inequality indexes defined for gross income, they are significant when net income is used. Given the relevant redistribution policies that characterize most of the OECD countries, inequality based on net income is a better concept to measure income dispersion. Thus, we see that there has been a reversal of the inequality trend and this can be enough to explain our results when using weighted regressions.

Furthermore, this reversal of the effects of the income variables can be influencing our results, since the weighted regressions assign more importance to inequality observations from OECD countries. To assess this possibility we introduce a cubic income function and we present the results in Table 6-12.

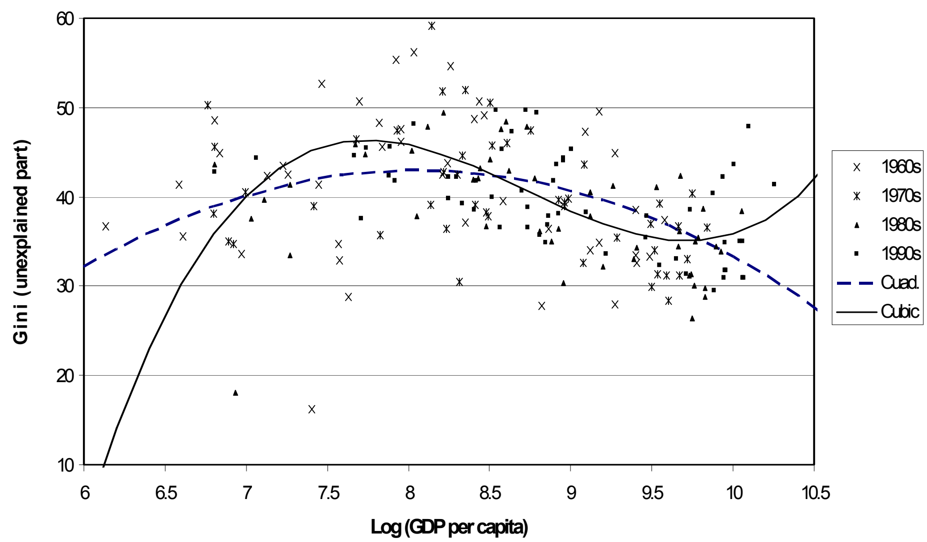
Table 6-12: Cubic income function using weighted linear regressions and PCSEs

	Gini Coefficient				**Atkinson index**			
	Dec. Average		*Year centered*		*Dec. Average*		*Year centered*	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
log (GDP)	1005.91	674.65	758.46	409.33	1036.12	670.97	944.62	587.55
	[221.48]***	[155.69]***	[214.46]***	[155.00]***	[316.78]***	[241.35]***	[289.68]***	[238.22]**
log (GDP) squared	-116.61	-78.26	-87.19	-46.52	-119.13	-77.81	-106.92	-66.09
	[25.74]***	[17.99]***	[24.97]***	[17.88]***	[36.46]***	[27.68]***	[33.43]***	[27.26]**
log (GDP) cubic	4.45	2.99	3.29	1.74	4.51	2.97	3.98	2.45
	[0.99]***	[0.69]***	[0.96]***	[0.68]**	[1.39]***	[1.05]***	[1.28]***	[1.04]**
Dummy: Socialist		-10.91		-12.85		-9.81		-13.05
		[1.61]***		[1.42]***		[2.51]***		[2.35]***
Dummy: Latin America		9.25		9.44		11.51		11.47
		[1.53]***		[1.50]***		[2.03]***		[1.92]***
Dummy: Sub-Sah. Africa		17.68		12.45		27.89		19.09
		[3.75]***		[2.86]***		[5.44]***		[4.74]***
Observations	193	193	177	177	135	135	136	136
Low income peak	2,368	2,320	2,380	2,508	2,506	2,362	2,747	2,993
High income hollow	16,243	16,516	19,377	23,087	17,801	16,190	21,835	21,925

Notes: Constant terms are not reported. Standard errors in brackets.
* significant at 10%; ** significant at 5%; *** significant at 1%

It is clear that the cubic function is present in our data. This specification is robust for all the different eight combinations of the conditional and unconditional curve using either data organized by decadal averages or centered on the first year of the decade, as well as for both inequality indexes. In addition, we report the two income level turning points for which the inequality trends change direction.

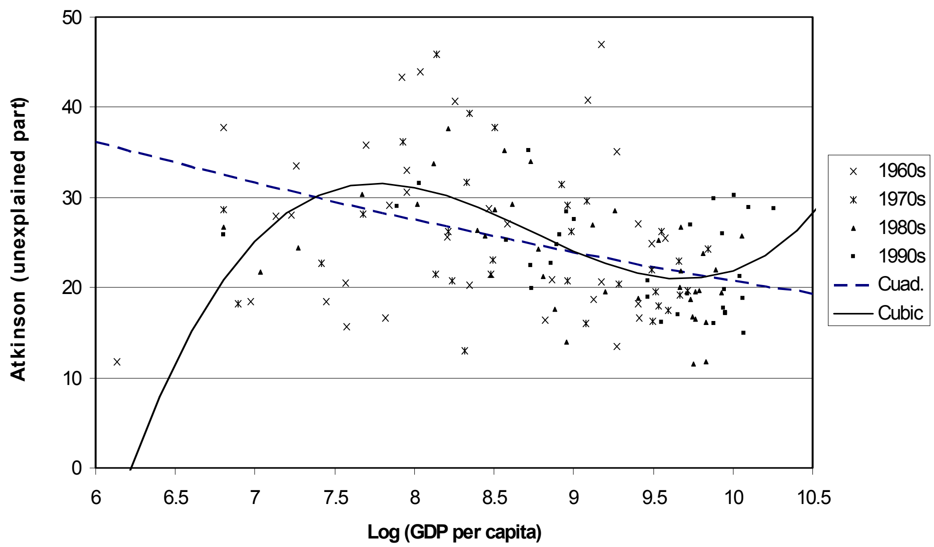
Figure 6-1: Gini coefficient and log(GDP per capita), cuadratic and cubic trends, decadal average data, weighted regressions and PCSEs,



In Figure 6-1 we plot the estimated quadratic and cubic functions with respect to the logs of GDP per capita and the unexplained part of the Gini coefficient when it is adjusted to account for the effect of the dummy variables. The cubic function itself does not contradict the Kuznets hypothesis, but adds an increasing trend after a certain high income threshold level is reached. Moreover, we organized the data by decade, and it can be observed that this threshold level has only been surpassed by some countries in the 1980s and especially in the 1990s. In fact, when we estimated the quadratic functions without the last decade, both income variables are again valid.

Finally, in Figure 6-2 we present the scatter plot when the Atkinson index is presented. For this case, we can draw the same conclusions as for the Gini coefficient. Note that both quadratic specifications have a bad fit into the data and for the Atkinson, we even have a slight U-pattern.

Figure 6-2: Atkinson index and log(GDP per capita), cuadratic and cubic trends, decadal average data, weighted regressions and PCSEs



6.5 Summary and conclusions

We use an improved inequality dataset that consistently uses the concept measured and thus, reduces the measurement error implicit in the widely used DS dataset. Following previous empirical specifications to test for inequality determinants (Higgins and Williamson, 1999; Barro, 2000) we use panel estimations for four decades and a representative sample of countries. We find clear support for the Kuznets hypothesis, for both the conditional and the unconditional version. Furthermore, this result is consistent over time, econometric techniques and two different ways to organize the data.

Additional inequality determinants also produce similar outcomes to previous studies. In particular, cohort size, democracy, schooling and dummy variables for socialist, Latin American and Sub-Saharan countries are generally significant and exhibit the expected effects. However, these determinants are not robust and they are correlated with the income variables, which creates possible multicollinearity problems when they are included in the regressions alone or together. A simple solution is to include only the dummy variables, which account for much of the increase of explanatory power characteristic of the conditional Kuznets version.

More importantly, we do not find evidence that trade openness influences inequality in a significant way. We use three different openness indicators and combine them with income levels to account for cross-country endowment differences. The Sachs and Warner index and the adjusted trade volume variable fail to report a measurable influence on the household distribution of income. Only the log of collected import duties has a direct and significant influence in some specifications, but the effect is not robust. If any, the distributional effects of more import duties present the expected direction predicted from the Stolper-Samuelson theorem. Another reason of this lack of influence, is that openness measurement is not adequately captured by the Sachs and Warner index, trade adjusted variables and the share of collected duties with respect to imports. Direct tariff indicators are only available for the 1990s and in a related study (Chapter 5) we used a cross-country estimation centered around 1994 and found that average tariffs significantly affect inequality as predicted by the Stolper-Samuelson theorem.

An improved feature of our econometric estimations is that we weight regressions to account for the fact that inequality statistics (and other statistics in general) are better assessed in developed countries. When the error terms are weighted by GDP per capita levels, the Kuznets hypothesis is somehow modified. The quadratic function is no longer valid, but instead, a cubic function is highly significant and robust to all specifications. This behavior can be explained by the recent inequality trend in OECD countries, where income dispersion has been increasing. Since the observations from these rich countries have more weight, it can account for a new inequality phenomena, that of a tilde-pattern where income inequality increases after a certain development level is attained.

These results do not contradict the original Kuznets effect, since for a range of incomes the curve is clearly discernible and statistically significant, but it may be a sign that highly-industrialized information driven economies may be experimenting new inequality consequences from changes in labor demand. In particular, this can be related to the surge of skilled-labor demand widely reported in the literature. On the other hand, the decline of the welfare state in some OECD countries can also be contributing to this inequality outcome.

Throughout the different empirical specifications, we employ the Atkinson index to assess the robustness of the results obtained from the Gini coefficient as an inequality indicator. In general, the results remain unchanged and this provides use with more confidence in our conclusions.

Finally, we concur with the previous empirical studies that used the DS dataset and find that the Kuznets hypothesis is a significant empirical regularity. However, the recent increase of income inequality in OECD countries may be challenging this view and in particular, points to the fact that a monotonic decrease in income dispersion is not a result of economic development and new labor demand forces can be reversing this pattern.

Chapter 7

Conclusions and Implications

The interaction between international trade and income inequality is an old topic of economic research, but the recent surge of economic globalization has revived academic and public interest. Concretely, the observation that wage dispersion in industrialized economies has increased in the last decades has spurred economic debate in OECD countries about the impact of trade liberalization. In addition, recent policy reforms in some developing countries have increased the interest on the possible effects of trade liberalization on the personal distribution of income.

The main purpose of this dissertation has been to analyze how trade openness affects income inequality from a global perspective. To perform this analysis we have assembled a new income inequality data, built a general equilibrium trade model that links tariff shifts with changes in household inequality, and tested the theoretical propositions derived from these models. In the following sections we analyze some of these issues in detail. First, we discuss different inequality empirical problems, diverse theoretical frameworks, the Stolper-Samuelson theorem and the Kuznets hypothesis. Furthermore, most of the chapters include important policy implications, and these are summarized in the last section.

7.1 Income inequality data and the reassessment of old empirical issues

The study of income inequality has been discouraged by serious data limitations. In general, personal income is not easily estimated and remains a sensitive issue when conducting household surveys. Moreover, such household income surveys are not regularly carried out in most countries and the comparability of observations between countries is exacerbated by the absence of an international institution that can guide and standardize the format of these surveys.

Under these circumstances, income inequality datasets are characterized by a rather heterogeneous compilation of observations. The intrinsic characteristics of each national survey, the concept measured (i.e. net income, gross income, expenditure), the reference unit (e.g. person, household, household equivalent) and the primary information source (national agencies or particular research studies) are only the main issues that can differentiate and qualify diverse inequality observations. Such disparities between household income surveys increase the measurement error intrinsic to personal income assessment.

In Chapter 3 we take on existing secondary inequality datasets and aim to reduce this particular type of measurement error. The changes we introduce are based on the recommendations of Atkinson and Brandolini (2001), who criticized previous datasets. In this chapter we conduct a comparability analysis between different grouping assumptions and conclude that it is not a sound practice to freely mix different concepts and reference units. Using this result, the new dataset we compile is characterized by different inequality-time series that diverge on the concept measured and the reference unit. In particular, we have three main series: Gross Income/Household, Net Income/Household and Expenditure/Person. Each series conveys complementary information that can be useful for different purposes. For instance, in Chapter 5 we use the Gross Income/Household series, which is an income concept that fits better in the analysis of the distributional effects of tariff protection, before any governmental redistribution policies are applied. When such governmental policies are to be considered the net income series is more relevant.

Another important distinction between the series is the country coverage. The expenditure series is almost entirely constituted by developing countries, while the net income series has a majority of OECD countries. Finally, the gross income series presents a balanced sample of countries and is comparable in this respect to the widely used secondary dataset compiled by the World Bank (Deininger and Squire, 1996).

A separate contribution of Chapter 3 is the use of parametric estimations of the Lorenz curve. For those inequality observations that provide sufficiently disaggregated information, we use seven different Lorenz curve models to estimate the best fit. Employing these estimates we obtain the whole income distribution, and with this information we calculate Atkinson indices of inequality and absolute poverty estimations.

Collapsing the information of a distribution into a single inequality index is a simplification that can overshadow particular inequality changes. While the Gini coefficient is the most widely used inequality measure, it assigns relatively more weight to changes in the middle of the distribution. A constant Gini coefficient over time can be ignoring important implicit inequality changes. Alternatively, the Atkinson family of indexes weighs more heavily the changes at the extremes of the distributions and thus, complements the information provided by the Gini coefficient. When we use both inequality indexes in empirical tests, we increase the robustness of the results.

The use of this improved inequality dataset is the cornerstone and main empirical contribution of the studies detailed in Chapters 5 and 6. As explained below, in Chapter 5 we test the main implication of the Stolper-Samuelson theorem when it is formally generalized to study household inequality. We find empirical support for the theorem's predictions and additionally, we can reconcile the apparent anomaly that middle-income countries have not experienced inequality improvements with increased trade openness. However, the lack of import tariff data limited the scope of the analysis and we were constrained to perform cross-country regressions centered in 1994. Although the disadvantages of these kinds of cross-country studies is widely recognized, we obtain a high degree of explanatory power when the error terms are associated with the per-capita income levels of each country. This econometric specification takes advantage of the fact that the quality of national statistics can be directly associated with the level of development of each country. In particular, we expect that rich countries produce higher quality observations than low-income countries, which cannot devote many resources to the compilation of national statistics.

In our econometric estimations we used the Kuznets hypothesis as an important inequality determinant to control for country differences. Chapter 6 is devoted to the empirical testing of this hypothesis. Kuznets associated inequality levels with per-capita income levels in the form of an inverted-u relationship. The most recent literature has documented the empirical validity of the Kuznets curve and it is regarded as a central inequality determinant. We confirm these previous findings when we use our new inequality dataset and when we control the error term for income levels. Furthermore, the results prevail when using the cross-country regressions of Chapter 5 or when we introduce panel estimations in Chapter 6. The interesting result of this chapter, however, is the emergence of a cubic relationship between income levels and income inequality when our panel regressions control for heteroskedasticity. This new pattern does not refute the Kuznets curve, but expands it to include the new trend of increased income dispersion experienced in highly industrialized economies. This observation confirms the recent observed pattern of increased personal and functional income inequality in OECD countries.

Finally, in Chapter 6 we also test the Stolper-Samuelson predictions using panel estimation. However, the lack of a long enough import tariff time-series and the statistically insignificance of the alternative openness measures we employ, provides an inconclusive test of the theorem. In this respect, the use of the new inequality dataset does not add any additional information to the existing literature. While some studies find either a positive or negative relationship, the bulk of the research points to an insignificant impact of trade liberalization on income inequality.

7.2 Income inequality, social welfare and trade

Income inequality is mostly associated with social justice and equity issues. Indeed, a vast literature has been devoted to study the social choice and ethical implications of income inequality.[1] However, while we generally regard inequality as an economic malaise, economists seldom consider it in social welfare estimations. As explained by Sen (1997) this a result of the preponderance of the Pareto condition in social choice theory. He argues that to make distributional judgments in welfare economics, the Pareto condition

[1] See for example Sen (1992, 1997).

must be excluded. That this condition is widely accepted and rarely questioned confirms the fact that economics has an almost single-minded concern with efficiency and is less concerned with distributional evaluations. Therefore, income inequality is often regarded as a necessary or acceptable drawback of economic policies aimed at increasing growth.

In Chapter 3 we move away from this standard approach and explicitly incorporate inequality considerations into social welfare estimations. Following the "veil of ignorance" insight first proposed by Harsanyi (1953, 1955), we can include distributional considerations into a Sen-type social welfare function. Formally, we directly associate income inequality measures with different social welfare functions.

We then embed this inequality indexes in a formal general equilibrium trade framework. Hence, we can trace the inequality and social welfare implications of diverse economic polices through the general equilibrium results. In particular, we map the distributional effects of import tariff changes. The use of an ownership matrix allows us to link the functional and the personal distribution of income. Standard trade models can be associated with specific ownership structure and thus, we obtain a series of distributional predictions from the Hecksher-Ohlin-Samuelson and the Ricardo-Viner models.

Within the HOS framework, we are able to generalize the Stolper-Samuelson results to the household distribution of income. In an economy with two production factors, one of them equally distributed (labor or unskilled labor) and the other concentrated (capital or skilled labor) the main predictions of the theorem still hold. The distributional effects of a Ricardo-Viner economy with an additional concentrated factor are, however, ambivalent. When the ownership matrix is better approximated by this three-factor approach we need more information about the distribution of the specific factors within the population to obtain less ambiguous results. As we already mentioned, the predictions obtained from both trade models are tested in Chapter 5 using the new income inequality dataset.

In Chapter 4 we analyze the social welfare implications of tariff protection. The insights obtained are then applied to study the theoretical setting of endogenous tariff formation. The efficiency properties of import protection are well studied in the literature. For instance, complete trade liberalization is the efficient trade policy for a small country. However, the widespread use of import protection instruments in many countries contradicts these theoretical predictions. One explanation can be that policy makers care about inequality-adjusted social welfare. In this context, the changes associated with import tariffs are more complex, i.e. the development level of each country is now relevant. In capital-abundant rich countries the social welfare distributional impact of a tariff increase is expected to be positive, while the inverse effect applies to labor-abundant poor countries. The distributional effects in middle-income countries that are not globally labor or capital-abundant are uncertain. Thus, in developed countries the distributional gains can compensate the negative efficiency losses associated with import protection and it is reasonable to assume that policy makers in these countries have incentives to maintain positive import tariff levels.

The favored explanation in the literature, however, is that policy makers do not only care about social welfare, but also about the political contributions of special interest groups (Grossman and Helpman, 2002). While the main forces behind the political support function remain, we complement the basic results of this political economy model with the inclusion of a inequality-adjusted social welfare function. The policy implications of this addition are explained in the corresponding section below.

7.3 The Stolper-Samuelson theorem

Despite being published more than 60 years ago, the Stolper-Samuelson theorem remains the central theoretical result linking trade with income distribution. In its original form, the theorem predicts the effect of tariff protection on the functional distribution of income. However, a generalization of its predictions to the personal distribution of income can be extracted using a rough simplification: poor countries are labor-abundant and labor is equally distributed, while rich countries are capital abundant and this production factor can be unequally distributed. In Chapter 4 we constructed a formal representation of this simplification, where we can relate tariff changes in a HOS environment directly to personal income inequality variations. In this context, the common perception based on the Stolper-Samuelson predictions is that trade improves income inequality in poor labor-abundant countries and deteriorates it in rich capital-abundant economies. Our literature survey illustrated how the theorem has been empirically tested with contradictory results.

Much of this dissertation has aimed at analyzing and expanding the theoretical implications of this theorem, while using a new income inequality dataset to conduct empirical tests. In Chapter 2 we have presented the main results of the trade and wages debate and in general, the basic consensus is that the Stolper-Samuelson predictions have been confirmed by the facts. However, the relatively low volume of trade between both groups of countries (north-south trade) can only partially explain the significant increase in wage inequality. Alternative explanations, such as skilled-biased technological change and outsourcing activities are regarded as more relevant determinants of the wage experience in OECD countries. Connecting these results with the observation presented in Chapters 3 and 6 that the personal distribution of income has also been deteriorating in highly-industrialized countries (mainly in the United States and the United Kingdom), one can generalize the conclusions of the wage debate to the household distribution of income. In particular, the empirical results of Chapter 5 using personal income inequality data support the Stolper-Samuelson predictions for capital-abundant countries.

The literature that analyzes the experience of developing countries is less conclusive. One possible explanation is the heterogeneity of economies that are bundled in this group. For instance, the endowment characteristics of middle-income and low-income countries are different. This fact can explain the apparent anomaly cited in the literature to refute the Stolper-Samuelson theorem. E.g., that the Southeast Asian experience of the 1960s and 1970s confirmed the predictions, while the Latin-America inequality experience of the 1980s and 1990s has contradicted it (Davis, 1996; Robbins, 1996a and 1996b). When assessed in relation to the whole world and not only with respect to richer countries, middle income countries (which characterize most of the Latin-American economies) are not labor-abundant. They can either be characterized as natural resource abundant or grouped in an intermediate diversification cone that is between the two polarized cones used in the original theorem. In Chapter 5 we view the world as a continuum of countries, each with different endowments that monotonically relate per-capita income levels with increased capital abundance. Within this context, the empirical results suggest that the main predictions of the Stolper-Samuelson theorem are supported when generalized to include the household distribution of income. Low-income countries benefit the most, in terms of inequality improvements, when import tariffs are reduced. The results for middle-income economies are uncertain and the richest countries present increased income inequality with trade liberalization.

Furthermore, the satisfactory goodness of fit of the regressions suggests that the HOS assumptions are a good approximation to real-world behavior. We reviewed in Chapter 2 the simplicity of these assumptions and how they have been constantly challenged in the literature. The strong empirical results support the use of this model, in particular, the small country assumption. Terms of trade effects are usually invoked to justify import protection; however, in an ever more integrated world economy these effects may have small aggregate long-run effects.

However, we can recall that the present surge of income dispersion in OECD countries has been associated with other economic events, such as the gradual dismantling of the welfare state (Atkinson, 2003) and skilled-biased technological change. Furthermore, rich countries possess capable social-safety networks that can deal with the potential income reduction of the bottom deciles of the distribution. It is precisely the lack of such functioning networks in low-income countries that emphasizes the negative income inequality and poverty effects of trade protection for this type of economies.

In Chapter 6 we use a panel estimations approach to assess the impact of trade policy on the household distribution. The inconvenience, however, is that we lack import tariff data for years before 1990. Thus, we are constrained to use other trade openness indicators during the four decade periods ranging from 1960 to 1990. In Chapter 2 we have explained the limitations of these alternative openness measures (i.e. Sachs and Warner index, adjusted volume of trade) and the expectation is that their combined use can increase the robustness of the estimations. Nevertheless, neither index has a significant impact on inequality under different specifications. We are uncertain if this outcome is a result of the intrinsic shortcomings of the indexes or if indeed, trade openness has a negligible impact on the personal income distribution. Thus, our findings add another uncertain answer to the existing contradictory and non-conclusive results present in the rest of the literature. This is not a direct contradiction of the Stolper-Samuelson theorem, but may be a result of other inequality determinants being more important and overshadowing the trade effects.

To sum up this section, our empirical results have mainly confirmed the effect of a tariff change predicted by the Stolper-Samuelson theorem. The use of alternative trade openness variables suggests that the distributional effects of trade are not significant.

7.4 The Kuznets curve

Another old economic idea that has survived until present time is the Kuznets hypothesis. In his seminal article, Kuznets (1955) observed that the distribution of income deteriorated in early stages of development and later began to improve after a certain threshold was crossed. The Kuznets hypothesis has been used in two different contexts.

First, it has been employed as a mechanism to explain dissimilar inequality levels between countries. As pointed in Chapter 3, within-country inequality is relatively stable in comparison to between-country income dispersion. When we introduced a quadratic income form in our inequality regressions in Chapters 5 and 6, these Kuznets variables are highly significant and robust to different econometric specification. Thus, income levels are an important and significant inequality determinant. These results confirm previous findings (Higgins and Williamson, 1999; Barro, 2000) pointing to the empirical relevance of the Kuznets curve and its stability over time.[2]

Secondly, the derived inverted-u relationship between per-capita income levels and inequality has been a central association often used to predict the distributional effects of growth. Low-income countries that experience accelerated growth are expected to encounter an increase in inequality until the threshold income level is attained. However, this initial trade-off between growth and inequality is not perceived as implicitly negative, but as a direct consequence of some specific sectors or economic groups taking off in the development process. Once this economic modernization encompasses most of society and public safety-nets are installed, additional growth spells are expected to decrease income inequality.

[2] We already mentioned that in Chapter 6 we found a cubic relation between per-capita income levels and inequality. This result does not contradict the Kuznets hypothesis, but points to a new phenomenon of increased inequality in highly industrialized countries. A possible explanation of this new development is the influence of skill-biased technological change on the functional distribution of income.

Despite its simplicity and empirical strength, there are some drawbacks to the Kuznets curve. Theoretically, it is not a formally constructed model based on solid microeconomic foundations. In its original unconditional version, shifts in labor demand are implicitly used to explain the inequality pattern. The conditional version includes additional variables –such as schooling, demographic factors and trade– that can affect labor demand and indirectly, income distribution. However, other factors may be at work and the income variables can then be only proxying these underlying forces.

Most notably, political and institutional variables are absent from the analysis. The government has a decisive role in the redistribution of income. It controls a wide range of instruments that can alter the final income of each household (e.g. taxes, transfers, privatizations, nationalizations, subsidies and import tariffs). In addition, political and social institutions may also have a significant influence on how income is distributed. Among others, the rule of law, corruption levels, unions bargaining power and labor market legislation. However, it is not far-fetched to think that these political and social variables are correlated with income levels and thus, the Kuznets hypothesis creates the right predictions albeit using an alternative explanation.

The Kuznets curve has also an important empirical downside. Assuming that economic development is a long run process, the hypothesis has to be tested using the evolution of inequality in a single country for a long period of time. We lack such lengthy inequality time-series and we have to use panel data estimations as an empirical substitution. Cross-country regressions have been widely criticized because of unobservable differences between countries. The use of pooled data can usually correct for this problem, but we cannot be certain that the unobservable variables are not also changing over time.

In fact, the statistical significance of three dummy variables in the regressions of Chapter 6 highlights both the theoretical and empirical weaknesses. The socialist country dummy asserts that income inequality is significantly lower in these countries when income levels are controlled for. However, we can explain this result given the preponderant role that the government was given in these societies and the ideological bias it had towards equity issues. The other two dummy variables are less easily explained. Latin-American and Sub-Saharan countries have above average inequality levels after income levels are considered. Barro (2000) has used additional explanatory variables to elucidate this effect, but without results. Therefore, there are some underlying variables that significantly affect inequality levels of which we cannot provide details; more than the fact that they are geographically concentrated in these two regions.

In summary, the Kuznets curve is a useful device to understand inequality levels and to think on how inequality changes over time. However, the lack of a solid theoretical basis about how inequality is determined and how it interacts with growth (Atkinson, 1970), limits its use for policy recommendations. In particular, the recent surge of inequality in OECD countries and the experience of Latin-American and Sub-Saharan countries are clear examples of these limitations.

7.5 Policy Implications

The wide concern on how globalization affects social and economic conditions is a broad question that we have partially analyzed in this research study. Although we have limited our attention to the effects of trade liberalization and have not covered an inclusive concept of globalization, some important policy implications can be drawn.

It is important to bear in mind that we are analyzing the marginal distributional effects of trade. While diverse inequality determinants have been suggested, we have already explained that per-capita income levels are the preferred inequality determinant used in the literature. Even when the Kuznets hypothesis can be proxying for other variables, trade is not regarded as a significant factor altering the level of income inequality. Thus, we are not considering how trade protection establishes inequality levels, but rather how changes in trade policy affect the existing distribution of income, which is mainly determined by other factors.

Moreover, there are multiple instruments a government can use to modify existing trade flows, e.g. tariffs, non-tariffs barriers (NTBs), discriminatory exchange rates and more recently, contingent protection and sanitary standards. Generally, several instruments are used at the same time, but they can also act as substitutes. These policy conditions make it extremely hard to measure trade openness and liberalization programs. Several measures have been proposed in the literature, with more or less success.

However, the Stolper-Samuelson predictions were originally designed to assess the impact on factor wages driven by a change in final prices. When international prices are normalized and transport costs ignored, domestic price changes are induced by an import tariff increase. Hence, the use of weighted average tariffs is the most consistent empirical setting to test the theorem's predictions. In Chapter 5 we have shown that the cross-country empirical evidence for 1994 clearly supports the Stolper-Samuelson theorem.

The policy implications of this outcome are clear. The poorer the country, the more incentives it has to reduce tariff protection. These incentives are reduced as per-capita income levels increase and for some middle-income countries it may prove that import protection does not have any positive distributional effects. Thus, import protection negatively affects the income of the poorer members in low-income countries. This negative outcome is reinforced by the inadequate public safety-nets present in these countries, which cannot compensate for income losses of its poorer members.

Positive tariffs under these circumstances can then be explained by endogenous tariff formation models. The influence of special interest groups is well researched in the literature (Grossman and Helpman, 2002). These groups lobby for tariff levels that benefit them directly and the policy makers subsides to their demands in relation to the weight he assigns to political contributions and the amount offered by the interest group. In Chapter 4 we have analyzed how these basic results can change when the policy maker includes equity issues in his objective function using an inequality-adjusted social welfare function. We found in this setting that in rich capital-abundant countries the positive distributional impact of tariffs increases the incentives to raise tariffs. On the contrary, in labor-abundant poor countries tariffs affect negatively social welfare, both by efficiency and distributional losses. In these countries, positive tariffs are explained by the lobby efforts of capital owners.

When broader trade openness indicators are used, the results are less informative. In our panel estimations of Chapter 6, we do not find any significant relationship between income inequality and the Sachs and Warner index or the adjusted volume of trade indicator. These openness indicators contain a wider set of information and this can be concealing the direct distributional influence of trade liberalization. Another possibility is that the marginal effect of trade is negligible and other inequality determinants are more relevant to explain distributional changes. However, we do not find any support for the hypothesis that trade deteriorates the personal distribution of income.

In summary, we found theoretical and empirical evidence supporting the hypothesis that trade liberalization benefits the income distribution of low-income countries. Data constraints have limited our analysis and thus, these results are not conclusive. However, we are more confident about the result that the marginal effect of trade openness is small and does not influence negatively income inequality in these countries. If one believes that trade has a positive impact on growth then trade liberalization can be regarded as a positive policy for developing countries.

For industrialized economies, the positive impact of tariff protection is also empirically supported. This confirms the main results of the trade and wages debate, which assign a positive but minor role to trade in the recent spell of wage and personal income dispersion. In these countries, however, advanced social security networks can better cope with income losses in poorer households, while the country benefits from improved efficiency through increased international trade.

Notwithstanding, we lack an integrated theory that can deal with the interaction between income distribution and other relevant economic variables. Within these limitations, the remarks produced in this dissertation apply to general trends observed in a cross-country sample. Income inequality determination and its dynamics are a theoretical area under development and the general analysis we have conducted here cannot substitute for comprehensive country-specific studies that integrate historical, political and social factors.

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