The relationship between economic growth and inequality: evidence from the age of market liberalism

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Abstract

Using a panel data set of selected countries, this paper shows that the inequality-growth relationship follows an ordinary-U curve during the period 1970-98, in which inequality first decreases and then increases with economic growth. In addition, there is some evidence that the increasing pattern may reverse at higher levels of income. A time-series approach shows that a substantial group of countries capture a minimum turning point in different years along the period and others follow a permanent positive trend. It also indicates that only a few countries reverse inequality in a latter stage and display a maximum turning point after the mid 1990s; these countries are associated with macroeconomic stability, high governance and moderate expansion of trade and FDI. Hence, the inequality-growth relationship during the era of market openness has tended to change towards a positive one, although it might reverse at a later stage.

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1. Introduction

In the post-war period, during the 1950s and 1960s, at a time of full employment and rapid growth, the distribution of income was not a major topic of discussion. However, there has emerged a renewed interest in this subject over recent years, owing to prolonged unemployment and unstable and slow economic growth on average during the last quarter of the twentieth century. The implementation of market oriented policies at a global scale since the late 1970s and the need to assess the performance of these policies is another aspect that has fostered renewed interest in the study of income distribution.

The analysis of the relationship between growth and inequality is one of the recent routes that have been followed to study the evolution of distribution. This analysis has not only revived old issues such as the Kuznets' inverted-U hypothesis (1955), but has also contributed to recent discussions like the pattern of inequality during the age of market liberalism. This paper concentrates on the former issue as it will be looking at the inequality-growth relationship over the last few decades.

Some studies have derived empirical support for an inverted-U curve using cross-country evidence in the absence of adequate longitudinal data on distribution (Bourguignon, 1994; Milanovic, 1995; Jha, 1996). However, it has been contended that this approach does not render appropriate conclusions as it does not deal with intertemporal relationships (Deininger and Squire, 1998: 276; De Gregorio and Lee 2002: 404). More recent studies have adopted a panel data approach by using the Deininger and Squire (1996) (D & S hereafter) data set and have obtained different forms of the inequality-growth relationship (Ram, 1997; Barro, 2000; Forbes, 2000). However, the D & S data set has been criticised for not generating an accurate outcome since many of its observations are not consistent and comparable, even after applying “high quality filters”, and because its coverage is limited and unbalanced (Atkinson et al., 2001; Galbraith and Kum, 2002).

Panel data analysis could also be undertaken by means of two additional sources available in the literature - the Luxembourg Income Studies (LIS) and the UTIP-UNIDO data sets. The former overcomes many of the problems of heterogeneity, since it is assembled from micro-level data, but its coverage is restricted mainly to a few wealthy countries in recent years, making it inappropriate for a global study of the inequality-growth relationship over the last decades. The latter comprises a large coverage, but it is assembled from industrial pay inequality, which is just a component of overall income inequality.

With the above in mind, for this study we use the Estimated Household Income Inequality (*EHII*) data set constructed by Galbraith and Kum (2003). It takes advantage of accurate observations in D & S and the information in the UTIP–UNIDO in order to replicate the coverage of the latter with estimated measures of household income inequality taking in to account the relationship between industrial pay inequality, household income inequality, and an additional set of variables. The result is a data set with large coverage that overcomes inconsistencies in D & S.

After assembling the variable on inequality and the variable on income it is possible to construct an unbalanced panel consisting of 116 countries and 2,289 observations over the period 1970-98. Moreover, the coverage of the data also allows us to construct a balanced panel consisting of 31 countries and 899 observations over the same period. We use both samples in order to test if gaps within the data can create any source of bias. So as to estimate the model consistently and efficiently we use GMM estimation for dynamic panel data models proposed by Blundell and Bond (1998).

The literature has conventionally applied quadratic equations. Although we follow this approach, we also extend the model in to a third degree polynomial to test the possibility that the inequality-growth relationship could be better described in terms of cycles along a process of adjustment toward a more globally competitive environment, as suggested by Jacobsen and Giles (1998). Simultaneously, orthogonal transformations are applied to reduce the degree of multicollinearity that characterises polynomial equations.

In the literature dealing with the evolution of income distribution it has been recently emphasised that further intertemporal evidence should ideally be based on time-series analysis from single countries (Bruno et al., 1998; Morrison, 2000). In this respect, Atkinson et al. (2001) state that increasingly economists are focussing attention on the long-run trend in income inequality and highlight the importance of time-series for this matter. They also contend that the increasing availability of estimates that range from 20 to 40 years in many nations is making it possible to examine long periods of distributional change through a time-series approach. Our data set allows us to conduct time-series analysis for 31 countries along 29 continuous estimations. This sample is obtained by splitting the balanced panel. This approach complements evidence obtained from the panel data analysis and enables us to date distributional changes across countries over the period.

In the time-series analysis linear and quadratic trends are explored and the model is also extended into a third degree polynomial to test the existence of any cyclical pattern while the problem of multicollinearity is addressed by using centered data. Some studies that have reported turning points in the trend of inequality have not addressed the issue of non-stationarity of the variables and have not tested for the presence of cointegrating regressions (Ram, 1993; Hsing and Smyth, 1994). In this sense, Jacobsen and Giles (1998: 408) highlight the adverse implications of modelling with non-stationary data, as this omission casts grave doubts on the reliability of the findings to date. In this study we address the issues of stationarity and cointegration. In addition, the existence of autocorrelation in the error term is also explored.

The panel data analysis shows an overall U-shaped relationship between inequality and growth at different levels of development and gives weak evidence of the presence of a local maximum over the long-run. The time-series analysis shows diverse patterns but in general illustrates that the majority of countries capture a minimum turning point in different years along the whole period and other countries show a permanent upward trend, only a few economies display a negative trend or no systematic relationship. Furthermore, the time-series approach reveals that rising inequality is likely to reverse at higher levels of per capita GDP as a few countries achieve a maximum turning point after the mid 1990s. It is worth noting that these countries are associated with macroeconomic stability, high governance, moderate expansion of trade and FDI and their period of increasing inequality starts earlier on average than the rest of the countries.

According to the theoretical foundations supporting the surge of market-oriented strategies in a global scope since the late 1970s, it was expected that income distribution would improve with economic growth. However, our findings do not support this view and are rather in keeping with recent studies indicating that inequality has tended to increase in many countries since the 1980s (Morrison, 2000; Gottschalk and Smeeding, 2000; Flemming and Micklewright, 2000; Atkinson and Bourguignon, 2000; Smeeding, 2002; Galbraith and Kum, 2002; Galbraith and Kum, 2003). On the other hand, the neoliberal view contends that inequality may begin to lessen over the long-run once the market forces react, and our findings partially seem to support this assertion.

The paper is organised as follows. Section 2 discusses neoliberal assumptions, its theoretical foundations and expectations. Section 3 provides a preliminary analysis of the evolution of growth, income distribution and the relationship between these two variables since 1970. In sections 4 and 5 the panel data analysis and the time-series analysis are undertaken respectively. The interpretation and discussion of results are presented in section 6. Finally concluding remarks are provided in section 7.

2. Assumptions, theoretical foundations and expectations

The neoliberal thesis claims to have theoretical support to offer countries improvements in income distribution for two main reasons. Firstly, it boosts exports, employment and output, and therefore provides additional resources that facilitate the distribution of income. Secondly, it facilitates the operation of market forces and the mechanism of prices which allows resources to be allocated more efficiently. The policy prescription recommended to achieve these goals can be summarised as liberalisation of trade, investments and the labour market; privatisation and fiscal discipline.

The cornerstone of this economic model is provided by trade openness and the theoretical pillar of this policy edifice is familiar from neoclassical trade theory (Corden, 1993). In terms of economic growth, trade liberalisation provides access to imported capital goods on more favourable terms that foster technological modernisation and productivity, and therefore expand output. In addition, this policy is assumed to give an ambiguous boost to exportables which reinforces export-led growth, while trade balances through a variable exchange rate. In particular, the theoretical foundation supporting distributional effects of trade is the Stolper-Samuelson theorem (FitzGerald, 1996: 32). Within this two-factor neoclassical model, liberalisation of foreign trade increases the use of the cheaper-abundant factor as exports and imports adjust according to the orthodox principle of comparative advantages, while the costly-scare factor is used less. This mechanism increases the income of the factor which is relatively most used in the export sector and which is also more abundant. This factor is conventionally assumed to be unskilled labour in developing countries; by the same token, income distribution tends to improve.

The opening of the capital account, accompanied by capital market liberalisation and the process of privatisation is expected to create preconditions for large capital flows from abroad. Efficiency is expected to be boosted by the transfer of technology and management know-how, which usually accompanies FDI. Moreover, such foreign flows are seen as mobilising external savings, which supplement domestic savings, and therefore raise investment and boost growth (Griffith-Jones, 1996: 127). The stimulus to exports is expected to increase supply and further investment in the trade sector. In addition, foreign investment emerges as a source of finance, while the share of commercial bank lending tends to fall, this pattern opens the possibility to allocate more resources to both government and private investment. In this sense, larger rates of investment encourage the expansion of exports and output. Consequently, capital account liberalisation also emphasise outward-looking growth. Furthermore, the distributional effect of foreign investment is caused by the expected flow of capital to the production of tradable goods that mainly uses the cheaper-abundant factor of the economy.

Labour market liberalisation is aimed at maintaining labour market flexibility by limiting union power and allowing wages and supply to respond flexibly to market signals (Barrett, 2001: 563-4). Under these circumstances, labour market liberalisation is intended to reduce market distortions, in order to lower the cost of labour and to encourage both competitiveness and employment, which benefits economic growth. The labour market, under conditions of liberalisation, is expected to adjust according to the principle of comparative advantage. In this sense, labour market liberalisation is linked to trade and capital account liberalisation because it also stimulates the production of tradable goods, which increases employment and wages in the exportable sector, and hence redistributes income.

As inflation is deemed to introduce distortions in relative prices and undermines the tradable goods sector, the neoliberal model advocates a reduction in the budget deficit in order to keep low and stable rates of inflation. By the same token, fiscal reform plays a crucial part in the model with an emphasis on both expenditure reduction and revenue increases (Bulmer-Thomas, 1996: 11). In this context, privatisation is deemed a condition for large capital flows from abroad and it is also considered a policy to cut expenditure (through eliminating subsidies) and to increase revenue (through asset sales or increased tax receipts). Stable and low rates of inflation, a reduction in the public deficit, and an overall macroeconomic discipline are expected to create conditions for economic certainty, which encourages capital inflows from abroad, savings, investments and in general a larger production of tradable goods. Hence, this pattern reinforces the outward-looking growth model. Under the neoliberal approach, it is also contended that low inflation rates prevent inequality due to the relative vulnerability to inflation of low income households.

Expectations for the inequality-growth relationship.

During the 1980s the prevailing global political economy added impetus to market oriented policies and discouraged any further attempt of protectionism. The ascendancy of neoliberal ideas during the Reagan era in the US, the reformist agenda of developing countries based on market-oriented policies, the collapse of the communist system, and the overall global expansion of economic liberalisation signalled a political and economic global shock that was characterised by placing special emphasis on outward-looking growth, market forces, a dominant role for the private sector in the economy, and the international mobility of capital. Under these circumstances and from the neoliberal perspective, we may expect improvements in the global distribution of income and an inverse relationship between income and inequality during the last two decades.

On the other hand, before the 1980s the prevailing economic policies of the post war period can be summarised as inward-looking development and protectionist strategies in developing countries; central planning methods in the former Soviet Union and Central and Eastern Europe countries, besides other republics; developmental strategies with staged economic liberalisation in East Asian countries; and limited economic liberalisation in developed economies. On that basis, it can be argued that during this period the primacy of the state played a more preponderant role than market forces. Consequently, over these years and from a neoliberal viewpoint, we may expect that inequality rises as income expands since market distortions and government interventions are usually deemed inefficient and inequitable in the neoliberal approach (Kanbur, 2000: 795). In this sense, we may expect that the relationship between the level of income and inequality before the 1980s presents a positive slope. Therefore, a long term relationship between economic growth and income inequality, over the post war period and from a neoliberal view point, may be depicted by an inverted-U curve with the turning point somewhere in the 1980s.

3. Preliminary evidence

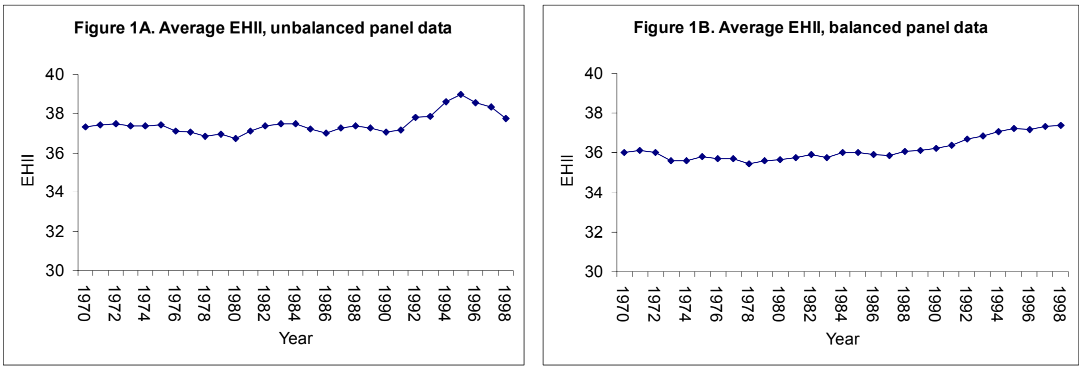
Trends in income distribution.

Initially, we explore the evolution of income distribution by plotting simple average values of the inequality measure (*EHII*). Figures 1A and 1B illustrate the unbalanced and balanced sample outlined earlier respectively. In general, it can be observed that over the 1970s, which is also the period of restricted economic liberalisation, inequality does not follow an increasing pattern, but declines slightly. On the other hand, the curves show an upward trend since the early 1980s and this trend seems to be reinforced during the late 1980s and early 1990s. In this respect some authors have also documented similar conclusions (Morrison, 2000; Gottschalk and Smeeding, 2000; Flemming and Micklewright, 2000; Smeeding, 2002; Galbraith and Kum 2002).

It should be added that only in the unbalanced sample, the period of rising inequality appears to reverse in 1996. In this sense, Galbraith and Kum (2003: 14) notice that the lower average of inequality over the late 1990s maybe spurious on account of variable lags in reporting underlying data to UNIDO and other agencies. As a matter of fact, the number of countries contained in our sample in the last years drops substantially. Consequently, the decreasing inequality illustrated in Figure 1 by the end of the period, maybe caused by gaps across the panel.[1]

[1] We also plot the unbalanced and balanced sample weighted by GDP, GDP per capita and population. The analysis is conducted for both developed and developing economies. The countries are divided according to the World Bank income classification using GNI per capita for 2000 and the two groups contain low and middle income economies and high income economies respectively. By separating the samples, it is visible that the upturn in inequality started later across developing economies, and it is

Figure 1. Average values of EHII for each year.



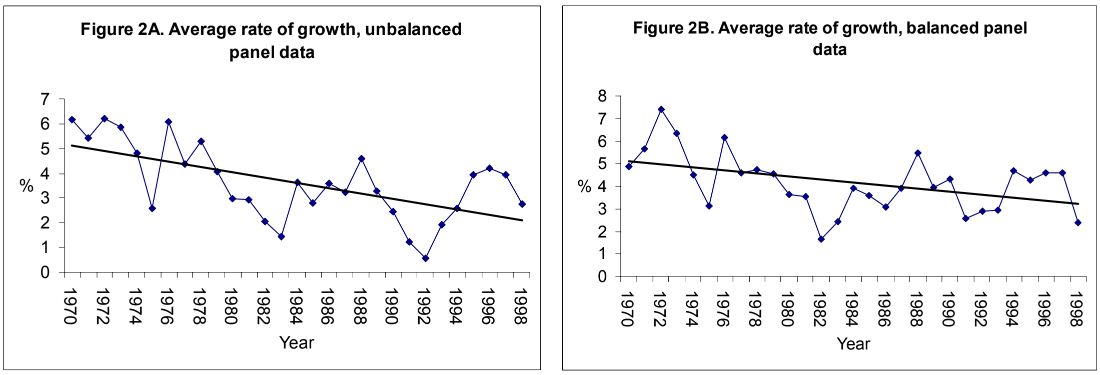
There is an increasing consensus in the literature claiming that inequality has risen over the age of free market liberalism, and the preliminary evidence above is in keeping with these findings. However, there are some discrepancies among the studies that try to determine the upturn period. In this context, Galbraith and Kum (2003) find the upturn beginning in 1979 for OECD countries and 1987 for non-OECD countries. Smeeding (2002) asserts that inequality rose from the late 1980s in almost every OECD nation, while it began to rise in the 1990s in Russia and Czech Republic. He also holds that from the 1970s inequality only increased in the United States and the United Kingdom, but the trend seems to have flattened out in both countries by the end of the 1990s. Gottschalk and Smeeding (2000) find that income inequality in over 20 wealthy nations declined through the 1970s and started increasing in the mid-1980s. Flemming and Micklewright (2000) state that earnings inequality increased through the 1990s in Central and Eastern Europe and the former Soviet Union. We will study upturn periods with further detail through continuous time-series across 31 countries later in this paper.

confirmed that the decreasing pattern of inequality since the late 1990s depends on the composition of the panel, as this trend is more robust in the unbalanced samples.

Trends in economic growth.

Figures 2A and 2B display the evolution of economic growth on yearly basis across the countries contained in our unbalanced and balanced sample respectively. The variable on economic growth is annual percentage growth rate of GDP based on constant U.S. dollars and is obtained from the World Development Indicators CD-Rom 2002. The rate of growth appears to be unstable and tends to slow down over the whole period, as it displays a downward trend. Hence, the composition of the panel does not seem to affect this pattern.[2]

Figure 2. Average values of rate of growth for each year.



These findings are in keeping with the perception of some authors who have stressed that over the last decades, economic growth proved to be unsteady and rather slow on average (Atkinson and Bourguignon, 2000: 2-3; Onaran, 2004: 2).

Through the ascendancy of market-oriented ideas in the early 1980s, some of the main expectations were to re-establish the rapid and sustained growth that characterised the boom of the Bretton Woods era, to improve income distribution and in general to re-establish the path to prosperity. Nevertheless, the empirical evidence

[2] The analysis is extended by plotting the unbalanced and balanced sample weighted by GDP and GDP per capita and is conducted for both developed and developing economies. The countries are divided according to the criteria already explained. In any case it is confirmed that the rate of growth is unsteady and follows a downward trend along the whole period.

exposed above indicates that during the era of economic liberalisation, rapid economic growth has not been restored, the rates of growth seem to be unsteady, and inequality has increased on average.

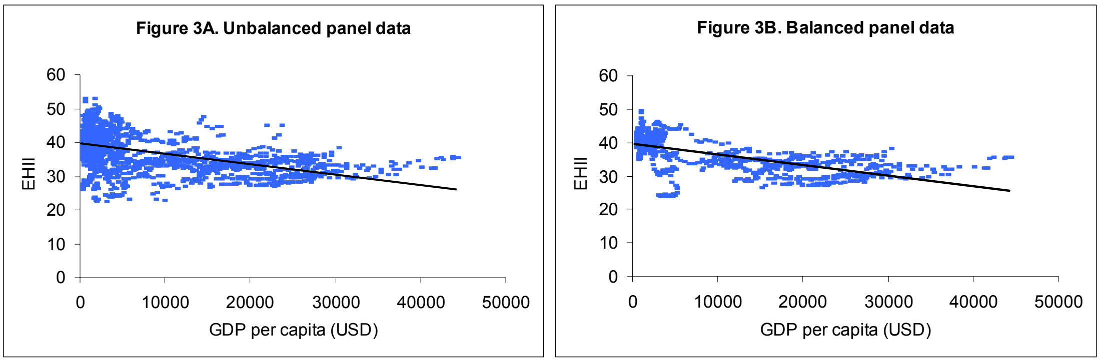
The relationship between inequality and growth.

Finally, figures 3A and 3B explore the pattern of the relationship between inequality and economic growth through both the unbalanced and balanced data set respectively. The variable on inequality is *EHII* as outlined earlier. Economic growth is represented by different levels of development or income through the GDP per capita expressed in 1995 U.S. dollars. Previous studies have also considered GDP per capita to illustrate the inequality-growth relationship (Deininger and Squire, 1998; Galbraith and Kum, 2002; De Gregorio and Lee, 2002). In both cases, it appears that inequality tends to decline with economic growth, independently of the level of development. However, it should be noted that inequality seems to increase slightly at high levels of GDP per capita.[3]

Although it is possible to observe a slight increase in inequality at high levels of income, in general figures 3A and 3B might suggest that inequality tends to decline with economic growth during the age of free market liberalism. However, this preliminary assertion deserves further attention because it was illustrated that inequality has actually risen over the last decades when *EHII* was explored ignoring its relationship with growth. Alternatively, another likely cause of this trend is that low income countries are normally associated with higher levels of income inequality.

[3] The overall samples are also split in low-middle income countries and high income countries sub-samples. In any case it is confirmed that inequality tends to decline on average with economic growth or at higher levels of income, independently of the level of development. In addition, a slight increase in inequality at a high level of income is also captured in every sub-sample.

Figure 3. Inequality-growth relationship.



On the other hand, it has been already argued that from a neoliberal perspective we may expect that the inequality-growth relationship follows an inverted-U curve over the period comprised in the sample – 1970-98. Nevertheless, the preliminary evidence explored above does not seem to support this view. In contrast, it appears to illustrate an ordinary-U curve in which most of the observations are located in the downward portion.[4] Therefore, we need to turn to quantitative methods so as to explore the possible existence of a systematic and convincing relationship between inequality and income level over the last decades.

4. Panel data approach.

The general regression panel data model for the income inequality-growth relationship follows:

$$EHII_{it} = \alpha_i + \beta_1 Y_{it} + \beta_2 Y^2_{it} + u_{it} \qquad (1)$$

[4] When we fit the samples, for both developed and developing economies, to five different equations – Linear, Logarithm, Polynomial, Power and Exponential - we find that in four out of six samples, the Polynomial equation following an inverse U-shaped curve displays the highest R square.

in which *EHII* is the inequality measure and *Y* is GDP per capita in 1995 US dollars. The subscripts *i* and *t* indicate country and year respectively. The error term u_{it} is assumed to satisfy white noise assumptions, i.e. zero mean, constant variance σ^2 and serially uncorrelated, which is denoted as $u_{it} \sim$ I.I.D. $(0, \sigma^2)$, α_i lets the intercept vary for each country and captures country-specific effects, finally β_1 and β_2 are parameters to be estimated.[5]

Before moving on the discussion about the estimation method, it should be emphasized that the quadratic function, or more generally, the second degree polynomial in *Y*, can be plotted as a parabola. This key feature of the model allows us to test formally for two different patterns of the inequality-income relationship. Firstly, if the sign of the coefficient β_2 on the quadratic explanatory term is negative ($\beta_2 < 0$), the curve will display a peak suggesting that a maximum point can be found in the equation. Under these circumstances, income inequality increases in the early stage of economic growth, reaches a peak, and then decline with a higher level of per capita income. It has already been pointed out that this inverse U-shaped pattern might correspond to the neoliberal prediction.

Secondly, if the sign of the coefficient β_2 on the quadratic explanatory term is positive ($\beta_2 > 0$), the curve will display a valley suggesting that a minimum point can be found in the equation. In this case, an ordinary-U shape instead of an inverted-U is captured, which implies that the degree of inequality first declines and then increases with further economic growth. The presence of a real rather than an inverse U-shaped

[5] Previous studies in the literature have also applied quadratic equations, but the formulations differ. For example, Deininger and Squire (1998) apply the specification suggested by Anand and Kanbur (1993) which includes income in the regression as *Y* and *1/Y*, De Gregorio and Lee (2002) apply the square specification as in Equation 1, and Galbraith and Kum (2002) employ a *log* transformation of GDP per capita. In this case we confine our attention to the square specification, because after conducting different regressions it proved to capture a more systematic relationship and the estimated parameters are slightly more significant than the other formulations.

relationship is expected to test recent findings of rising inequality over the last decades, which is rather the period of increasing economic liberalisation.

Unbalanced sample

Initially, we regress Equation 1 with the unbalanced sample employed in the preliminary analysis. The overall fit of the model is examined by performing two formal specification tests. Firstly, The Breusch and Pagan Lagrange Multiplier test (1980) rejects the standard OLS assumption that the intercept value is the same across countries, and therefore there are country-specific effects in the model.[6] Secondly, the Hausman test (1978) suggests that the country-specific effects are correlated with the regressor in the equation.[7] The no correlation assumption is an important pillar of the random-effects model (REM), but in this case is violated. Hence, the random-effects estimates are inconsistent and the fixed-effects specification (FEM) is more robust.

The specification tests and the results obtained from the pooled regression and the two panel estimations are reported in Table 1 from column 1 to column 3. It is interesting to note that the coefficient of Y^2 is significant and positive in the three equations. Consequently, this analysis captures a U-shape where income inequality first diminishes and then is found to rise with increasing output.

Before adopting the FEM as the final estimation, it is important to test whether the model satisfies white noise assumptions, by the same token an autocorrelation test on the error term u_{it} should be available. We find that the first and second order AR

[6] The Breusch and Pagan Lagrange Multiplier test (1980), based on OLS residual and under the null hypothesis: $\sigma^2_\varepsilon = 0$, i.e., $\alpha_i = \alpha$, is distributed as a χ^2 with one degree of freedom (Greene, 2000: 572-3). The LM test statistic is equal to 10,081.52, which far exceeds the 5 percent critical value of the χ^2 distribution with one degree of freedom, 3.84. As the null hypothesis is rejected, it is concluded that there are country-specific factors, and the OLS regression is inappropriate.

[7] Under the null hypothesis that the country-specific effects and the regressors are uncorrelated, the Hausman test (1978), is based on an asymptotic χ^2 distribution with two degrees of freedom. The Hausman test statistic is equal to 49.58, which exceed the 5 percent critical value of the χ^2 distribution with two degrees of freedom, 5.99. Since the null hypothesis is rejected, the random-effects estimators are inconsistent and the FEM is preferred.

tests, conducted on the fixed-effects regression and reported in column 3 of Table 1 are not satisfied.[8] So as to address this problem, it is required to explore the possibility that autocorrelation may arise owing to model miss-specification, to be precise, because of an omitted lagged dependent variable. So, Equation 1 is extended and transformed into a dynamic panel data model (DPDM) by adding a lagged endogenous variable as follows:

$$EHII_{it} = \alpha_i + \gamma EHII_{it-1} + \beta_1 Y_{it} + \beta_2 Y^2_{it} + \eta_i + u_{it} \qquad (2)$$

On the other hand, the inclusion of a lagged dependent variable introduces a source of persistence over time, correlation between the right hand regressor $EHII_{it-1}$ and the error term u_{it}. In addition, DPDM are characterised by individual effects η_i caused by heterogeneity among the individuals.[9] Hence, it is necessary to adopt further estimation and testing procedures for this model.

In order to estimate the model consistently and efficiently we use a generalized method of moment estimation (GMM) for DPDMs proposed by Blundell and Bond (1998). Initially, the estimation method eliminates country-effects (η_i) by expressing Equation 2 in first differences as follows:

$$EHII_{it}\text{-}EHII_{it-1} = \gamma(EHII_{it-1}\text{-}EHII_{it-2}) + \beta_1(Y_{it}-Y_{it-1}) + \beta_2(Y^2_{it}-Y^2_{it-1}) + (u_{it}\text{-}u_{it-1}) \qquad (3)$$

in addition, on the basis of the following standard moment condition:

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[8] The AR test statistic of order one is equal to 52.46 and the AR test statistic of order two is equal to 35.97, both with a negligible *P* value. The tests of serial autocorrelation up to order two are not satisfied as they reject the null hypothesis: $\rho_1 = \rho_2 = 0$. We also find evidence of serial autocorrelation when conducting the OLS and random-effects regressions as reported in Table 1, column 1 and column 2 respectively.

[9] For an elaboration in this point see Badi H. Baltagi, *Econometric analysis of panel Data* (Sussex: John Wiley & Sons, 2001) 2nd Ed., pp. 129-30.

$E(EHII_{i,t-s}\,\Delta u_{it}) = 0$, for $t = 3,\ldots,N$ and $s \geq 2$

that is, lagged levels of $EHII_{it}$ are uncorrelated with the error term in first difference. The method uses lagged levels of $EHII_{it}$ as instruments to control for likely endogeneity of the lagged dependent variable, reflected in the correlation between this variable and the error term in the equation in first differences. The resulting GMM estimator is known as the difference estimator and was proposed by Arellano and Bond (1991).

However, Blundell and Bond (1998: 115-6) state that the GMM estimator obtained after first differencing has been found to have large finite sample bias and poor precision. They attribute the bias and poor precision of this estimator to the problem of weak instruments, as they assert that lagged levels of the series provide weak instruments for the first difference. So as to improve the properties of the standard first-differenced GMM estimator Blundell and Bond justified the use of an extended GMM estimator, on the basis of the following moment condition:

$E[\Delta EHII_{it-1}\ (\eta_i\ + u_{it})] = 0$

that is, there is no correlation between lagged differences of $EHII_{it}$ and the country specific effect. The method therefore uses lagged differences of the endogenous variable as instruments for equations in levels, in addition to lagged levels of $EHII_{it}$ as instruments for equations in first differences. The extended GMM, therefore, encompasses a regression equation in both differences and levels, each one with its specific set of instrumental variables. This type of estimation, called system estimator, not only improves the precision but also reduces the finite sample bias.

The model assumes that the disturbances u_{it} are not serially correlated. If this is the case, there should be evidence of first order serial correlation in differenced residuals (*i.e.* $u_{it} - u_{it-1}$), but no evidence of second order serial correlation (Arellano et al., 2002: 5-8). It is an important assumption because the consistency of the GMM estimators hinges upon the fact that $E[\Delta u_{it}\, \Delta u_{it-2}] = 0$. Thus, tests of autocorrelation up to second order in the first-differenced residuals are required. Moreover, so as to assess the validity of the instruments a Sargan test of overidentifying restrictions proposed by Arellano and Bond (1991) is reported.

Table 1. Unbalanced panel data

	(1) OLS	(2) REM	(3) FEM	(4) GMM sys	(5) GMM sys	(6) GMM sys Orthogonal
$EHII_{t-1}$				0.680 *	0.680 *	0.702 *
Y	-6.38E-04 *	-1.18E-04 *	5.69E-05	-6.07E-04 *	-9.18E-04 *	-1.09E-03 *
Y^2	1.21E-08 *	5.70E-09 *	2.72E-09 *	2.11E-08 *	4.63E-08 *	6.15E-08 *
Y^3					-4.99E-13 *	-8.24E-13 *
Constant	40.325 *	37.511 *	36.973 *	13.539 *	14.016 *	13.399 *
BP LM test	[0.000]					
Hausman test		[0.000]				
Sargan test:				[0.862]	[0.818]	[0.787]
AR(1) test:	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
AR(2) test:	[0.000]	[0.000]	[0.000]	[0.804]	[0.873]	[0.883]
Wald test for Y^3					[0.000]	[0.000]
Observations	2289	2289	2289	2173	2173	2173
Countries	116	116	116	116	116	116
Min turning Point	26,269	10,322		14,387	12,394	11,505
Max turning Point					49,470	38,265

Notes:

Dependent variable: *EHII*

BP LM, Hausman, Sargan, serial correlation and Wald test are *P* values

* Significant at 5%; **Significant at 10%

Column 4 of Table 1 provides the results obtained from the GMM system estimation. The tests of serial correlation in the first differenced residuals are in both cases consistent with the maintained assumption of no serial correlation in the disturbances u_{it},[10] while the Sargan test of overidentifying restrictions is unable to reject the validity of the instruments.[11] Under these circumstances, it is possible to treat the results as valid. In this case, we also find that the coefficient of Y^2 is positive and significant, as in the outcomes obtained previously. In order to determine the value of the minimum turning point we follow Hsing and Smyth (1994) and Jacobsen and Giles (1998) procedure. Based on estimated parameters, taking the first derivative of the dependent variable with respect to *Y* and setting the first condition equal to zero, it corresponds to $14,387.[12]

Cyclical pattern.

The preliminary analysis of the unbalanced data set gives some evidence of decreasing inequality by the late 1990s, although this trend might be the result of discontinuity in the *EHII* data set. Through the panel data approach we asses the existence of a second turning. In this sense, Equation 3 is extended into a third-degree polynomial by adding the cube of income per capita as follows:

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[10] Under the null hypothesis of no autocorrelation, the tests are asymptotically distributed as N(0,1) In this case, The second-order serial correlation test statistic is equal to 0.249 and the *P* value is equal to 0.805; therefore, the test fail to reject the null that the first differenced error term is not second order serially correlated. The first-order serial correlation test statistic is equal to -4.314 with a negligible *P* value; hence, by construction, the test rejects the null that this process does not exhibit first-order serial correlation.

[11] Under the null hypothesis that the instruments are not correlated with the error process, the Sargan test is asymptotically distributed as a chi-square with as many degrees of freedom as overidentifying restrictions. In this case, the Sargan test statistic is equal to 63.67 and the *P* value is equal to 0.862; so, the test is unable to reject the validity of the instruments.

[12] This GMM system regression does not include differential intercept dummies. When yearly dummy variables are incorporated into the equation the minimum turning point increases up to $17,769, but when country dummy variables are added both the Sargan test and the first-order serial correlation test are not satisfied. The first-differenced GMM estimators are also obtained. We find that without differential intercept dummies the minimum turning point is $18,287 and with yearly dummies the minimum turning point is $15,103. When adding country dummies, there is some evidence of serial correlation in the disturbances and the Sargan test is not satisfied. (Results not reported).

$$I_{it} - I_{it-1} = \gamma(I_{it-1} - I_{it-2}) + \beta_1(Y_{it} - Y_{it-1}) + \beta_2(Y^2_{it} - Y^2_{it-1}) + \beta_3(Y^3_{it} - Y^3_{it-1}) + (u_{it} - u_{it-1}) \quad (4)$$

The results from the GMM system regression are reported in column 5 of Table 1. The cubic term enters negatively and significantly in the equation implying that inequality reaches a peak and then reverses with the presence of a second turning point. It is worth nothing that the coefficients of the lagged dependent variable, income per capita and its square remain statistically significant and their signs do not change. Moreover, the magnitude of the coefficients does not change substantially. So as to confirm whether Y^3 belongs in the model, a Wald test for excluding variables is conducted. The test leads to the conclusion that the unrestricted regression or the cubic equation is more appropriate.[13]

Some authors have claimed that the long-run income distribution may be better described in terms of long period cycles that may be modelled by a polynomial function to the nth degree (Hsing and Smyth, 1994: 113; Jacobsen and Giles, 1998: 420), while they also stress the possibility of a high degree of correlation among the independent variables. The Multicollinearity problem may arise in polynomial equations because the explanatory variable appears with various powers. Thus, the various *X's* are likely to be highly correlated.[14]

With the above in mind, an orthogonal transformation as in Doornik et al. (2002: 35), is performed to reduce multicollinearity. This transformation takes each observation in deviation from the future means, together with a standardisation.

[13] Under the null hypothesis: $\beta_3 = 0$, the Wald test follows a χ^2 distribution with 1 df equal to the restrictions imposed by the null hypothesis. In this case, the Wald test statistic is 7.65 and the *P* value is almost zero, indicating that the restricted regression is not valid.

[14] Terms like X^2, X^3, X^4, etc are all nonlinear functions of *X* and therefore, strictly speaking, do not violate the multicollinearity assumption of the classical model. Nevertheless, the correlation coefficient will show the *X's* to be highly correlated, which will make it difficult to estimate parameters precisely in polynomial equations. On the other hand, if the purpose of econometric analysis is just forecasting or prediction, as in the present case, multicollinearity is not a serious problem since the higher the R^2, the better the prediction. (For a discussion see Gujarati, 2003: 227, 343-4, 369).

Results are shown in column 6 of Table 1. We find that Y^3 also enters negatively and significantly, whereas the Wald test emphasises that the restricted or quadratic equation is not valid. The minimum and maximum turning points correspond to \$11,505 and \$38,265 respectively. Although the value at which the maximum turning point is located in the orthogonal equation is lower than that of the original GMM system equation, it is still in a relatively high position, suggesting that increasing inequality reverses at a high level of development.

In order to test if this cyclical pattern is associated with the level of development, the overall sample is split in developed and developing countries according to the income classification outlined earlier. Table 2 illustrates the outcome of the GMM system regressions for both sub-samples; it also shows results when orthogonal transformations are applied in the cubic equations. In any case, the Wald test for excluding variables is unable to reject the null hypothesis that the coefficient on Y^3 is equal to zero. These findings suggest that income distribution follows a cyclical pattern during the age of economic liberalisation, in which inequality tends to decline with economic growth after a prolonged period of time, independently of the level of development.

These results are in keeping with the preliminary evidence obtained from the unbalanced data sample. On the other hand, Table 2 shows that the first-order serial correlation test is not satisfied in the developed countries sub-sample. Hence, results from this group must be taken with reservations.[15] We now test the existence of a cyclical pattern through a balanced panel data set.

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[15] Some authors have demonstrated that GMM estimators generally perform better with a relatively large *N* (Blundell and Bond, 1998; Judson and Owen, 1999). On the other hand, the size of *N* in the developed countries sub-sample is relatively small, which might be a cause of imprecision and lack of efficiency. So as to overcome any presence of small sample bias, the overall sample is also split by adopting different criteria. The first group comprises countries with low and lower-middle income per capita, while the second comprises countries with upper-middle and high income per capita. In this

Table 2. Unbalanced panel data (developed and developing countries)

	Developing countries			Developed countries		
	GMM sys	GMM sys	GMM sys Orthogonal	GMM sys	GMM sys	GMM sys Orthogonal
$ehii4_{t-1}$	0.715 *	0.703 *	0.731 *	0.626 *	0.638 *	0.605 *
Y	-1.05E-03 *	-3.60E-03 *	-3.75E-03 *	-3.10E-04 **	-1.88E-03 *	-1.23E-03 *
Y^2	1.67E-07 *	1.05E-06 *	1.12E-06 *	1.04E-08 *	8.74E-08 *	5.69E-08 *
Y^3		-6.81E-11 *	-7.10E-11 *		-1.09E-12 *	-6.84E-13 *
Constant	12.194 *	13.832 *	12.710 *	13.958 *	22.498 *	19.715 *
Sargan test:	[0.519]	[0.505]	[0.743]	[1.000]	[1.000]	[1.000]
AR(1) test:	[0.000]	[0.000]	[0.000]	[0.132]	[0.115]	[0.126]
AR(2) test:	[0.798]	[0.747]	[0.749]	[0.580]	[0.500]	[0.560]
Wald test for Y^3		[0.000]	[0.000]		[0.003]	[0.002]
Observations	1484	1484	1484	689	689	689
Countries	89	89	89	27	27	27
Min turning point	3,140	2,165	2,077	14,907	14,977	14,722
Max turning point		8,146	8,480		38,319	40,688

Notes:
Dependent variable: *EHII*
Sargan, serial correlation and Wald test are *P* values
* Significant at 5%; **Significant at 10%

Balanced sample.

We apply the balanced panel data set to explore the income-inequality relationship. Results obtained from the overall sample are reported in Table 3. The GMM system method applied in the quadratic regression fits a U-shaped pattern, in which the predicted turning point is $16,750. This level of GDP per capita is larger than its counterpart predicted in the unbalanced sample ($14,387), because the balanced data set contains a larger proportion of developed economies. On the other hand, neither the GMM system method nor the orthogonal transformation captures a cyclical pattern when the equation is extended into a third degree polynomial, since the

way, the size of *N* does not drop drastically in any sub-sample. We conduct GMM system regressions for quadratic and cubic specifications and also apply orthogonal transformations for both sub-samples. In any case, the first and second order serial correlation tests are satisfied, whereas the Wald test leads to the conclusion that the Y^3 should not be excluded from the model in any of the sub-samples (results not reported).

coefficients for Y^2 and Y^3 are not statistically significant in any case. Moreover, the Wald test for excluding variables does not reject the restricted equation, suggesting that the cubic model is inappropriate. In this case, the overall sample is not split in sub-groups since every country will be analysed separately through a time-series approach.

Table 3. Balanced panel data

	GMM sys	GMM sys	GMM sys Orthogonal
$EHII_{t-1}$	0.7701 *	0.7855 *	0.9238 *
Y	-2.91E-04 *	-3.77E-04 **	-1.98E-04
Y^2	8.69E-09 *	1.39E-08	1.07E-08
Y^3		-7.97E-14	-1.27E-13
Constant	9.460 *	9.120 *	3.209 *
Sargan test:	[1.000]	[1.000]	[1.000]
AR(1) test:	[0.035]	[0.033]	[0.028]
AR(2) test:	[0.548]	[0.552]	[0.604]
Wald test for Y^3		[0.665]	[0.375]
Observations	868	868	868
Countries	31	31	31
Min turning point	16,750	15,689	11,702
Max turning point		100,423	44,607

Notes:
Dependent variable: *EHII*
Sargan, serial correlation and Wald test are *P* values
* Significant at 5%; **Significant at 10%

The empirical evidence above points in favour of an ordinary U-shaped relationship between income inequality and growth over the period 1970-1998. This finding is robust and fits both developed and developing economies. On the other hand, the presence of a maximum turning point over the long-run, vanishes when we use the balanced panel data set and this is in keeping with the preliminary evidence provided earlier. Hence, the evidence of a cyclical pattern is weak.

Cross-country analysis

A number of studies have found an inverted-U relationship between income and inequality by using cross-sectional analysis in the absence of adequate longitudinal-data (Bourguignon, 1994; Milanovic, 1995; Jha, 1996). However, it has been stressed that this approach does not yield appropriate conclusions as it does not deal with intertemporal relationships (Deininger and Squire, 1998: 276; De Gregorio and Lee, 2002: 404). In order to explore the potential bias that might arise between the panel data estimates and cross-section approach, we group the data in 5-year average periods and obtain six samples for unbalanced and balanced data sets.[16] We do indeed find that the quadratic terms display a negative sign in the log specification, as in De Gregorio and Lee (2002), suggesting the existence of an inverted-U curve; but their coefficients are significant only in the first four equations of the unbalanced sample. The Anand-Kanbur specification also reveals the existence of an inverted-relationship, as in Deininger and Squire (1998), but only in the first three equations of the unbalanced sample, and only in one of them the coefficient of the inverse term is significant. The remaining regressions and the linear specification capture an ordinary-U pattern, but the significance of the coefficients is ambiguous. Results are illustrated in Table 4

This approach derives weak empirical support for the Kuznets hypothesis. Moreover, globalisation does not seems to be a factor affecting the traditional inverted-U relationship found in cross-sectional data. In contrast, the relationship between income and inequality seems to depend on the specifications of the equations and on the number of observations. In general, this approach lacks robustness and its results are ambiguous.

[16] Only the last sample comprises a four-year averages period between 1995 and 1998

Table 4. Cross-country regressions

Especification	1970-1974		1975-1979		1980-1984		1985-1989		1990-1994		1995-1998	
Unbalanced												
Level												
Y	-7.05E-04	*	-7.03E-04	*	-6.70E-04	*	-6.03E-04	*	-4.58E-04	*	-3.73E-04	**
Y^2	1.14E-08		1.26E-08		1.19E-08		1.11E-08		7.11E-09		4.84E-09	
Log												
Log Y	11.719	*	-0.734	***	12.419	*	5.541		2.534		2.203	
$(\text{Log } Y)^2$	-0.885	*	-1.27E-08	*	-0.909	*	-0.463	**	-0.262		-0.236	
Anand-Kanbur												
Y	-5.19E-04	*	-4.95E-04	*	-4.29E-04	*	-2.83E-04	*	-2.23E-04	*	-2.00E-04	*
1/Y	-302.435		-625.614	**	-443.328		259.298		251.013		347.788	
Observations	77		85		96		93		100		84	
Balanced												
Level												
Y	-8.89E-04	**	-8.50E-04	*	-8.59E-04	*	-6.90E-04	*	-5.84E-04	*	-5.23E-04	*
Y^2	1.98E-08		1.86E-08		1.80E-08	***	1.24E-08		9.30E-09	***	8.14E-09	***
Log												
Log Y	7.080		4.758		4.742		4.130		4.946		5.883	
$(\text{Log } Y)^2$	-0.601		-0.451		-0.454		-0.400		-0.436		-0.474	
Anand-Kanbur												
Y	-4.36E-04	*	-3.64E-04	*	-3.46E-04	*	-2.77E-04	*	-2.45E-04	*	-2.16E-04	*
1/Y	461.593		823.663		930.918		1076.318		945.411		638.064	
Observations	31		31		31		31		31		31	

Notes:
Dependent variable: *EHII*
* significant at 1%, ** significant at 5%, *** significant at%

Although the panel data analysis determines the level of income in which the minimum turning point occurs, it does not date the minimum and does not determine when the maximum occurs either, if any. Moreover, although the panel data analysis obtains conclusions for two different sub-samples, it does not reach conclusions for specific country cases. With the above in mind, we complement our findings through a time-series analysis. This approach allows us to explore particular country cases in order to obtain further evidence and to predict both date and level of GDP per capita in which turning points occur.

5. Time-series approach.

Some authors have pointed out that in order to explore the evolution of inequality, further intertemporal evidence should ideally be based on time-series analysis from single countries (Bruno et al., 1998; Morrison, 2000). Moreover, Atkinson et al. (2001: 22-3) notice that the availability of 20 to 40 years of estimates on income inequality in many nations makes it possible to examine the determinants and consequences of long periods of distributional change, e.g., the relationship between inequality and growth. In this context, it is worth complementing the panel data analysis through a time-series approach to obtain additional conclusions.

So as to conduct the time-series analysis, we take the balanced panel data and decompose it into countries. In this way, it is possible to obtain 31 time-series with 29 observations each, along the period 1970-98. Initially we test a systematic relationship between inequality and growth by applying linear and quadratic equations in levels and log transformation of *Y* and the functional form suggested by Anand and Kanbur (1993) as follows:

Linear

Level $$EHII_t = \alpha + \beta_1 Y_t + u_t \quad (5)$$

Log $$EHII_t = \alpha + \beta_1 lnY_t + u_t \quad (6)$$

Quadratic

Level $$EHII_t = \alpha + \beta_1 Y_t + \beta_2 Y^2{}_t + u_t \quad (7)$$

Log $$EHII_t = \alpha + \beta_1 lnY_t + \beta_2 (lnY_t)^2 + u_t \quad (8)$$

Anand-Kanbur $$EHII_t = \alpha + \beta_1 Y_t + \beta_2 1/Y_t + u_t \quad (9)$$

The process to select the model is conducted under the following criteria. Firstly we determine if the linear model can be rejected in favour of a quadratic equation or the Anand-Kanbur specification. To reject the linear model, at least one of the equations from 7 to 9 has to meet two conditions – the Lagrange Multiplier test for adding variables has to reject the restricted regression[17] and all the coefficients in the equation have to be statistically significant at any conventional level - otherwise the model is assumed to be linear.

If more than one equation satisfies the two conditions above, three additional fitness tests for model selection are undertaken – Akaike information criteria (AIC), Schwarz information criteria (SIC) and Ramsey's RESET test (RRT).[18] The equation that performs better across these tests is selected as the appropriate nonlinear model.

The existence of a cyclical pattern in the long-run income distribution that may follow long waves is also explored. In this sense, Equation 7 and Equation 8 are extended into a third degree polynomial by adding a cubic term as follows:

Cubic equations

Levels $$EHII_t = \alpha + \beta_1 Y_t + \beta_2 Y^2{}_t + \beta_3 Y^3{}_t + u_t \qquad (10)$$

Logs $$EHII_t = \alpha + \beta_1 lnY_t + \beta_2 (lnY_t)^2 + \beta_3 (lnY_t)^3 + u_t \qquad (11)$$

The linear and quadratic models are rejected and the inequality-growth relationship is regarded as cyclical, if at least one of the two equations above satisfies the Lagrange Multiplier test for adding variables and all the coefficients in the

[17] The LM statistic follows the chi-square distribution with df equal to the number of restrictions imposed by the restricted regression, one in the present case. The null hypothesis is "the restricted regression is adequate, i.e. the additional coefficient is equal to zero".

[18] In comparing two or more competing equations, the equation with the lowest value of AIC is preferred. Like AIC, the lower the value of SIC the better the model. The RRT is a general test of specification error that can be conducted on the basis of the F test under the null hypothesis that the model is correct.

equation are statistically significant at any conventional level. If both of the cubic equations satisfy the previous conditions, the three additional fitness tests for model selection, as described earlier, are conducted so as to determine the preferred specification. In total, seven regressions for every country case are undertaken, the results are available upon request.

The time-series analysis does not lead to the existence of a universal trend of inequality, since it captures quadratic and cubic patterns with diverse turning points as well as linear trends both positively-sloped and negatively-sloped. In only two countries it is not possible to capture any systematic relationship. Before moving further to a discussion about the results, it is important to raise three additional considerations about the estimation procedure.

Firstly, it is worth noting that the Durbin-Watson *d* test and the Breusch-Godfrey (BG) test[19] show evidence of autocorrelation in most of the country-cases – only in two countries it is not detected by the tests. In this context, some authors examining the pattern of income inequality through time-series analysis have stressed that in the presence of residual autocorrelation results are flawed (Fosu, 1993; Jacobsen and Giles, 1998). Thus, we correct for the presence of autocorrelation by using Cochrane-Orcutt method as in Hsing and Smyth (1994) and the Prais-Winsten method.

If autocorrelation persists, we test the possibility that it may arise due to model-mis-specification by adding a lagged dependent variable. However, the inclusion of a lagged dependent variable introduces a source of persistence over time

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[19] One of the main assumptions underlying the *d* statistic is that the disturbances u_t are generated by the first-order autoregressive scheme: $u_t = \rho u_{t-1} + \varepsilon_t$. It is therefore used to test first order serial autocorrelation under the null hypothesis H_0: $\rho = 0$. The BG test allows for higher-order AR(ρ) schemes and follows a chi-square distribution with ρ df. For this particular case, we test up to second order serial autocorrelation under the null hypothesis $H0$: $\rho_1 = \rho_2 = 0$; that is, there is no serial correlation of first and second order.

– correlation between the right hand regressor $EHII_{t-1}$ and the error term u_t. Due to the presence of simultaneity, the method of two-stage least squares (2SLS) and instrumental variables is performed. In this way, it is possible to obtain consistent and efficient estimators. We notice that after applying this approach, serial autocorrelation persists. Thus, it is possible to argue that most of the equations in the time-series analysis suffer from pure autocorrelation and not necessarily from specification bias as the equations in the panel data approach.

It should be added that any of the three methods outlined earlier are able to correct for autocorrelation in 13 out of 29 country-cases, in the corresponding selected equation or in any other suitable specification. With the above in mind, the first differenced method is performed in the particular country-cases with persistent autocorrelation. The application of this method solves the AR problem; however, the corresponding relationship vanishes as the coefficients of the explanatory variables are no longer significant. Under these circumstances, we take the results from the selected equations as valid and allow for autocorrelation only in these country-cases.[20]

Secondly, the estimation of models with non-stationary data can lead to spurious regressions. Jacobsen and Giles (1998: 408) point out that modelling the relationship between income distribution and economic growth with non-stationary data casts grave doubts on the reliability of the findings to date. On the other hand, if a time-series has a unit root, its first differences can be stationary; that is, the original time-series is I(1). A series is integrated of order d or I(d) if after being differenced d times it becomes stationary. In addition, although linear combinations of I(1) series can produce another I(1) series, there are special cases in which their combination can

[20] Bruno et al. (1998) explored data for India and found an ordinary U-shaped relationship between Gini index and the domestic product per person. However, when they took first differences of the equation they found that the relationship vanishes. Nevertheless, they proceeded to draw conclusion from the equation in levels.

cancel out the stochastic trends of the variables and will generate one which is I(0). When such a combination exists, the I(1) series are said to be co-integrated and their parameters are interpreted as long run parameters.

We determine the order of integration of each series via the Augmented Dickey Fuller (ADF) test of stationarity.[21] The nature of the unit root process may have three forms; therefore the ADF test is estimated under three different null hypotheses as follows:

Y_t is a random walk: $\Delta Y_t = \delta Y_{t-1} + u_t$ (12)

Y_t is a random walk with intercept: $\Delta Y_t = \beta_1 + \delta Y_{t-1} + u_t$ (13)

Y_t is a random walk with intercept

around a stochastic trend: $\Delta Y_t = \beta_1 + \beta_{2t} + \delta Y_{t-1} + u_t$ (14)

Where Y_t can be any variable

the test is applied in levels, first differences and second differences for every equation above in order to determine whether the variables are I(0), I(1) or I(2). In every case two lags are considered.

To test for co-integration between the series, the augmented Engle-Granger (AEG) is conducted.[22] In this case the three forms described from (12) to (14) are also applied in every equation between (5) and (11), that is linear, quadratic and cubic forms in level and log specification plus the Anand-Kanbur form. The co-integration

[21] The ADF test starts with $Y_t = \rho Y_{t-1} + u_t$. For theoretical reasons it is manipulated to obtain $Y_t - Y_{t-1} = \rho Y_{t-1} - Y_{t-1} + u_t = (\rho - 1)Y_{t-1} + u_t$ which can be alternatively written as $\Delta Y_t = \delta Y_{t-1} + u_t$. Under the null hypothesis $\delta = 0$ ($\rho = 1$); that is, there is a unit root – the time series is nonstationary - the estimated t value of the coefficient of Y_{t-1} on (12) follows the τ statistic.

[22] To perform the AEG test, it is necessary to estimate a regression and apply the ADF test on the obtained residuals. Although the AEG test also follows the τ statistic, the ADF critical values are not appropriate; therefore Engle-Granger critical values are required.

test is conducted in levels so as to determine if the residuals are I(0). Initially two lags are applied, if no co-integration is found the number of lags is changed.

Results from the unit root test of stationarity and the test for co-integration are available upon request. We observe that the test equation (12), with no intercept and trend, captures more I(0) variables than the other two specifications. If the variables are first differenced the number of stationary variables rises. Moreover, when the variables are second differenced, almost all series (234 out of 248) are I(2) if the test equation (12) is applied.

In a substantial number of equations (172 out of 217) their linear combination is I(0) when the test equation with no intercept or trend is applied on the residuals and two lags are used. This outcome is consistent with the results obtained from the unit root analysis. The number of co-integrated equations declines when the other two specifications are conducted. It is worth noting that many of the regressions that are not co-integrated become an I(0) linear combination if the number of lags used in the AEG test is changed. It should be added that only in one country, Bolivia, the null hypothesis of nonstationarity in the residuals is not rejected in all its regressions. On the other hand, all the selected models of the remaining countries are co-integrated regressions.

Finally, we transform the explanatory variables to reduce collinearity by expressing them in the deviation form (i.e. deviation from the mean value), as suggested by Draper and Smith (1998: 371-2). In this case, the data are said to be centred around their average value, or often just centred. After applying transformations in quadratic and cubic equations we observe that pair-wise correlation between linear and square regressors decreases substantially. Pair-wise correlations between the linear and cubic regressors and square and cubic regressors also tend to decrease, although in some cases, especially between the linear and cubic regressors, the correlation reduction is moderate. Nevertheless, in any case improvements are achieved (results not reported).

Once the method to reduce multicollinearity is undertaken, we notice that in seven out of 31 countries, the model selected originally is not adequate as some of the coefficients are no longer significant. In these specific country-cases we proceed to select a new equation that satisfies the model selection criteria described so far.

6. Interpretation of the time-series results.

Table 5 sums up the results obtained from the time-series analysis. It indicates the selected model for every country and the year and level of per capita GDP in which the turning points occur. It also shows general results obtained after applying procedures to correct for autocorrelation, to reduce multicollinearity and to test for co-integration.

In addition, Table 5 reveals the existence of different patterns. Five countries follow a linear positive trend along the period. Nine countries show a local maximum, most of them during the early 1970s, but a subsequent local minimum that is followed by a period of rising inequality (max-min trend hereafter); in five countries of this group the final increasing period is longer than nine years; Chile shows a short positive trend over the last years, but it also displays a long increasing trend along the first two decades. Seven countries present a U-shaped relationship, four of them display the minimum turning point along the 1970s, two more in the late 1980s and only Singapore in the 1990s. Six countries initially show an ordinary-U trend, but a subsequent local maximum after the mid 1990s that reverses the period of rising inequality (min-max trend hereafter); in five countries of this group the minimum turning point occurs along the 1970s and therefore the positive trend lasts several years, only in Korea the minimum turning point occurs in the late 1980s and hence the increasing period is relatively shorter.

Table 5. Results from the time-series analysis

			1st turning point		2nd turning point				Linear
Country	Curve shape	Function	Year	PGDP	Year	PGDP	AR	Multicollinearity	combination
Bolivia	Linear (-)	Level					Not corrected		
Malaysia	Linear (-)	Level					(P - W)		I(0)
Egypt	Linear (+)	Level					(P - W)		I(0)
Finland	Linear (+)	Level					Not corrected		I(0)
Greece	Linear (+)	Level					(P - W)		I(0)
Hungary	Linear (+)	Log					Not corrected		I(0)
Turkey	Linear (+)	Level					(P - W)		I(0)
US	Ordinary-U	Log	1971-1972	17,684			(P - W)	reduction	I(0)
Syria	Ordinary-U	Log	1974-1975	579			Not corrected	reduction	I(0)
Colombia	Ordinary-U	Log	1975-1976	1,639			(P - W)	reduction	I(0)
Sweden	Ordinary-U	Level	1978-1979	21,890			(P - W)	reduction	I(0)
Canada	Ordinary-U	Log	1987-1988	19,084			(C - O)	reduction	I(0)
Spain	Ordinary-U	Log	1988-1989	13,298			Not corrected	reduction	I(0)
Singapore	Ordinary-U	Log	1993-1994	22,218			(P - W)	reduction	I(0)
Ecuador	max-min	Log	1971-1972	941	1974-1975	1,293	Not corrected	reduction	I(0)
Japan	max-min	Level	1977-1978	25,423	1987-1988	35,740	(C - O)	reduction	I(0)
Denmark	max-min	log	1971-1972	24,299	1989-1990	31,689	(P - W)	reduction	I(0)
Mexico	max-min	Log	1971-1972	2,371	1989-1990	3,125	Not corrected	reduction	I(0)
Ireland	max-min	Log	1976-1977	9,449	1989-1990	13,915	Not corrected	reduction	I(0)
Mauritius	max-min	Log	1973-1974	1,491	1992-1993	3,235	Not corrected	reduction	I(0)
India	max-min	Log	1982-1983	243	1994-1995	362	Not corrected	reduction	I(0)
Indonesia	max-min	Log	1974-1975	372	1995-1996	1,088	No AR	reduction	I(0)
Chile	max-min	Level	1990-1991	3,317	1995-1996	4,745	Not corrected	reduction	I(0)
UK	min-max	Log	1971-1972	12,116	1994-1995	19,138	Not corrected	reduction	I(0)
Norway	min-max	Level	1974-1975	19,171	1995-1996	34,458	Not corrected	reduction	I(0)
Austria	min-max	Level	1976-1977	20,168	1998-1999	31,355	(P - W)	reduction	I(0)
Netherland	min-max	Level	1977-1978	20,664	2001-2002	32,031	(P - W)	reduction	I(0)
Italy	min-max	Level	1979-1980	14,549	2000-2001	20,955	No AR	reduction	I(0)
Korea	min-max	Level	1987-1988	6,605	1999-2000	12,582	(P - W)	reduction	I(0)
Kenya	NSR								I(0)
Zimbabwe	NSR								I(0)

Notes:

min-max: The first turning point is a local minimum and the second turning point is a local maximum

max-min: The first turning point is a local maximum and the second turning point is a local minimum

NSR: No systematic relationship

P - W: Autocorrelation corrected through the Prais-Winsten method

C - O: Autocorrelation corrected through the Cochrane-Orcutt method

I(0): The linear combination of the variables in the equation is I(0), that is, co-integrated regression

It is worth noting that two countries show a negative linear pattern - Bolivia and Malaysia. However, the former is not the result of economic growth and falling inequality, rather the result of negative rate of growth and rising inequality over the sample. The latter captures a linear trend, but with weak evidence.[23] Finally, in only two countries it is not possible to capture any systematic trend - Kenya and Zimbabwe. Not surprisingly, these countries have shown low rates of growth over the period, which reduces variability in the explanatory variables and makes it difficult to conduct an accurate regression analysis.

Although the time-series approach does not lead to the existence of a common trend to explain the relationship between per capita GDP and income distribution, it shows that a large number of countries tend to increase inequality with economic growth during relatively long periods over the sample. For some countries this positive relationship is permanent and for others starts at different years, only for a few countries the relationship reverse after a prolonged period of rising inequality. It is interesting to note that those developed and developing countries that change towards a positive relationship show minimum turning points over different years; however, most of the developed countries display the trough along the 1970s; whereas most of the developing ones display the trough after the mid 1980s. This, fact suggests that developed economies tended to start a period of rising inequality earlier and this is in keeping with the preliminary evidence in section three and Galbraith and Kum (2003)

Table 6 concentrates the characteristics for every type of relationship captured in the time-series analysis. It has been noticed that countries following a min-max trend mainly display the local minimum along the 1970s and the local maximum after the mid 1990s, only Korea displays a latter trough; in this sense, the average minimum and maximum turning points occur around 1978 and 1998 respectively. The

[23] The coefficient on the explanatory variable in the linear equation for Malaysia is just statistically significant at 10 % and the F test of overall significance is just satisfied also at 10 %. This country-case also captures a cubic relationship, but it vanishes when we correct for multicollinearity. By analysing raw data we observe that the inequality-growth relationship in Malaysia rather follows a cyclical pattern with several turning points over the sample that might be modelled as a 4^{th} degree polynomial.

trough across countries following the max-min trend mainly occurs over the late 1980 and early 1990s and the average is around 1990. The trough for those countries that capture the U shape is more diverse as it can occur either along the 1970s or in the late 1980s mainly, the average turning point lies around 1981, but this figure is not representative of an overall trough due to diversity across countries following this trend. It should be stressed that those countries showing evidence of reaching a peak after the mid 1990s, present a positive relationship between growth and inequality over a long period that starts mainly along the 1970s; in addition, their positive trend period tends to start earlier than those countries which continue to show an increasing pattern after the mid 1990s.

Table 6. Characteristics by type of relationship and level of development

Number or countries	Relationship	Trade Growth % 1970-1998	FDI Growth % 1970-1998	Inflation SD 1970-1998	Governance 1996	Turning point		Turning point	
						Year	Location	Year	Location
2	Linear (-)	1.09	11.76	1091.53	0.18				
5	Linear (+)	1.58	20.60	11.24	0.50				
7	Ordinary-U	1.48	11.21	5.63	0.88	1981	min		
9	max-min	1.90	10.58	24.37	0.54	1977	max	1990	min
6	min-max	0.54	3.54	4.61	1.27	1978	min	1998	max
2	NSR	1.52	8.06	9.42	-0.42				

Notes:
min-max: The first turning point is a local minimum and the second turning point is a local maximum
max-min: The first turning point is a local maximum and the second turning point is a local minimum
NSR: No systematic relationship

The economic liberalisation process has been conducted through two main stages, especially in developing countries. The first one has been mentioned earlier and involves the implementation of a set of economic policies, which is in essence the orthodoxy that dominated the 1980s and early 1990s. The second stage has emerged since the late 1990s; it emphasises a set of socio-political norms advocating principles of governance based on efficiency and effectiveness of the modern state and is an attempt to socialise and humanize the earlier technocratic elements.[24] It should be added that macroeconomic stability is considered an essential requisite for the operation of markets and free mobility of capital.

On this basis, we explore how trends in the growth-inequality relationship can be associated with different policies and norms involved in the economic liberalisation process. So as to represent the set of socio-political norms, the analysis includes the average of aggregate governance indicators for the year 1996. The set of economic policies is represented by the annual rate of growth of trade volume and FDI inflows. Fiscal discipline and macroeconomic stability are represented through standard deviation of inflation.[25] A simple average of every indicator is worked out for every group of countries classified according to the different patterns captured in the time-series analysis. Results are illustrated in Table 6.

We notice that those countries which have achieved decreasing inequality after the mid 1990s (the min-max trend), present a higher governance indicator compared to the rest of countries. Their corresponding governance indicator is 1.27, whereas it is 0.50 for those countries that have experienced a continuous upward trend, and 0.54 and 0.88 for those countries that have shown max-min and ordinary-U patterns respectively. Furthermore, countries in the min-max group present a lower standard

[24] The original set of economic norms is also called the Washington Consensus or First Generation Reforms, see Williamson (1990) and Ortiz (2003). The set of socio-political norms, often also called the Post Washington Consensus or Second Generation Reforms focuses on issues of civil society participation, social capital formation, capacity building, safety nets, transparency and accountability, institution building, among others). For further discussion see Higgott (2000).

[25] The aggregate governance indicator is obtained from the World Bank website. It is the average of six indicators measuring the following dimensions of governance: voice and accountability, political stability and absence of violence, government effectiveness, regulatory quality, rule of law, and control of corruption. Its score lies between -3.0 and 3.0 with higher score corresponding to better governance. Trade volume is the sum of exports and imports of goods and services measured as a share of GDP, inflation reflects the annual percentage of change in consumer prices, the source is World Bank (2002). FDI inflow is measured as a percentage of GDP and is obtained from UNCTAD (2003) and World Bank (2002)

deviation of inflation and lower rates of growth in terms of trade volume and FDI in relation to the rest of the countries.

7. Concluding remarks.

The panel data analysis captures a general pattern that resembles a U-shaped curve, in which inequality first decreases, reaches a trough and then increases with economic growth, and the pattern seems to apply in both developed and developing countries. When we test for the existence of cycles this approach shows week evidence of a local maximum at further stages of development.

The time-series analysis is carried out across different country-cases in order to date turning points and to explore further intertemporal evidence. This approach does not indicate a single trend to explain the relationship between inequality and per capita GDP; however, it shows that a substantial number of countries comprised in the analysis increase inequality with economic growth during relatively long periods over the sample. For some countries this positive trend is permanent and for others begins at different years with the presence of a minimum turning point; in addition, a group of six countries show evidence that the trend can reverse at further stages of output as they capture the presence of a later peak. The time-series analysis also shows evidence that periods of rising inequality tend to start earlier in developed economies than in developing ones

The implementation of outward-oriented policies started in some economies during the late 1970s, notoriously the US and the UK, while other countries adopted them along the 1980s. In this context, the results suggest that a positive relationship between growth and inequality started in some countries before they embarked in structural reforms. As a result, other factors like stagflation in the 1970s due to oil price shocks or rising interest rates and the debt crisis in 1982 (Galbraith and Kum, 2002: 14) might have contributed to drive inequality up. Moreover, we observe that the rise of inequality continues along the sample, which suggests that the surge of market liberalism did not improve income distribution; in contrast, it seems to reinforce the change towards a positive relationship between growth and inequality. Through the ascendancy of market-oriented ideas it was expected to boost economic growth to reduce inequality and therefore to achieve a negative relationship between this two variable; however, the results undermine these expectations and their theoretical foundations.

On the other hand, the results suggest that a period of rising inequality is likely to reverse over the long-run as some of the countries that capture the minimum turning point in early years show evidence of improving income distribution in recent years. This finding is consistent with previous studies claiming that in an environment of greater competition income distribution may widen in an initial period due to changes and adjustments in markets; however, as the period of adjustment continues market forces react, individuals adapt and the levels of inequality may began to lessen (Jacobsen and Giles, 1998: 419-20). We also find that time is not the only factor affecting this process, because macroeconomic stability, a good level of governance and gradual expansion of openness are additional factors associated to the fall in inequality in a latter stage. In this respect Chu and Tanzi (1998: xiv) argue that sound macroeconomic and structural policies are consistent with sustainable economic growth and improved equity over the long term; in addition, Angeles-Castro (2005) shows that those countries which are associated with a high governance indicator and a more stable economy are likely to mitigate the adverse effect that FDI might cause and are likely to obtain benefits from trade in terms of income distribution.

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