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THE RELATIONSHIP BETWEEN ECONOMIC GROWTH AND INEQUALITY: EVIDENCE FROM THE AGE OF MARKET LIBERALISM*

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RESUMEN

Mediante un panel de países selectos, este artículo muestra que la relación desigualdad-crecimiento sigue una forma de U durante el periodo 1970-98, en el cual la desigualdad inicialmente decrece y posteriormente se incrementa con el crecimiento económico. La evidencia también muestra que este patrón creciente puede revertirse a mayores niveles de ingreso. Usando series de tiempos se ilustra que varios paises presentan un punto mínimo en diversos años a lo largo del periodo, mientras que otros siguen una tendencia positiva permanente. Adicionalmente se observa que, un pequeño grupo de países revierte la tendencia de desigualdad creciente a niveles altos de ingreso y despliegan un punto máximo a finales del periodo. Estos países están asociados con estabilidad macroeconómica, eficiencia gubernamental y expansión

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moderada del comercio y la IED. Entonces, la relación desigualdad-crecimiento durante la era del liberalismo de mercado tiende a cambiar hacia una relación positiva, aunque puede revertirse posteriormente.
Clasificación JEL: C22, C23, O15, O57
Palabras clave: Distribución del ingreso, crecimiento económico, modelos dinámicos de datos en panel, análisis de series de tiempo

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 rises 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.
JEL classification: C22, C23, O15, O57
Keywords: Income distribution, economic growth, dynamic panel data models, time-series analysis

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 in a global scale since the late 1970s and the need to assess their performance 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 adopted a panel data approach by using the Deininger and Squire (1996) (D & S thereafter) data-set, which improves coverage and extends the period slightly in relation to previous data-sets used in cross-country analysis. 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 still limited and unbalanced (Atkinson and Brandolini 2001, Galbraith and Kum 2002).

In order to overcome the problem of sparse coverage presented in the D & S data set, different panel studies grouped the data in averages or intervals or selected a sub-sample of countries, and obtained diverse, even contradictory, forms of the inequality-growth relationship. Ram (1997) used a selected sub-sample of 19 developed economies and found an ordinary-U curve, whereas Barro (2000) derived support for an inverted-U curve by organising the sample in ten-year intervals. Deininger and Squire (1998) showed that the inequality-growth relationship is downward sloping by grouping the data in decadal averages; in contrast, Forbes (2000) found that it is upward sloping when the data are restricted to a subset of five-year intervals. In all these cases there is a risk of bias in the construction of the averages and intervals or in the selection of the subsets, and the results seem to be affected by the way the data are organised.

The World Income Inequality Database version 1.0 (WIID) (UNU/WIDER 2000) has also become a source of information for panel data studies of the inequality-growth relationship (Grün and Klasen 2003). The D & S data-set was its starting point and represents half of its observations. However, WIID still suffers from sparse coverage when "reliable filters" are applied or more homogeneous sub-samples are obtained.

Panel data analysis could also be undertaken by means of two additional sources available in the literature, the Luxembourg Income Study and the UTIP-UNIDO (UTIP 2002) 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 data-set in order to replicate the coverage of the latter with estimated measures of household income inequality. The EHII takes into account the relationship between industrial pay inequality and household income inequality, and also takes into account an additional set of variables in order to adjust for the different weight of manufacturing across economies. The result is a data-set with large coverage and representation of overall household inequality, which is aimed at overcoming inconsistencies in D & S. Therefore, this recent source of information offers the possibility to explore further evidence of the trends in distributional changes.

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. Due to the relatively large coverage of the sample, we apply the observations directly and therefore eliminate any risk of bias that might result from the construction of averages and intervals or from the selection of sub-samples. Moreover, the coverage of the data allows us to construct a balanced panel consisting of 31 countries and 899 observations over the same period, the related literature had not analysed balanced samples before. 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 a generalised method of moments (GMM) estimation for dynamic panel data models proposed by Blundell and Bond (1998). The large coverage of the sample, across countries and over time, helps to improve the precision of the estimators.

The literature has conventionally applied quadratic equations. Although we follow this approach, we also extend the model into 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.* (2002) 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. 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. Early studies reporting turning points in the trend of inequality did not address the issue of non-stationarity of the variables and did not test 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 suggests that the presence of a local maximum over the longer-run depends on the composition of the panel. 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, Cornia and Court 2001, 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 two discusses orthodox assumptions and expectations; it also provides a preliminary analysis of the evolution of growth, income distribution and the relationship between these two variables since 1970. In section three and four the panel data analysis and the time-series analysis are undertaken respectively. The interpretation and discussion of results are presented in section five. Finally concluding remarks are provided in section six.

2. ASSUMPTIONS, EXPECTATIONS AND PRELIMINARY EVIDENCE

2.1 THE ORTHODOX VIEW

Standard neoclassical theory asserts that economic liberalisation offers 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.

Since the early 1980s the prevailing global political economy has added impetus to outward-oriented policies and has discouraged attempts of protectionism. In this context, the dominant policy prescription has mainly involved liberalisation of trade, investments and the labour market, besides privatisation and fiscal discipline. Under these circumstances and from an orthodox perspective, we may expect improvements in the global distribution of income and an inverse relationship between the level of income and inequality.

On the other hand, before the 1980s the prevailing economic policies of the post-war period can be summarised as protectionist strategies and inwardlooking development 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. Hence, over these years and from a neoclassical viewpoint, we may expect that inequality rises as income increases, because market distortions and government interventions are usually deemed inefficient and inequitable (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 may be depicted by an inverted-U curve with the turning point somewhere after the early 1980s.

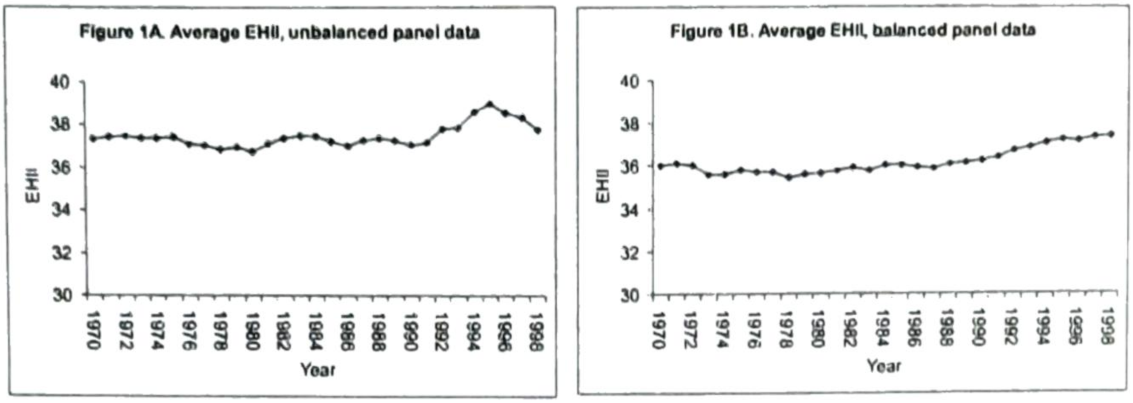
2.2 PRELIMINARY EVIDENCE

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 can be the result of variable lags in reporting underlying data to UNIDO. As a matter of fact, the number of countries contained in our sample in the last years drops

FIGURE 1

Average values of EHII for each year



substantially. Consequently, the decreasing inequality illustrated in Figure 1 by the end of the period, maybe caused by gaps across the panel.[1]

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 around 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

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

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.

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 samples respectively. The variable on economic growth is annual percentage growth rate of GDP based on constant U.S. dollars; the source is World Bank (2002). The rate of growth appears to be unstable and tends to slow down over the whole period, as it displays a downward trend, and the composition of the panel does not seem to affect this pattern.[2]

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 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

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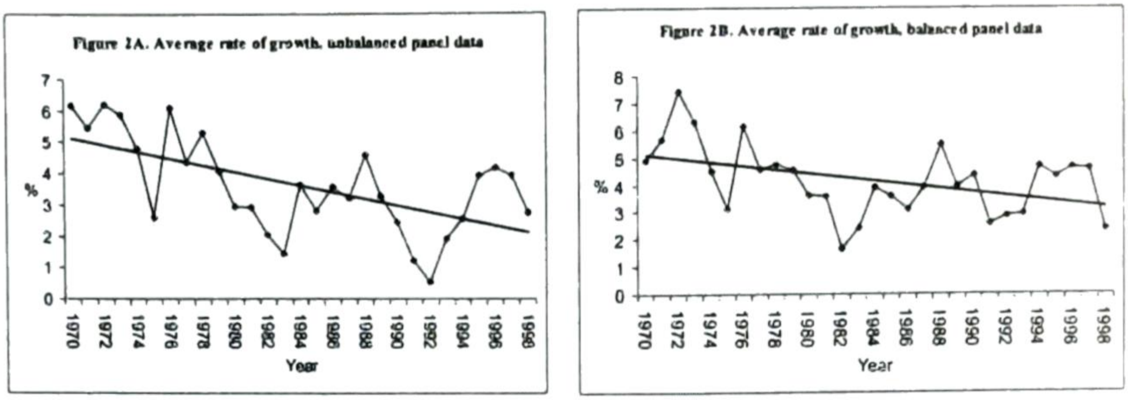
[2] The analysis is extended by plotting the unbalanced and balanced sample weighted by GDP and GDP per capita and is conducted for both developed and developing economies. The countries are divided according to the criteria already explained. In any case it is confirmed that the rate of growth is unsteady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

FIGURE 2

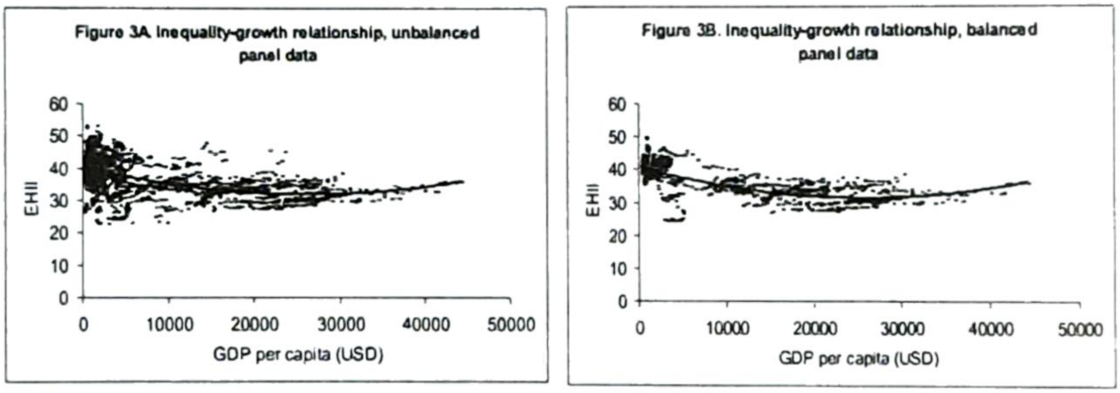
Average values of rate of growth for each year



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]

FIGURE 3

Inequality-growth relationship



[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.

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, a likely cause of this trend is that low income countries are normally associated with higher levels of income inequality.

On the other hand, it has been already argued that from the perspective of standard theory we may expect that the inequality-growth relationship follows an inverted-U curve over the period comprised in the sample 1970-1998. 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. When we fit the balanced and unbalanced data sets, for both developed and developing economies plus the overall samples, to three different equations –Linear, Logarithm and Polynomial– we find that in the six samples, the Polynomial equation following an inverse U-shaped curve displays the highest *R* square. However, we have to keep in mind that the number of regressors should be the same for the comparison to be appropriate, as the *R* square tends to rise with the number of regressors, and in this case the polynomial specification has one more. In this sense, we obtain the Akaike Information Criteria (AIC) and the Schwarz Information Criteria (SIC) and observe that the polynomial specification has the lowest values in every of the six samples.[4] In the following section we turn to quantitative methods to explore in more detail the possible existence of a systematic and convincing relationship between inequality and income level over the last decades.

[4] The AIC is a measure of the goodness of fit of an estimated statistical model. The AIC methodology finds the model that best explains the data with a minimum of parameters. It imposes a penalty for adding regressors to the model. In comparing two or more competing equations, regardless the number of regressors, the equation with the lowest value of AIC is preferred. Like AIC, the lower the value of SIC the better the model.

3. PANEL DATA APPROACH

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

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

in which *EHII* is the inequality measure and *Y* is GDP per capita in 1995 U.S. dollars.[5] The subscripts *i* and *t* indicate country and year respectively. The error term u_{it} is assumed to satisfy white noise assumptions, that is independently and identically distributed with 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.[6]

It is worth nothing that by applying quadratic equations we expect to find the corresponding turning point within the sample range because the preliminary analysis provides some evidence suggesting that the pattern of inequality tends to change across the observations. The presence of a maximum turning point within the sample range ($\beta_2 < 0$) would confirm orthodox assumptions and expectations. On the other hand, the presence of a real rather than an inverse U-shaped relationship ($\beta_2 > 0$) would confirm recent findings of rising inequality since the 1980s. It should be added that, depending on the functional form, inflection points might occur outside the range of the observations producing

[5] Some previous studies in the related literature have investigated a reverse causality, that is to say the effect of inequality on growth (Odedokun and Round 2001, Banerjee and Duflo 2003). This paper focuses on the effect of growth on inequality, because is aimed at testing if the expansion of output can bring a distributional effect during the era of market openness.

[6] 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.

a relationship with either a positive or negative slope over the whole range of the sample. In the next sections we will test equations with different functional forms in order to confirm or reject the presence of a turning point within the sample.

3.1 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) (LM) rejects the standard OLS assumption that the intercept value is the same across countries, and therefore there are country-specific effects in the model.[7] Secondly, the Hausman test (1978) suggests that the country-specific effects are correlated with the regressor in the equation. The no correlation assumption is an important pillar of the random-effects model (REM), but in this case is violated.[8] 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, including the level of GDP per capita in which the turning points occur, are reported in table 1 from column 1 to column 3. In order to determine the value of each turning point we follow Hsing and Smyth (1994) and Jacobsen and Giles (1998) procedure. The procedure is based on estimated parameters, takes the first derivative of the dependent variable with

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[7] The LM test, based on the OLS residuals and under the null hypothesis: $\alpha_i = \alpha$, that is, the classical regression model with a single constant term is appropriate, is distributed as a χ^2 with one degree of freedom (Greene 2000, 572-3). In this case, 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.

[8] 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.

respect to Y and sets it equal to zero. It is interesting to note that the coefficient of Y^2 is significant and positive in the first two columns and the corresponding inflection points are located at different levels of income, but both within the sample range.[9] These two functional forms capture a U-shape where income inequality first diminishes and then is found to rise with increasing output. On the other hand, the coefficient of Y^2 in the FEM, presented in column 3, is also significant and positive, but this model captures the turning point outside the range of observations and displays a positive and monotonic relationship over the sample. This result suggests that inequality has tended to increase with economic growth along the whole period.

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 (AR) test on the error term u_{it} should be available. We find that the first and second-order AR tests, conducted on the fixed-effects regression and reported in column 3 of table 1 are not satisfied.[10] So as to address this problem, it is required to explore the possibility that autocorrelation may arise owing to model misspecification, 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:

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

On the other hand, the incorporation of a lagged dependent variable introduces a source of persistence over time, correlation between the error term u_{it} and the right hand regressor $EHII_{it-1}$. In addition, DPDMs are characterised by individual

[9] In the sample, the minimum value is 84 and the maximum value is 44,206, expressed in 1995 U.S. dollars.

[10] 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.

effects η_i that result from heterogeneity among the individuals.[11] Therefore, it is necessary to adopt different testing and estimation methods in this case.

So as to estimate the model consistently and efficiently we use a generalised method of moments estimation for DPDMs proposed by Blundell and Bond (1998). Initially, the estimation procedure eliminates country-effects (η_i) by expressing equation (2) in first differences as follows:

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

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

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

that is, the error term in first differences is uncorrelated with lagged levels of $EHII_{it}$. This econometric technique applies instruments to control for endogeneity of the lagged dependent variable in first differences using lagged levels of $EHII_{it}$. The endogeneity is reflected in the correlation between the error term in the equation in first differences and the lagged dependent variable. The GMM estimator obtained from this procedure was proposed by Arellano and Bond (1991) and is known as the difference estimator.

On the other hand, the GMM, estimator obtained after first-differencing, has been found to have large finite sample bias and poor precision (Blundell and Bond 1998, 115-6). Both drawbacks are attributed to the problem of weak instruments, since lagged levels of the series represent weak instruments for the first difference. Blundell and Bond justified the use of an extended GMM estimator in order to improve the properties of the standard first-differenced GMM estimator. The extended estimator is constructed on the basis of the following moment condition:

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

[11] For an elaboration in this point see Baltagi (2001, 129-30).

TABLE 1

Unbalanced panel data

	(1) OLS	(2) REM	(3) FEM	(4) sys-GMM	(5) sys-GMM	(6) sys-GMM 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	2,289	2,289	2,289	2,173	2,173	2,173
Countries	116	116	116	116	116	116
Min turning Point	26,269	10,322	-10,452	14,387	12,394	11,505
Max turning Point					49,470	38,265

Notes: Dependent variable: EHII, p values in parenthesis, * Significant at 5%,
** Significant at 10%.

that is, there is no correlation between the country-specific effect and lagged differences of $EHII_{it}$. In short, this econometric technique applies two sets of instruments. The first is lagged levels of $EHII_{it}$ as instruments for equations in first differences and the second is lagged differences of the endogenous variable as instruments for equations in levels. This method, called system GMM estimator (sys-GMM), reduces the finite sample bias and improves the precision.

In this econometric technique, it is assumed that the disturbances u_{it} are serially uncorrelated. In this sense, there should not be evidence of second-order serial correlation in differenced residuals (u_{it} - u_{it-1}), whereas there should be evidence of first-order serial correlation (Doornik et al. 2002, 5-8). The assumption of no correlation in the disturbances u_{it} is important because the consistency of the GMM estimators relies on the fact that $E[\Delta u_{it}\, \Delta u_{it-2}] = 0$. Consequently, tests of first-order and second-order autocorrelation need to be conducted in the first-differenced residuals. Furthermore, a Sargan test of overidentifying restrictions, proposed by Arellano and Bond (1991), is reported so as to assess the validity of the instruments.

Column 4 of table 1 provides the results obtained from the sys-GMM 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},[12] while the Sargan test of overidentifying restrictions is unable to reject the validity of the instruments.[13] Under these circumstances, it is possible to treat the results as valid. In this case, we find that the coefficient of Y^2 is positive and significant and the minimum point corresponds to $14,387. Hence, this sys-GMM regression indicates that the relationship between growth and inequality tends to follow a U-shaped curve over the sample.[14]

The panel used in this study is larger than those used in previous studies exploring the relationship between economic growth and income inequality. Due to an improved coverage of the panel we apply the observations directly and do not organise them in averages or intervals or in sub-samples of countries,

[12] 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 0.804; therefore, the test fails 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.

[13] 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.

[14] This sys-GMM 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.

as in previous studies. This approach eliminates the risk of bias that might result from the selection of subsets or from the construction of averages and intervals. The large panel helps to improve the precision of the GMM estimators, as they generally perform better with a larger N and T (Judson and Owen 1999).

3.2 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 point. In this sense, equation (3) is extended into a third-degree polynomial by adding the cube of income per capita as follows:

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

The results from the sys-GMM 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. 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.[15] It should be added that the maximum is located outside the range (see footnote 9); consequently, this regression suggests that a maximum turning point might occur, but at high levels of output beyond the range of observations; however, further considerations need to be done.

[15] 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.

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 n^{th} 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.[16]

With the above in mind, an orthogonal transformation as in Doornik *et al.* (2002, 25, 35), is performed to reduce multicollinearity. This transformation takes each observation in deviation from the future means, together with a standardisation. 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. The value at which the maximum turning point is located in the orthogonal equation is lower than that of the original sys-GMM equation and it is within the sample range. Nevertheless, 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 sys-GMM regressions in quadratic and cubic equations 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 rejects the null hypothesis that the coefficient on Y^3 is equal to zero. These findings suggest that the relationship between inequality and growth follows a U-shaped curve during the age of economic liberalisation, but inequality tends to decline with economic growth after a prolonged period of time, independently of the level of development.

[16] 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).

TABLE 2
Unbalanced panel data (developed and developing countries)

	Developing countries			Developed countries		
	sys-GMM	sys-GMM	sys-GMM Orthogonal	sys-GMM	sys-GMM	sys-GMM Orthogonal
$EHII_{t-1}$	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, *p* values in parenthesis, * Significant at 5%, ** Significant at 10%.

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.[17] We now test the existence of a cyclical pattern through a balanced panel data-set.

[17] Some authors have demonstrated that GMM estimators generally perform better with a relatively large *N* (Blundell and Bond, 1998; Judson and Owen, 1999) as noted before. 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 way, the size of *N* does not drop drastically in any sub-sample. We conduct sys-GMM 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. The results are available upon request.

3.3 BALANCED SAMPLE

We apply the balanced panel data-set to explore the income-inequality relationship. Due to sparse coverage of previous data on income distribution, balanced panels had not been used before in the related literature. The results obtained from the overall sample are reported in table 3. The sys-GMM 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 sys-GMM method nor the orthogonal transformation captures a cyclical pattern when the equation is extended into a third degree polynomial, since the 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.

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 longer-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 depends on the composition of the panel.

3.4 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,

TABLE 3

Balanced panel data

	sys-GMM	sys-GMM	sys-GMM 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,
p values in parenthesis, * Significant at 5%,
**Significant at 10%

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 cross-country samples for unbalanced and balanced data-sets.[18] Equations in levels, logs and the Anand-Kanbur specification are applied in every sample. Results are illustrated in table 4.

[18] 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 10%

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 equations in levels capture an ordinary-U pattern, but the significance of the coefficients is ambiguous.

This approach derives weak empirical support for the Kuznets hypothesis. Moreover, globalisation does not seem 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.

Although the panel data analysis determines the level of income in which the inflection points occur, 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 (developed and developing countries), 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 levels of GDP per capita in which turning points occur.

4. 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.* (2002, 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. 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 transformations 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[19] 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 –AIC, SIC and Ramsey's

[19] 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" in other words the additional coefficient is equal to zero.

RESET test (RRT)–.[20] 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 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 of the tests and regressions 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.

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[20] See footnote 4 for an explanation of AIC and SIC. 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.

Firstly, it is worth noting that the Durbin-Watson d test and the Breusch-Godfrey (BG) test[21] 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 (C-O) as in Hsing and Smyth (1994) and the Prais-Winsten method (P-W).

If autocorrelation persists, we test the possibility that it may arise due to model misspecification by adding a lagged dependent variable. However, the inclusion of a lagged dependent variable introduces a source of persistence over time, 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 the (C-O) and the (P-W) methods are able to correct for autocorrelation in 14 out of 27 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 in order to keep the

[21] 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 H_0: $\rho_1 = \rho_2 = 0$; that is, there is no serial correlation of first and second-order.

systematic relationship, and allow for autocorrelation only in these country-cases where autocorrelation is persistent.[22]

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 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.[23] 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.

[22] 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.

[23] 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 non-stationary– the estimated t value of the coefficient of Y_{t-1} on (12) follows the τ statistic.

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 cointegration between the series, the augmented Engle-Granger (AEG) is conducted.[24] In this case the three forms described from equations (12) to (14) are also applied in every equation from (5) to (11). The cointegration test is conducted in levels so as to determine if the residuals are I(0). Initially two lags are applied, if no cointegration is found the number of lags is changed. The results from the unit root test of stationarity and the test for cointegration 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 non-stationarity 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 (that is 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

[24] 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 t statistic, the ADF critical values are not appropriate; therefore Engle-Granger critical values are required.

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.

Once the method to reduce multicollinearity is undertaken we notice that in some 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.

5. 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 cointegration.

From table 5 we observe 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 curve, 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.

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.[25] 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 reverses 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

[25] 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 4th degree polynomial.

TABLE 5

Results from the time-series analysis

Country	Curve shape	Function	1st turning point Year	1st turning point PGDP	2nd turning point Year	2nd turning point PGDP	AR	Multicollinearity	Linear 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)
Netherlands	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 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. It should be stressed that the positive trend of these countries tends to start earlier than those countries which continue to show an increasing pattern after the mid-1990s. The 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.

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 humanise the earlier technocratic elements.[26] 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 an average of aggregate governance indicators for the year 1996. The set of economic policies is represented by the annual rate of growth

[26] 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).

TABLE 6

Characteristics by type of relationship

Number of countries	Relationship	Trade Growth % 1970-1998	FDI Growth % 1970-1998	Inflation SD 1970-1998	Governance 1996	Turning point Year	Turning point Location	Turning point Year	Turning point 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

of trade volume and FDI inflows. Fiscal discipline and macroeconomic stability are represented through standard deviation of inflation.[27] 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 the countries. Their corresponding governance indicator is 1.27, whereas it is 0.50 for those economies that have experienced a continuous upward trend, and 0.54 and 0.88 for those that have shown max-min and ordinary-U patterns respectively. Furthermore, countries in the

[27] The aggregate governance indicator is obtained from World Bank (2003). 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).

min-max group present a lower standard deviation of inflation and lower rates of growth in terms of trade volume and FDI in relation to the rest of the countries.

6. 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, the results seem to depend on the composition of the panel, as only the unbalanced sample shows evidence of a local maximum at further stages of development.

The time-series analysis is carried out across countries with continuous observations 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 shows 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, especially in 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, 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. In fact, globalisation is sometimes presented in the relevant literature as a cause for the deterioration of income distribution in recent decades. In this respect Cornia and Court (2001,1) argue that liberal economic policy regimes is a crucial factor causing the widespread surges in inequality around the world and Smeeding (2002, 28) holds that globalisation is one force widening income inequality in countries.

On the other hand, a substantial part of the literature on changes in income distribution during the last two decades attributes the rise in inequality to the skill-biased technological change (SBTC) (Berman *et al.* 1998, Acemoglu 2002). According to this argument, countries tend to experience a fall in relative demand for unskilled labour and an increase in that for skilled labour, due to an acceleration of technical change over the past few decades, this process is expected to exacerbate inequality.[28] Both explanations (globalisation and technical change) dominate the relevant literature (especially for the study of developed countries) and have been dubbed the "transatlantic consensus".[29]

However, Singh (2001) argues that this consensus is unsatisfactory and underlines an alternative analysis, which emphasises the role of institutions (unions, minimum wages), macroeconomic conditions, and social norms.

Through the ascendancy of market-oriented ideas it was expected to boost economic growth, to reduce inequality, and therefore to achieve a negative relationship

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28 The SBTC hypothesis is oriented to explain the expansion of the wage gap in developed countries over the last two decades; nevertheless, evidence from the developing world is also consistent with it, despite the opposite Stolper-Samuelson prediction (Berman *et al.* 1998).

29 The skill enhancing trade hypothesis is oriented to explain the expansion of the wage gap in developing economies (Robbins 1996). It is based on arguments that, to some extent, can be similar to those used in the "transatlantic consensus" for developed countries, as it claims that economic liberalisation and the intrinsic adoption of new technologies are an important reason of income dispersion in the developing world. According to this approach, the inflowing technology resulting from economic liberalisation is accompanied by a change in demand in favour of skilled workers, and this shift can outweigh the reduction in skilled labour demand predicted by standard trade theory (Arbache *et al.* 2004).

between these two variable; however, the results undermine these expectations and reveal that there is room for alternative explanations like those mentioned above.

On the other hand, the results suggest that a period of rising inequality is likely to reverse over the longer-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 Tanzi and Chu (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.

The fact that some countries can reverse inequality over the longer-run represents a challenge for the SBTC hypothesis. In this respect, Card and DiNardo (2002) hold that a fundamental problem for this approach is that wage inequality has tended to stabilise in some countries, like the US, in the 1990s, despite continuing advances in technology.[30] Moreover, the evidence in this

[30] Stabilisation of wage inequality is a puzzle for the SBTC hypothesis; nevertheless some authors have derived arguments that could explain this problem. Pissarides (1997) states that in developing countries either the new technology or the importation and assimilation process can be skill-biased and give a temporary and relative advantage to skilled labour that leads to higher relative wages only during the period of transition towards a higher level of technology. He also argues that the response of relative supply of skilled and unskilled labour to trade openness can also explain a temporary increase of wage differentials. In addition, Goldin and Katz (1998) hold that within firms, demand for skill rises when new technologies are introduced, but it declines once the other workers have learned to use the new equipment; hence, this process can follow a technological cycle.

paper suggests that macroeconomic conditions and social norms can be associated to changes in income distribution, as suggested by Singh, at least in the longer-term.

Hence, the inequality-growth relationship has tended to change towards a positive one during the last three decades, and the possible explanations of this trend can be widespread economic liberalisation, and technological change and the resulting acceleration in skill bias. Nevertheless, additional factors are also relevant to explain the peak in this relationship in some countries since the mid-1990s.

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