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Trends, Breaks and Persistence in Top Income Shares

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Abstract

We examine the top income share data of a sample of countries to empirically examine for the presence of structural breaks, linear trends and persistence. The analysis of the data is carried out separately for each individual country using novel econometric procedures that are both appropriate and robust. Various theories have been put forward to explain the causes of structural breaks in long run data, such as the introduction of assembly lines from the time of World War I and the ICT revolution. What we find is that there is no clear evidence that Anglo Saxon countries have similar trends as opposed to Nordic, Continental European or other Asian countries. The results are varied and no clear conclusion can be made. Further, the top income share data is found to be highly persistent, suggesting that shocks to the data are likely to be long-lived.

Keywords: Unit Roots, Top Income Shares, Structural Breaks

JEL classification: C22, C32, N30

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1. Introduction

In 1953, Simon Kuznets and Elizabeth Jenks published *Shares of Upper Income Groups in Income and Saving*, where they produced the first comparable long-run income distribution series. One year later, in his famous presidential address to the American Economic Association, Kuznets first addressed the 'character and causes of long-term changes in the personal distribution of income' (Kuznets, 1955). In his speech, Kuznets emphasized the need to develop proper definitions of inequality and outlined the properties of the data required for the study of inequality development over time. Since then, many efforts have been made to provide inequality data. While the primary focus has been on building micro-panel data sets based on national household surveys, this focus on microdata and the consequent lack of long-run data meant that the long-run analysis of inequality remained under-researched. This however changed, when Piketty (2001, 2003) constructed a series of top income shares in France, spanning the entire twentieth century. This led to a building up of interest in the long-run developments of inequality, and similar efforts of building long data sets for many other countries. The data on top income shares has been used in many studies to draw attention to the rich and their income levels by uncovering the top income distributions. This approach contributes to the set of studies that have focussed on top income distributions rather than the overall measures of inequality such as the Gini. As pointed out by Roine and Waldenström (2015), top income shares are not just about the rich and, in the absence of available alternatives, they provide a useful general measure of inequality over time, even if they say nothing meaningful about the changes happening within the lower part of the distribution.

In his book, Piketty (2001) documents that for France inequality increased from the beginning of the twentieth century to World War I, after which it decreased until the late 1970s, and then the trend started to rise again. This study has proven to be highly influential, prompting a range of studies investigating the trends in top income shares in other countries such as UK (Atkinson 2003), USA (Piketty and Saez 2003), continental Europe and the developed countries (Atkinson and Piketty 2007), and emerging market countries (Atkinson and Piketty 2010, Alvaredo et. al. 2013).

In general, the studies find that the measures of inequality have differing trends depending on the period of time and the associated underlying economic conditions. For example, the causes for decline in top income shares over the first half of the twentieth century have been attributed to the loss of large amounts of wealth to capital owners caused by exogenous shocks, thereby decreasing their income share (Roine et. al. 2009). This decline in wealth continued to fall decades after World War II due to high taxes. However, after 1980 it has been argued that that top income shares have increased in Anglo-Saxon countries but not in Continental European countries (Roine et. al. 2009), and this has not been due to increases in capital incomes but rather due to increased wage inequality (Piketty and Saez, 2006).

There have been calls for exploiting the dynamics of long run inequality data across time paying attention to the variation of countries, using appropriate econometric methods to determine whether structural breaks are present in the trend, as well as the underlying signs and magnitudes of trend (or no trend) in the regimes demarcated by the breaks. However, a problem with such studies that identify breaks is the nature of persistence of the data in long time series is ignored, leading to potentially misleading results. In this paper we address this gap by making a robust test for trends, structural breaks and persistence in top income shares for eleven countries, which include Anglo Saxon countries, continental Europe and Asian countries. The analysis of breaks, trends, and persistence in the data is carried out separately for each individual time series. What we find is that there is no clear evidence that Anglo Saxon countries have similar trends as opposed to Continental European or other Asian countries. The results are varied and no clear conclusion can be made. What we argue for is that countries cannot be readily aggregated in to groups such as Anglo-Saxon or Nordic etc., as each individual country has different dynamics. Further, if regression based analysis on long run top income share data is to be carried out, then the country specific characteristics may need to be accounted for given the possibility of structural breaks and the underlying persistence that are found to exist in the data.

The rest of the paper is organized as follows. The next section presents the literature review and discusses some methodological issues regarding the estimation of trends and breaks in inequality. Section 3 presents the testable hypotheses that underlie the observed trends in inequality and explains the econometric methodology used to test these hypotheses. Section 4 reports the empirical results. The final section concludes.

2. Literature Review

Atkinson and Piketty (2007) argue that the top 1% income share maintained a relatively high level up until World War I. This was followed by a drop that took place during World War II and the Great Depression, although the fall in top income shares was more gradual for those countries that stayed out of World War II. From then on, the top income share declined steadily over the twentieth century up until around 1980, when it began to increase again. According to Atkinson and Piketty (2007), Anglo-Saxon countries (such as Australia, New Zealand, the USA) have experienced a substantially greater increase than non-English speaking countries (such as France, Sweden, Norway, Finland, the Netherlands).

Despite the strong emphasis in the top income share literature on the diverging patterns between Anglo-Saxon countries and continental Europe, recent studies covering many other countries have provided deeper insights into the long-run evolution of inequality. Atkinson and Piketty (2010) and Atkinson et. al. (2011) provide evidence on inequality trends across six different groups of countries; namely, Anglo-Saxon, Continental European, Nordic, Asian, African and Latin American countries. According to Roine and Waldenström (2015), almost all countries exhibit a secular decline in top income shares over the twentieth century. Divergences within country groups appear, however, from 1980 onwards, with substantial increases for the Western English-speaking countries as well as China and India; a modest increase in some Nordic countries and Southern European countries; and on increase or decrease in some Continental European countries and Japan. These results suggest that the Kuznets proposal that inequality follows an inverted U-shape does not fit all countries. For those countries without the upturn around 1980, inequality presents an L-shape instead.

The literature on inequality has proposed several theories aimed at explaining the trends and structural breaks present in inequality series. Inequality developments have been explained in terms of technological breakthroughs, trends in globalization, and the link between inequality and economic growth.

Skill-biased technological change has long been suggested as one of the main factors shaping inequality over time (Murphy, 1989, Krueger, 2012). According to the proponents of this theory, in the absence of a growing supply of skilled workers, technological change will increase the wage difference between skilled and unskilled workers. Atkinson (2008) suggests that if countries are affected by the same technological change, the impact on wages will depend on the ability of each country to supply workers with higher skills. Therefore, according to Atkinson (2008), skill-biased technological change does not automatically lead to wage differences and higher inequality. Also, Caselli (1990) points out that not all technological changes are in fact skill biased. Furthermore, some technological changes may have boosted the productivity of low-skilled workers (Mokyr, 1990).

Regarding the role of globalization in explaining inequality, the findings in the literature are polarized. Whereas some authors conclude that globalization accentuates inequality (Firebaugh, 2003; Wade, 2004), others suggest that economic integration has played an important role in closing the inequality gap (Dollar and Kraay, 2002). The theoretical foundations of the causal link between globalization and inequality are grounded in trade theory. Whereas classical trade theory predicts an increase in inequality in countries with relatively abundant supplies of skilled labour and capital, modern trade theory is less clear-cut. Melitz (2003) and Melitz and Ottaviano (2008) suggest increasing returns in the top, while Leamer (2007) and Venables (2008) conclude that both the top and the bottom of the income distribution will benefit, to the detriment of the middle-income individuals. Globalization, along with information technology, may also play an important role in explaining the increasing wage dispersion observed for "stars" in certain professions (Rosen, 1981).

The link between inequality and growth has long been studied in both the theoretical and the empirical literature, with controversial results. On the one hand, several authors suggest that inequality may be good for growth if high inequality provides incentives to work harder and invest in order to take advantage of high rates of returns (Mirrlees, 1971, Lazear and Rosen, 1981) or if higher inequality fosters aggregate savings and capital accumulation (Kaldor, 1955, Bourguignon, 1981). On the other hand, greater inequality may limit growth if higher taxation and regulation implemented to tackle inequality in turn reduces the incentive to invest (Bertola, 1993, Alesina and Rodrick, 1994, Perotti, 1996); or if inequality implies under-investment by the poor in the presence of financial market imperfections (Galor and Zeira, 1993); or in the presence of skilled-biased technological change, as explained above.

While there has been a continuously evolving discussion of the time-varying nature of inequality for various developed countries, the econometric analysis is limited. This may be due to the fact that the income distribution data is relatively new (Atkinson and Leigh 2013), as we have already mentioned. One of the few econometric applications on time series data pertaining to inequality is that of Roine and Waldenstrom (2011), RW hereafter, where they apply multiple structural change tests within a single equation framework as proposed by Bai and Perron (1998, 2003), and a system of equations framework following the recent methodology developed by Qu and Perron (2007). The empirical analysis of RW attempts to test for and identify common breaks on the data of top income shares of eighteen countries using two separate time series data sets; one that covers a sample spanning almost a century and another that focusses on the post war period. Consequently, RW synthesise the conclusions drawn from the estimated structural breaks. While the study by RW is highly insightful, a major drawback is that they study assumes the inequality data to be stationary. In other words, shocks to the inequality data (top income shares) are likely to be transitory in nature. This however, needs to be tested empirically rather than making an assumption and as we will find in this paper, we cannot reject the possibility that shocks to top income shares are highly persistent in nature.

A recent study by Islam and Madsen (2015) tests whether income inequality is persistent by employing a long panel data set of 21 OECD countries over a time period spanning 1870 to 2011. The employ the Bai and Carrion-i-Silvestre (2009) panel unit root tests that allows for multiple structural breaks. They find that the shocks to income inequality are likely to be temporary. A possible drawback from this approach is that the time dimension is long and the cross sectional dimension is short. The time dimension could have been individually exploited given the sufficient data points that are available. Besides, the choice of allowing up to five breaks can raise problems. The incorporation of five structural breaks is not an appropriate strategy if one wants to determine if a unit root is present. This is because the unit root process can be viewed as a limiting case of a stationary process with multiple breaks, one that has a break (permanent shock) every period.

3. Hypothesis Testing Framework and econometric methodology

As explained above, Atkinson et al. (2011) argue that there has been a sharp drop in top income shares in the first half of the 20th century, around World War II and the Great Depression, whereas in the second half of the 20th century, there has been an increase in top income shares. These arguments suggest a set of hypotheses to be tested:

Hypothesis I: Whether we can detect structural breaks at the points that allow us to demarcate the three regimes: prior to Great Depression or World War II, following from this point of time up to the 1970s; and then the period thereafter. Since World War II the high rates of marginal taxation for the top income earners can be a cause for a structural break.

Hypothesis II: Whether the trend of top income shares can be found to be increasing or stagnant prior to the Great Depression, then decrease between World War II and the mid-1970s, and since then increase again (Piketty and Saez 2003). These regime may coincide with the start of assembly lines (early part of the twentieth century) or the ICT revolution of the 1970s and 1980s (Roine and Waldenstrom 2015).

Hypothesis III: Allowing for these structural changes if they exist, do we find evidence of persistent inequality? If shocks to inequality are not transitory, then technological innovations or financial shocks are likely to have persistent effects; which have consequences for policy such as advocating redistributive measures (Christopolous and McAdam 2017).

When estimating a trend in times series data, past studies have had to deal with the possibility of whether the data contained a unit root. This issue has been raised by Perron (1988) who concluded that the correct specification of the trend function would be affected due to the presence of a unit root. If for example, the time series data contains a unit root, then using ordinary least squares to test for the presence of a trend will suffer from severe size distortions. Conversely, if the time series data does not contain a unit root, or in other words is a trend stationary process, but is modeled as a unit root process, the tests will be inefficient and will lack power relative to the trend stationary process (see Perron and Yabu 2009a). Further, if one allows for the possibility of structural breaks in the time series data, the issue of determining the presence of a unit root in the data becomes complicated. For example, one can falsely conclude a data series to be a unit root process by neglecting a structural break in what is an otherwise trend stationary process (Perron 1989). Alternatively, in a difference stationary process, neglecting a trend break can one to incorrectly suggest the presence of stationarity (Leybourne, Mills, and Newbold 1998). Accordingly, recent studies have allowed for the presence of structural breaks when testing for the presence of unit roots. However, the estimates of the break dates that are obtained by minimizing these unit root tests are, in general, not consistent for the true break dates (Vogelsang and Perron 1998). Besides, these unit root tests suffer from the problem that they provide little information regarding the presence and number of trend breaks. Conversely, testing whether a time series process can be characterized by a broken trend is complicated by the fact that the nature of persistence in the errors is usually unknown. Indeed, inference based on a structural change test on the level of the data depends on whether a unit root is present while tests based on differenced data can have very poor properties when the series contains a stationary component (Vogelsang 1998). This circular testing problem underscores the need to employ break testing procedures that do not require knowledge of the form of serial correlation in the data.

Based on the above arguments, we choose to estimate the trend function based on the general model given by:

$$y_t = \mu_0 + \beta_0 t + \sum_{i=1}^{K} \mu_i DU_{it} + \sum_{i=1}^{K} \beta_i DT_{it} + u_t, \qquad t = 1,2,\dots,T \tag{1}$$
$$u_t = \rho u_{t-1} + \varepsilon_t, \qquad t = 2,3,\dots,T, \qquad u_1 = \varepsilon_1$$

where y_t denotes the data on top income shares, $DU_{it} = I(t > T_i)$, $DT_{it} = (t - T_i)I(t > T_i)$, $i = 1,2,\dots,K$. A break in the trend occurs at time, $T_i = [T\lambda_i]$, where $\beta_i \neq 0$, and λ_i is the break fraction. The date(s) for any break(s) in the series and the number of breaks (K) is unknown. No assumptions are made with regards to the nature of the error term, i.e. u_t can be either $I(0)$, that is, $|\rho| < 1$, or $I(1)$ that is, $\rho = 1$. To determine whether structural breaks exist we test the null hypothesis $H_0: \beta_i = 0$ against the alternative $H_1: \beta_i \neq 0$. Perron and Yabu (2009a) propose a novel method to detect a break in the trend function based on a Feasible Quasi Generalized Least Squares (FGLS) method and a further second break using a sequential approach due to Kejriwal and Perron (2010).

The motivation for adopting this approach is as follows: First, simulation evidence presented in Vogelsang and Perron (1998) and Lee and Strazicich (2001) suggests that the estimates of the break dates obtained by minimizing/maximizing the commonly used unit root tests (such as the Lee and Strazicich (2003) tests) over all possible break dates are unlikely to provide consistent estimates of the true break dates. Secondly, the unit root tests (such as Lumsdaine and Papell (1997), Zivot and Andrews (1992)) typically employed suffer from serious power and size distortions due to the asymmetric treatment of breaks under the null and alternative hypotheses. If breaks are indeed present, this information is not exploited to improve the power of the testing procedure. More importantly, these tests are subject to a spurious rejection problem when breaks are present under the unit root null hypothesis. Finally, many of the commonly used unit root tests provide little information regarding the existence or number of trend breaks. At an intuitive level, it seems more natural to be first able to ascertain if breaks are at all present before proceeding to conduct unit root tests allowing for such breaks. In the absence of breaks, these tests suffer from low power due to the inclusion of extraneous break dummies thereby potentially leading the researcher to estimate a differenced specification when a level specification is in fact more appropriate. Indeed, as stressed by Campbell and Perron (1991), proper specification of the deterministic components is essential to obtaining unit root tests with reliable finite sample properties.

The first step tests for one structural break in the slope of the trend function using procedures that are robust to the stationarity/non-stationarity properties of the data. The tests employed are designed to detect a break in slope while allowing the intercept to shift. A rejection by these robust tests can therefore be interpreted as evidence of a structural break in trend. Given evidence in favor of a break, we then proceed to test for one against two slope breaks using the extension of Perron and Yabu (2009) proposed by Kejriwal and Perron (2010). Again, this latter test allows us to distinguish between one and two breaks while being agnostic to whether a unit root is present. Given the number of sample observations available to be approximately 85, we allow for a maximum of two breaks in our empirical analysis. There are two reasons for this. As we have explained earlier, we expect according to the observations made by Piketty and Saez (2003) that there should be two breaks to account for the U-shape trend in top income shares data. Secondly, from an econometric viewpoint, allowing for a large number of breaks is not an appropriate strategy if one wants to determine if a unit root is present. The reason is that a unit root process can be viewed as a limiting case of a stationary process with multiple breaks, one that has a break (permanent shock) every period. Further, as discussed in Kejriwal and Perron (2010), the maximum number of breaks should be decided with regard to the available sample size. Otherwise, sequential procedures for detecting trend breaks will be based on successively smaller data subsamples (as more breaks are allowed) thereby leading to low power and/or size distortions. It is therefore important to allow for a sufficient number of observations in each segment and choose the maximum number of permissible breaks accordingly.

To briefly describe the Perron and Yabu (2009a) procedure which is to detect a break in the trend function based on a Feasible Quasi Generalized Least Squares (FGLS) method; first, the following auto regression on the error term in (1) is estimated:

$$\hat{u}_t = \alpha\hat{u}_{t-1} + \sum_{i=1}^{k}\varphi_i\hat{u}_{t-i} + e_{tk} \tag{2}$$

where the lag length k is chosen using the Bayesian Information Criteria (BIC). The estimate of α is obtained using OLS, denoted $\tilde{\alpha}$. Perron and Yabu (2009) use a bias corrected version of $\tilde{\alpha}$, denoted by $\tilde{\alpha}_M$, to improve the finite sample properties of the tests, proposed by Roy and Fuller (2001). In the next step, Perron and Yabu (2009a) calculate the super-efficient estimator of α given by:

$$\tilde{\alpha}_{MS} = \begin{cases} \tilde{\alpha}_M \text{ if } |\tilde{\alpha}_M - 1| > T^{-1/2} \\ 1 \quad \text{if } |\tilde{\alpha}_M - 1| \leq T^{-1/2} \end{cases} \tag{3}$$

Using a super-efficient estimate is crucial for obtaining nearly identical limit properties in the I (0) and I(1) cases. The estimate $\tilde{\alpha}_{MS}$ is then used to construct the quasi differenced regression

$$(1-\tilde{\alpha}_{MS})y_t = (1-\tilde{\alpha}_{MS})x'_{L1,t}\Psi + (1-\tilde{\alpha}_{MS})u_t;\ t = 2,3,\dots,T$$
$$y_t = x'_{L1,1}\Psi + u_1 \tag{4}$$

where $\Psi = (\mu_0, \beta_0, \mu_1, \beta_1)'$. The resulting estimates from the regression are denoted as $\tilde{\Psi}^{FG} = \left(\tilde{\mu}_0^{FG}, \tilde{\beta}_0^{FG}, \tilde{\mu}_1^{FG}, \tilde{\beta}_1^{FG}\right)'$. The Wald test $W_{QF}(\lambda)$ for a particular break function λ_1, where the subscript QF denotes the Quasi Feasible GLS is given by

$$W_{QF}(\lambda_1) = \left(\tilde{\beta}_1^{FG}(\lambda_1)\right)^2 / \sqrt{\left[\left(\tilde{h}_v(\lambda_1)\right)\{(X^{\alpha\prime}X^{\alpha})^{-1}\}\right]} \tag{5}$$

where $X^{\alpha} = \left[x_{L1,1}, (1-\tilde{\alpha}_{MS})x_{L1,2}, \dots, (1-\tilde{\alpha}_{MS})x_{L1,T}\right]'$. The quantity $\tilde{h}_v(\lambda_1)$ is an estimate of 2π times the spectral density function of $v_t = (1-\alpha L)u_t$ at frequency zero. If $|\tilde{\alpha}_{MS}| < 1$, a kernel-based estimator given by

$$\tilde{h}(\lambda_1) = T^{-1}\sum_{t=1}^{T}\hat{v}_t^2(\lambda_1) + 2T^{-1}\sum_{j=1}^{T-1}k(j,\tilde{l})\sum_{t=j+1}^{T}\hat{v}_t(\lambda_1)\hat{v}_{t-j}(\lambda_1) \tag{6}$$

is employed where $\hat{v}_t(\lambda_1)$ are the least squares residuals from (3). The function $k(j,\tilde{l})$ is the quadratic spectral kernel and $\tilde{l}$ is the bandwidth. When $\tilde{\alpha}_{MS} = 1$, the estimate suggested is an autoregressive spectral density estimate that can be obtained from the regression:

$$\hat{v}_t = \sum_{i=1}^{k}\xi_i\hat{v}_{t-i} + e_{tk} \tag{7}$$

where the lag length k is again chosen using the BIC. Following Andrews (1993) and Andrews and Ploberger (1994), Perron and Yabu (2009) consider the Mean, Exp, and sup functionals of the Wald test for different break dates. They found that with the Exp functional, the limit distribution in the I(0) and I(1) cases are nearly identical. They recommend the following statistic to determine the structural break:

$$ExpW = ln\left[T^{-1}\sum_{\lambda_1\in\Lambda_1}exp\left(1/2\,W_{QF}(\lambda_1)\right)\right] \tag{8}$$

In the spirit of Perron and Yabu (2009), Kejriwal and Perron (2010) propose a sequential procedure that allows one to obtain a consistent estimate of the true number of breaks irrespective of whether the errors are I(1) or I(0). The first step is to conduct a test for no break versus one break. Conditional on a rejection, the estimated break date is obtained by a global minimization of the sum of squared residuals. The strategy proceeds by testing each of the two segments (obtained using the estimated partition) for the presence of an additional break and assessing whether the maximum of the tests is significant. Formally, the test of one versus two breaks is expressed as:

$$ExpW(2|1) = \max_{1\le i\le 2}\left\{ExpW^{(i)}\right\} \tag{9}$$

where $ExpW^{(i)}$ is the one break test in segment i. We conclude in favour of a model with two breaks if $ExpW(2|1)$ is sufficiently large.

In the second stage of the empirical analysis we conduct robust estimations of the trend. If no structural breaks are found to be present in the data, then we estimate the trend function for the entire sample. However, if breaks are found to be present in the data, we delineate the sub-samples from the break points and conduct robust trend estimation for each of the regimes demarcated by the breaks points. To this end we apply an appropriate econometric method of robust trend estimation due to Perron and Yabu (2009b) that allows one to be agnostic to the nature of persistence of errors in the trend function.

The procedure is quite similar in spirit to the Perron and Yabu (2009b) procedure so we omit the details and outline the main differences. The reader is referred to the Perron and Yabu (2009a) paper for further details. First, the residuals $\hat{u}_t$ in (2) are now obtained from a regression of y_t on $x_t = (1, t)'$. Next, the super-efficient estimate $\tilde{\alpha}_{MS}$ (obtained as discussed earlier) is used to estimate the quasi-differenced regression

$$(1 - \tilde{\alpha}_{MS}L)y_t = (1 - \tilde{\alpha}_{MS}L)x'_t\Psi^0 + (1 - \tilde{\alpha}_{MS}L)u_t; t = 2,3, \dots, T$$
$$y_t = x'_1\Psi + u_1 \quad (10)$$

where $\Psi^0 = (\mu_0, \beta_0)'$. Denote the estimate of β_0 from this regression by $\hat{\beta}_0$. Then, using the notation $x^{FG} = (x_1^{FG}, x_2^{FG}, \dots, x_T^{FG})'$ with $x_1^{FG} = (1,1)'$; $x_t^{FG} = [1 - \tilde{\alpha}_{MS}, t - \tilde{\alpha}_{MS}(t-1)]$ for $t = 2,3, \dots, T$; a $100(1-\alpha)\%$ confidence interval for β_0; again valid for both I(1) and I(0) errors, is obtained as

$$\hat{\beta}_0 \pm c_{\alpha/2}\sqrt{(\tilde{h}_v)\{(X^{\alpha\prime}X^{\alpha})^{-1}\}} \quad (11)$$

where $c_{\alpha/2}$ is such that $P(x > c_{\alpha/2}) = \alpha/2$ for $x \sim N(0,1)$ and $\tilde{h}_v$ is already defined.

In the final stage of empirical analysis, we conduct unit root tests to ascertain the nature of persistence in the top income shares data. If there is evidence of structural breaks, we apply a new class of unit root tests which allows for breaks under both the null and alternative hypotheses (Carrion-i-Silvestre, Kim, and Perron 2009).

The tests are extensions of the feasible point optimal statistic of Elliott et al. (1996) and the *M* class of tests due to Ng and Perron (2001).

Consider equation (1); the estimates of the break fractions λ_i and the regression parameters are obtained by minimizing the sum of squared residuals from the quasi-differenced regression analogous to (4). The sum of squared residuals evaluated at these estimates is denoted by $S(\alpha(\hat{\lambda}), \hat{\lambda})$, where $\alpha(\hat{\lambda}) = 1 - c(\hat{\lambda})/T$. The feasible point optimal statistic is then given by:

$$PT - GLS = S(\alpha(\hat{\lambda}), \hat{\lambda}) - \alpha(\hat{\lambda})S(1, \hat{\lambda})/s^2(\hat{\lambda}) \quad (12)$$

Where $s^2(\hat{\lambda}) = s_{ek}^2/[1 - b(1)]^2$ and $s_{ek}^2 = (T-k)^{-1}\sum_{t=k+1}^{T}\hat{e}_{tk}^2$; $b(1) = \sum_{j=1}^{k}\hat{b}_j$ Both $\hat{b}_j$ and $\hat{e}_{tk}^2$ are obtained using OLS estimation of the following equation:

$$\Delta \tilde{y}_t = b_0 \tilde{y}_{t-1} + \sum_{j=1}^{k} b_j \Delta \tilde{y}_{t-j} + e_{tk}$$

where $\tilde{y}_t = y_t - \widehat{\Psi}_2{}' x_{Li,t}(\hat{\lambda})$; $x_{Li,t}(\hat{\lambda}) = [1, t, DU_{it}(\hat{\lambda}), DT_{it}(\hat{\lambda})]$; i denotes the number of breaks; and $\widehat{\Psi}_2{}'$ is the OLS estimate of the quasi differenced regression (4).

Carrion-i-Silvestre et al. (2009) also consider extensions of the M-class of tests analysed in Ng and Perron (2001). These extensions involve the inclusion of multiple structural breaks, building on the work of Perron and Rodriguez (2003). The statistics computed by Carrion-i-Silvestre et al. (2009) are similar to Ng and Perron (2001) where the null hypothesis is that of a unit root against the alternative of stationarity with the symmetric treatment of structural breaks in the null and alternative hypothesis. These statistics are computed as follows:

$$MPT = \left[c^2(\hat{\lambda}) T^{-2} \sum_{t=2}^{T} \tilde{y}_{t-1}^2 + \left(1 - c(\hat{\lambda})\right) T^{-1} \tilde{y}_T^2\right] / s^2(\hat{\lambda}) \tag{13}$$

$$MZa = [T^{-1} \tilde{y}_T^2 - s^2(\lambda)](2T^{-1} \sum_{t=2}^{T} \tilde{y}_{t-1}^2)^{-1} \tag{14}$$

$$MSB = (T^{-2} \sum_{t=2}^{T} \tilde{y}_{t-1}^2)^{1/2} / s^2(\hat{\lambda}) \tag{15}$$

$$MZt = [T^{-1} \tilde{y}_T^2 - s^2(\lambda)]\left(4s^2(\hat{\lambda}) T^{-2} \sum_{t=2}^{T} \tilde{y}_{t-1}^2\right)^{-1/2} \tag{16}$$

where $s^2(\hat{\lambda})$, $\tilde{y}_t$ and $c(\hat{\lambda})$ have already been defined. The computation of the critical values of these powerful unit root tests are described by Carrion-i-Silvestre et al. (2009).

Such a symmetric treatment of breaks alleviates these unit root tests from size and power problems that plague tests based on search procedures (for instance, Zivot and Andrews 1992, Lumsdaine and Papell 1997). If no evidence is found of structural breaks, we apply standard (no break) unit root tests developed by Elliott, Rothenberg, and Stock (1996) and Ng and Perron (2001). There is always a potential power issue associated with unit root tests allowing for multiple breaks, given that a unit root process is observationally equivalent to a stationary process with multiple breaks in the limit. Simulation evidence presented in Carrion-i-Silvestre, Kim, and Perron (2009) shows that the tests allowing up to two breaks have decent finite sample power when the data generating process is driven by one or two breaks. Indeed, they have much better properties than unit root tests based on search procedures given that they exploit information regarding the presence of breaks.

4. Data and Empirical Results

The data spans the period of 1921 to 2000. The two exceptions are India, which ends in 1999, and the Netherlands which begins in 1915 and ends in 1999.[1] Figure 1 below shows the trending behaviour of top income shares of the selected countries in this study. We can note by eyeballing the data, that the underlying trends do not seem similar when compared separately, and a case for one or more structural breaks does

[1] Special thanks to Daniel Waldenström for making the data available on his website.
The data is available at http://www.uueconomics.se/danielw/Data.htm. For source and description of data please see Roine and Waldenström (2011).

seem plausible for selected countries.

[Figure 1 about here]

The main source for the construction of top income shares data is by using the personal income tax returns on the national level. Income shares are calculated following a methodology first outlined in Piketty (2001, 2003) which in turn builds on the work by Kuznets (1953). Top income shares are constructed by dividing the number of top share tax units and their incomes, with the reference tax population and their total income. The income is gross total income before taxes and transfers (see Roine and Waldenström 2011 for details).

Hypothesis I : Structural Breaks

We test for the presence of structural breaks using the procedure by Perron and Yabu (2009) and Kejriwal and Perron (2010) allowing for up to 2 breaks, where the null hypothesis is that a series does not contain a break against the alternative that there are breaks. Table 1 reports the test results and, where present, the likely date of the break.

Table 1: Structural Break Test Results

	ExpW 0|1	ExpW 1|2	# of breaks	TB 1	TB 2
Australia	0.74		0		
Canada	3.11^{**}	31.57	2	1932	1979
New Zealand	0.03		0		
USA	3.67^{***}	0.64	1	1973	
France	0.93		0		
Sweden	8.45^{**}	7.33	2	1971	1983
Norway	14.73^{***}		1	1988	
Japan	0.97		0		
Finland	3.21^{**}	34.75	2	1973	1986
Netherlands	0.05		0		
India	0.25		0		

From the empirical results we can see that the structural change points in the data do not conform with the views of Piketty and Saez (2003) except for Canada.

For the remaining 10 countries chosen, we find two structural breaks for two countries (Sweden and Finland) and the break dates are quite similar. Two other countries (USA and Norway) are found to contain a single break. For the countries where we find evidence of breaks, the preponderance of break locations are in the 1970s and 1980s. Six countries show no evidence of any structural change.

Hypothesis II: Trends

Next, for countries where a break is identified, we partition the sample into separate regimes and estimate the linear trends for each regime following the method due to Perron and Yabu (2009b) as described in the previous section. The trend estimates for pre-break and post-break regimes are reported in Table 2. For those countries that exhibit two breaks, we partition the data in to three regimes, whereas for a single break case, the number of regimes is two. However, for meaningful estimates to be obtained, a sufficient number of observations is necessary for estimation of a trend in each regime.

Given that some of the break points are found to be in the 1980s, the trend estimates for the post break regime in this case are not reported, simply because the estimates are not possible and in those cases we have highlighted that there are too few data points to obtain meaningful results. Where estimates are obtained, the associated confidence intervals are reported within parentheses. For those countries where no breaks are found, the trend estimates are based on the whole sample of data points.

Table 2 Trend Estimation Results

	Regime I	Regime II	Regime III
Australia	–0.0109 90% conf. int. (–0.0220, 0.0001) 95% conf. int. (–0.0241, 0.0022)	N/A	N/A
Canada	Too few observations	– 0.0176** 90% conf. int. (–0.0319, –0.0034) 95% conf. int. (–0.0346, –0.0007)	0.0260** 90% conf. int. (0.0102, 0.0418) 95% conf. int. (0.0073, 0.0448)
New Zealand	–0.0040 90% conf. int. (–0.0190, 0.0109) 95% conf. int. (–0.0218, 0.0137)	N/A	N/A
USA	–0.0133* 90% conf. int. (–0.0259, –0.0007) 95% conf. int. (–0.0283, 0.0016)	0.0272** 90% conf. int. (0.0091, 0.0453) 95% conf. int. (0.0057, 0.0487)	N/A
France	–0.0105 90% conf. int. (–0.0219, 0.0010) 95% conf. int. (–0.0241, 0.0031)	N/A	N/A
Sweden	–0.0169** 90% conf. int. (–0.0243, –0.0095) 95% conf. int. (–0.0256, –0.0081)	Too few observations	0.0254** 90% conf. int. (0.0216, 0.0291) 95% conf. int. (0.0209, 0.0298)
Norway	–0.0150** 90% conf. int. (–0.0212, –0.0088) 95% conf. int. (–0.0224, –0.0076)	Too few observations	
Japan	–0.0103 90% conf. int. (–0.0296 ,0.0091)	N/A	N/A

	95% conf. int. (–0.0333, 0.0128)		
Finland	–0.0108 90% conf. int. (–0.0274, 0.0057) 95% conf. int. (–0.0305, 0.0088)	Too few observations	Too few observations
Netherlands	–0.0186** 90% conf. int. (–0.0252, –0.0120) 95% conf. int. (–0.0264, –0.0108)	N/A	N/A
India	–0.0046 90% conf. int. (–0.0229, 0.0138) 95% conf. int. (–0.0263, 0.0172)	N/A	N/A

** and * denote significance at 5% and 10% levels respectively; the numbers in brackets are the confidence intervals. NA denotes not applicable, given there are no breaks.

First we consider the trend estimates of Canada which is the only country to contain significant structural breaks, where the break locations are in line with the views advocated by Piketty and Saez (2003). In regime I (1921 – 1932) the number of observations is too low to obtain meaningful estimates of a trend. However, in Regime II (1932 – 1979), we find a significant negative trend; followed by Regime III (1979 – 2000) where the trend is positive. There is some support in this case of the view that inequality started to increase since the 1970s. Based on the finding of two structural breaks for Sweden and Finland, we find meaningful trend estimates only in in a single regime. In the case of Sweden for example, in Regime I (1921 – 1971), the estimate is negative; Regime II (1971 – 1983) contains too few observations to obtain meaningful estimates. While in the case of Regime III (1983 – 2000) the number of observations is not quite enough, the estimates reported should be treated with caution. However, if we are to consider these estimates, it seems that the trend is increasing. In the case of Finland however, we only obtain estimates for Regime I (1921 – 1973) which are found to be insignificant. In the case of Norway (1921 – 1988) the trend is negative, but with few observations in the second regime, we cannot produce a meaningful trend estimate.

Overall, there is some if not overwhelming evidence in favour of the trends advocated by Piketty and Saez (2003). Using the piecewise linear method of fitting linear trends to regimes demarcated by structural breaks, there does seem to be some evidence that top income shares declined until the 1970s and thereafter the trend (albeit not significant in some cases) may have reversed.

Hypothesis III: Unit Roots

Following the results in Table 1, we employ the unit root tests proposed by Carrion-i-Silvestre *et al* (2009) which allow for breaks under both the null and alternative hypotheses. For Australia, France, Japan, Netherlands, India, where no structural break is found, the M-class tests proposed by Elliott *et al* (1996) and Ng and Perron (2001) are applied. The results of the tests are reported in Table 3 below.

Table 3: Unit Root Tests

	MZA	MZt	MSB	MPT	ADF	PT
	With Structural Breaks					
Canada	-9.18	-2.10	0.229	21.94	-2.25	21.99
USA	-10.44	-2.27	0.217	14.55	-2.56	17.41
Sweden	-7.28	-1.86	0.255	31.56	-1.99	36.19
Norway	-6.91	-1.85	0.268	16.35	-2.02	18.72
Finland	-7.53	-1.86	0.247	23.93	-1.98	24.29
	No Structural Breaks					
Australia	-2.23	-1.05	0.469	10.89	-1.68	13.99
N. Zealand	-13.02	-2.47	0.189	9.35	-2.67	9.18
France	-0.04	-0.03	0.710	31.12	-1.92	53.95
Japan	-1.13	-0.63	0.55	17.28	-1.89	25.97
Netherlands	-11.68	-2.36	0.202	8.10	-2.10	8.19
India	-2.74	-1.12	0.407	8.74	-1.33	9.53

No Structural Breaks: unit root statistics are computed using Ng and Perron (2001) and Elliot *et al* (1996). The number of lags is chosen by Modified Akaike Information Criterion (MAIC) as recommended by Ng and Perron (2001). *With Structural Breaks*: the unit root test statistics allowing for a break in both the null and the alternative using Carrion-i-Silvestre *et al* (2009).

The results of the unit root tests with no structural breaks show that we are unable to reject the null hypothesis of a unit root for the entire sample of countries. Following Carrion-i-Silvestre et. al. (2009), we employ the same battery of tests, the Generalized Least Squares Dickey Fuller ($DF - GLS$), the Point Optimal (P_T) tests and the M-class tests (MZ_a, MZ_t, MSB, MP_T), this time allowing for a single structural break in the trend. The results show that even after accounting for structural breaks in both the null and the alternative hypotheses, we find no evidence in which to reject the null hypothesis of a unit root. Overall, from our results, we are unable to reject the hypothesis of a unit root in top income share data. Thus, shocks to top income shares are likely to be highly persistent.

5. Conclusion

This paper adds to the literature on the long-run development in top income shares by testing three hypotheses. First, we test for structural breaks in the series using robust methods that are agnostic regarding the stationarity or nonstationarity of the series. Second, using the piecewise-linear method of fitting linear trends to regimes demarcated by structural breaks, we estimate the trends in the inequality series for the pre-break and/or inter-break, and post-break regimes. Finally, we test for the degree of persistence in the analysed series. Through testing these hypotheses, we obtain a comprehensive time series characterization of long-run inequality behaviour for a set of eleven countries.

With the exception of Canada and USA, our results on structural breaks do not entirely support the views of Piketty and Saez (2003) or Roine and Waldenström (2011). In the case of Australia, the trend is negative throughout the sample, whereas for New Zealand, we do not find any evidence of a significant trend. However, we find some evidence of a decreasing trend in top income shares up to the 1970s followed by the upturn around the 1980s. Top income shares appear to be highly persistent despite the presence (or not) of structural breaks in the data. Contrary to Piketty and Saez (2006), we find that the pattern of trends in continental European countries is mixed. For France, there is no significant trend for the entire sample, and while we find evidence of structural breaks for Finland, there is no evidence of any significant trend in the major sub-sample; the other sub-samples are too short (containing few observations) to make any meaningful estimates of the trend. For the Netherlands, there is no structural break and there is evidence of a negative unbroken trend for the entire sample. For Norway, there is a negative trend for most part of the sample, while for Sweden, the trend is similar to that of the Anglo-Saxon countries; first, a declining trend in top income share followed by a brief interval where no clear trend estimation is possible due to few observations, followed by an increasing trend. In the case of the two Asian countries, there is no evidence of any structural break and no significant trend. Our results show that there is no common trending behaviour when comparing groups of countries such as Nordic countries with Anglo-Saxon or continental Europe. This result departs from that of Roine and Waldenstrom (2011) in terms of the lack of common break dates and from Atkinson and Piketty (2007) with regards to the heterogeneity of trends within groups such as Anglo-Saxon countries or Nordic countries. It has been argued that technology shifts that are skills biased, can change the trend of inequality. We see some evidence of this, that there is a change in the trend for Anglo-Saxon countries such as Canada and USA, and a continental European country, being Sweden. The trends coincide with the views that the introduction of assembly lines may have caused a decrease in inequality while the ICT revolution led to an increase in inequality. However, this does not happen for countries such as Australia or France. This however, is not completely unexpected as technological changes do not take place at the same time around the world due to adoption lags (Comin and Mestieri 2013).

Finally, a test is carried out on how persistent shocks are to the top income shares. We find that using unit root tests that allow for structural breaks (where we do find evidence of breaks) and those that do not contain breaks (where the data does not show evidence of any breaks), the conclusion is clearly in favour of inequality being highly persistent to shocks. This view is contrary to that of Islam and Madsen (2015) but supports the conclusions of Christopoulos and McAdam (2017).

References

Alesina, A. and Rodrik, D., 1994. Distributive politics and economic growth, *Quarterly Journal of Economics*. Vol. CIX (2), pp. 465-490.

Alvaredo, F., Atkinson, A.B., Piketty, T. and Saez, E., 2013. The Top 1 Percent in International and Historical Perspective. *Journal of Economic Perspectives*, Vol. 27(3), pp. 3-20.

Atkinson, A.B., 2008. The Changing Distribution of Earnings in OECD Countries Oxford and New York. Oxford University Press, Oxford.

Atkinson, A.B. and Søgaard, J.E., 2016. The Long-Run History of Income Inequality in Atkinson, A.B., Piketty, T. (Eds.), 2007. Top Incomes over the Twentieth Century: A Contrast Between European and English-Speaking Countries. Oxford University Press, Oxford.

Atkinson, A.B., Søgaard, J.E., 2013. The long-run history of income inequality in Denmark. *Scandinavian Journal of Economics*, Vol. 118(2), pp. 264-291.

Atkinson, A.B., Piketty, T. (Eds.), 2010. Top Incomes: A Global Perspective. Oxford University Press, Oxford.

Atkinson, A.B., Piketty, T. and Saez, E., 2011. Top Incomes in the Long Run of History. *Journal of Economic Literature*. Vol. 49(1), pp. 3-71.

Bertola, G., 1993. Factor Shares and Savings in Endogenous Growth. *American Economic Review*, Vol. 83(5), pp. 1184-1198.

Bourguignon, F., 1981. Pareto Superiority of Unegalitarian Equilibria in Stiglitz' Model of Wealth Distribution with Convex Saving Function. *Econometrica*, Vol. 49(6), pp. 1469-75.

Burkhauser, R.B., Hahn, M.H. and Wilkins, R.J., 2015. Measuring Top Incomes Using Tax Record Data: A Cautionary Tale from Australia. *Journal of Economic Inequality*, Vol. 13(2), pp. 181-205.

Carrion-i-Silvestre, J.L., Kim, D. and Perron, P., 2009. GLS-Based Unit Root Tests with Multiple Structural Breaks Under Both the Null and the Alternative Hypotheses. *Econometric Theory*, Vol. 25(6), pp. 1754-1792.

Caselli, F., 1999. Technological revolutions. *American Economic Review*, Vol. 89(1), pp. 78–102.

Christopoulos, D., & McAdam, P. 2017. On the Persistence of Cross-Country Inequality Measures. *Journal of Money, Credit and Banking*, *49*(1), 255-266.

Comin, D. A., and Mestieri, M. 2013. *Technology diffusion: Measurement, causes and consequences* (No. w19052). National Bureau of Economic Research.

DeBacker, J., Heim, B., Panousi, V., Ramnath, S., and Vidangos, I., 2013. Rising inequality: transitory or persistent? New evidence from a panel of US tax returns. *Brookings Papers on Economic Activity*, Vol. 2013(1), pp. 67-142.

Dollar, D. and Kraay, A., 2002. Spreading the wealth. *Foreign Affairs*, Vol. 81, pp. 120–133.

Elliott, G., Rothenberg, T. and Stock, J., 1996. Efficient Tests for an Autoregressive Unit Root. *Econometrica*, Vol. 64, pp. 813 – 836.

Firebaugh, G., 2003. *The New Geography of Global Income Inequality*. Cambridge, MA: Harvard University Press.

Fuest, C. and Huber, B., 2001. Labor and Capital Income Taxation, Fiscal Competition, and the Distribution of Wealth. *Journal of Public Economics*, Vol. 79(1), pp. 71-91.

Galor, O. and Zeira, J., 1993. Income distribution and macroeconomics. *Review of Economic Studies, Vol.* 60(1), pp. 35-52.

Hamilton, K. and Hepburn, C. 2014. Wealth. *Oxford Review of Economic Policy*, Vol. 30(1). pp. 1-20.
Ihori, T., 2001. Wealth Taxation and Economic Growth. *Journal of Public Economics*, Vol. 79(1), pp. 71-91.
Islam, M.R. ans Madsen, J.B., 2015. Is Income Inequality Persistent? Evidence using Panel Stationarity Tests, 1870 – 2011. *Economics Letters*, Vol. 127, pp. 17-19.
Kaldor, N., 1955. Alternative Theories of Distribution. *The Review of Economic Studies,* Vol. 23(2), pp. 83-100.
Krueger, A., 2012. The Rise and Consequences of Inequality", *Presentation made to the Center for American Progress in Washington, DC*, January
Kuznets, S., 1955. Economic growth and income inequality. *American Economic Review*, Vol.45(1), pp. 1-28.
Kuznets, S., & Jenks, E. 1953. Shares of Upper Income Groups in Savings. In *Shares of Upper Income Groups in Income and Savings* (pp. 171-218). NBER.
Lazear, E.P. and Rosen, S., 1981. Rank-Order Tournaments as Optimum Labor Contracts. *Journal of Political Economy*, Vol. 89(5), pp. 841-864.
Leamer, E.E., 2007. A flat world, a level playing field, a small world after all, or none of the above? A review of Thomas L Friedman's the World is flat. *Journal of Economic Literature*, Vol.45(1), pp. 83-126.
Lee, J. and Strazicich, M. 2003. Minimum Lagrange Multiplier Unit Root Test with Two Structural Breaks. *Review of Economics and Statistics*, Vol. 85, pp. 1082-1089.
Leigh, A., 2007. How Closely do Top Income Shares Track Other Measures of Inequality?. *Economic Journal*, Vol. 117(524), pp. F589-F603.
Leybourne, S., Mills, T. and Newbold, P., 1998. Spurious rejections by Dickey-Fuller tests in the presence of a break under the null. *Journal of Econometrics*, Vol. 87, pp. 191-203.
Melitz, M.J., 2003. The impact of trade on intra-industry reallocations and aggregate industry productivity. *Econometrica*, Vol. 71(6), pp. 1695-1725.
Melitz, M.J., Ottaviano, G., 2008. Market size, trade, and productivity. *Review of Economic Studies*. Vol. 75, pp. 295–316.
Mirrlees, J., 1971. An Exploration in the Theory of Optimum Income Taxation. *Review of Economic Studies*, Vol. 38(114), pp. 175-208.
Mokyr, J., 1990. The Lever of Riches: Technological Creativity and Economic Progress. Oxford University Press, Oxford.
Murphy, K.J., 1999. Executive compensation. In: Ashenfelter, O., Card, D. (Eds.), Handbook of Labor Economics, Vol. 3. North-Holland, Amsterdam.
Nelson, C.R. and Plosser, C.I., 1982. Trends and Random Walks in Macroeconomic Time Series. *Journal of Monetary Economics*, Vol. 10(2), pp. 139-162.
Ng, S. and Perron, P., 2001. Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power. *Econometrica*, Vol. 69, pp. 1519-1554.
Perron, P., 1989. The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis. *Econometrica*, Vol. 57, pp. 1361-1401.
Perotti, R., 1996. Growth, Income Distribution and Democracy: What the Data Say. *Journal of Economic Growth*, Vol. 1(2), pp. 149-187.
Perron, P. and Yabu, T., 2009a. Testing for Shifts in Trend with an Integrated or Stationary Noise Component. *Journal of Business and Economic Statistics*, Vol. 27, pp. 369-396.
Perron, P. and Yabu, T. 2009b. Estimating Deterministic Trends with an Integrated or Stationary noise component. *Journal of Econometrics*, Vol. 151, pp. 56-69.
Piketty, T., 2001. Income inequality in France, 1900–1998. CEPR Discussion Paper

No. 2876, Centre for Economic Policy Research, London.
Piketty, T., 2003. Income inequality in France, 1901–1998. *Journal of Political Economy*. Vol.111(5), pp. 1004-1042.
Piketty, T. 2005. Top Income Shares in the Long Run: An Overview. *Journal of the European Economic Association*, Vol. 3(2-3), pp.1-11.
Piketty, T., 2007. Income inequality in France, 1900–1998. In: Atkinson, A.B., Piketty, T. (Eds.), Top Incomes over the Twentieth Century: A Contrast Between European and English-Speaking Countries. Oxford University Press, Oxford.
Piketty, T. and Saez, E., 2014. Inequality in the Long Run," *Science*, Vol. 344 (6186), pp. 838-843.
Roine, J., Vlachos, J., & Waldenström, D. 2009. The long-run determinants of inequality: What can we learn from top income data? *Journal of Public Economics*, *93*(7), 974-988.
Roine, J. and Waldenström, D. 2011. Common Trends and Shocks to Top Incomes: A Structural Breaks Approach. *Review of Economics and Statistics*, Vol. 93, pp.832-846.
Roine, J. and Waldenström, D. 2015. Long-Run Trends in the Distribution of Income and Wealth. In Atkinson, A.B. & Bourguignon, F. (eds.) *Handbook of Income Distribution*, 2(A), pp. 469-592. Elsevier B.V., USA.
Venables, A.J., 2008. Rethinking economic growth in a globalizing world: an economic geography lens. Working Paper No. 18, The International Bank for Reconstruction and Development/World Bank, on behalf of the Commission on Growth and Development.
Wade, R.H., 2004. Is Globalization Reducing Poverty and Inequality?. *World Development*. Vol. 32(4), pp 567-589.

Figure 1. Top income shares.