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ON THE DYNAMICS OF INCOME INEQUALITY

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Abstract

We analyse the dynamics of income inequality using top income share data of selected countries. We contribute to the recent studies that 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. We examine the trends and conclude there is no clear evidence that Anglo Saxon countries have similar trends with Nordic countries as has been suggested in recent studies. Finally, shocks to the top income share data is not transitory, which have consequences for policy such as advocating redistributive measures.

JEL Codes: C22, C32, N30

Keywords: Inequality, Top Income Shares, Trends, Persistence

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, efforts have been made to provide data on inequality. While the primary focus has been on building micro-panel data sets based on national household surveys, the consequent lack of data spanning long time periods 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 data sets spanning long time periods 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 2005), 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 (2014).

There have been calls for exploiting the dynamics of long run inequality data over time paying attention to the variation of countries, using 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 the 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, 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, USA) have experienced a substantially greater increase than non-English speaking countries (such as France, Sweden, Norway, Finland, 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 which include Nordic, Anglo Saxon and Asian, exhibit a secular decline in top income shares over the twentieth century. Recent studies conclude that 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 no increase or decrease in some Continental European countries and Japan. These results suggest that Kuznet's proposal that inequality follows an inverted U-shape does not apply to all countries.

The literature on inequality has proposed several theories aimed at explaining the trends and structural breaks present in inequality data. Inequality has 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, 1999, 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 (1999) 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. While 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. Classical trade theory predicts an increase in inequality in countries with relatively abundant supplies of skilled labour and capital, on the contrary, 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 (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). One of the few econometric applications on time series data pertaining to inequality is that of Roine and Waldenstrom (2011), 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 Roine and Waldenstrom (2011) attempts to test for and identify common breaks in 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. While their study is highly insightful, a major drawback is that their study assumes the inequality data to be stationary. This property needs to be empirically tested. As we will find in this paper, we reject that shocks to top income shares are transitory in nature.

A recent study by Islam and Madsen (2015) tests whether income inequality is persistent by employing a long panel data set of Gini coefficients and top 10% income shares for 21 OECD countries over the period 1870–2011. They employ the individual and panel stationary tests due to Carrion-i-Silvestre (2005) allowing for a maximum of five structural breaks. The test is based on the Kwiatkowski et. al. (1992) (KPSS) test. They compute the bootstrap distribution following Maddala and Wu (1999) with 10,000 replications to take account of cross-sectional dependence in the estimates of the KPSS test statistics in order to reduce the bias and increase the power of the tests. As a robustness test, they employ the Bai and Carrion-i-Silvestre (2009) panel unit root tests that allows for multiple structural breaks. They conclude that the shocks to income inequality are temporary. The methods applied are comprehensive and show that there are mechanisms that bring income shares to a constant level. However, in another more recent and comparable study, Christopoulos and McAdam (2017) examine inequality persistence in a multi-country unbalanced panel using a range of stationary and long memory tests. They analyse the Gini index for 47 countries spanning a time period of at least 30 years. The tests employed include panel unit roots with and without breaks. The test for unit roots with breaks is based on a novel procedure that allows for a Fourier function. Finally a panel fractional unit root test is also conducted. Conducting these battery of tests, they find no evidence of shocks being transitory to inequality measures. The results of Christopolous and McAdam (2017) contradict those of Islam and Madsen (2015).

In this study we try to address the mixed results on the persistence of income inequality by adopting a time series approach. Besides, the measurement of trends in top income shares data, we address the issue of persistence allowing for the possible presence of structural breaks. We motivate our use of methods by taking into account the following considerations: First, the unit root tests provide little information regarding the existence or number of trend breaks. Intuitively, it would be reasonable to first determine if structural breaks exist in the data before proceeding to conduct unit root tests allowing for such breaks. The reason is that such tests suffer from low power due to the inclusion of extraneous break dummies. This leads to the possible estimation of a differenced specification when a level specification is in fact more appropriate. Campbell and Perron (1991) argue that the proper specification of the deterministic components is essential to obtaining unit root tests with reliable finite sample properties (see Ghoshray et. al. 2014). Secondly, the unit root tests typically employed suffer from serious power and size distortions when structural breaks are included only under the null or only under the alternative hypotheses. If structural breaks are present in the data, this information is not exploited to improve the power of the testing procedure. Further, these tests are subject to a spurious rejection problem when breaks are present under the unit root null hypothesis (see Ghoshray et. al. 2014). When testing for structural breaks in top income shares as well as persistence in the data, we take in to account these issues.

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 two or three regimes: prior to Great Depression or World War II, following from this point of time up to the 1980s; 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. For example, it has been argued that between 1950 and 1980 most countries went through a relatively Egalitarian phase, when low inequality prevailed.

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 regimes 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 exogenous shocks, such as 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). Alternatively, if shocks to inequality are transitory then it implies that opportunities exist for distributional mobility that allow income shares to be brought towards a constant level in the long run (Islam and Madsen 2015). It has been argued that since the 1980s, inequality has been extreme and persistent. Is there an argument that countries which never were directly involved in the war have not been inclined to impose a post-war Egalitarian regime? Is it the case that as a result, the top income shares have been persistent?

To determine whether shocks to inequality are transitory or not, past studies have had to deal with the possibility of whether the inequality data contained a unit root (see for example, Islam and Madsen 2015, Christopoulos and McAdam 2017). Besides, estimation of trends in inequality is also dependent on the presence of a unit root. Perron (1988) 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 modelled 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 lead 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:

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

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 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. 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 (2009a) 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 may 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 (2009a) 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:

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

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

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

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 (2009a) 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 (2009a), 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.

Following this procedure, the residuals $\hat{u}_t$ in (2) are obtained from a regression of y_t on $x_t = (1, t)'$. The super-efficient estimate $\tilde{\alpha}_{MS}$ (obtained as discussed earlier) is used to estimate the quasi-differenced regression

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

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

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

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, et. al. 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:

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

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 comprises of top 1% income shares of eleven countries, being Australia, Canada, New Zealand, USA, France, Sweden, Norway, Japan, Finland, Netherlands, and India. The data spans the period of 1921 to 2000. The two exceptions are India, which begins in 1922 and ends in 1999; and the Netherlands which begins in 1915 and ends in 1999. The data is available at http://www.uueconomics.se/danielw/Data.htm. All the details with regards to the source, description and construction of data can be found in Roine and Waldenstrom (2011). It may be noted that top income share data can be obtained from sources such as the World Wealth and Income Database (WWID) (http://wid/world/data/) which cover the same countries but more recent data up to 2014. However, from the WWID, updated data up to 2014 can be found only for four countries, being Australia, New Zealand, France and USA. Besides, the construction of the top 1% income shares is carried out using a different procedure than that to Roine and Waldenstrom (2011). Further, for various countries, such as Norway, Sweden, Netherlands and New Zealand, there are several periods of missing observations. These problems can be obviated by using the data set provided by Roine and Waldenstrom (2011).

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 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 (2009a) 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 about here]

From the empirical results we can see that the structural change points in the data do not conform to 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 about here]

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 (but not for structural breaks) 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, and New Zealand, 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 bottom half of Table 3 below.

[Table 3 about here]

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 selected countries. Turning to the countries where we found that structural breaks exist, being Canada, USA, Sweden, Norway and Finland, we employ the unit root tests that allow for structural breaks. 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 the known structural breaks in the trend. The results show that even after accounting for structural breaks in both the null and the alternative hypotheses, we cannot 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 not transitory in nature.

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. This does not happen for countries such as Australia or France, which 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). One could argue that the major shocks such as the World Wars and the Great Depression had a persistent effect on income inequality. During the period of the World Wars and the Great Depression, the high taxes had a persistent effect on capital owners with their wealth and income being affected. Alternatively when major structural change occurred in the early 1980s with deregulation and privatisation in many developed countries, inequality increased and persisted over time. The persistence in inequality could be caused by the institutions that greatly advantage the elite in providing access to economic opportunities. Holter (2015) documents several reasons why persistence may exist in top income shares which include the returns to investment in human capital, progressive taxation, and the presence of credit constraints. The finding of persistent inequality can have consequences for distributional mobility and there may be a need for policy intervention.

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Figures and Tables

Figure 1. Top income shares.

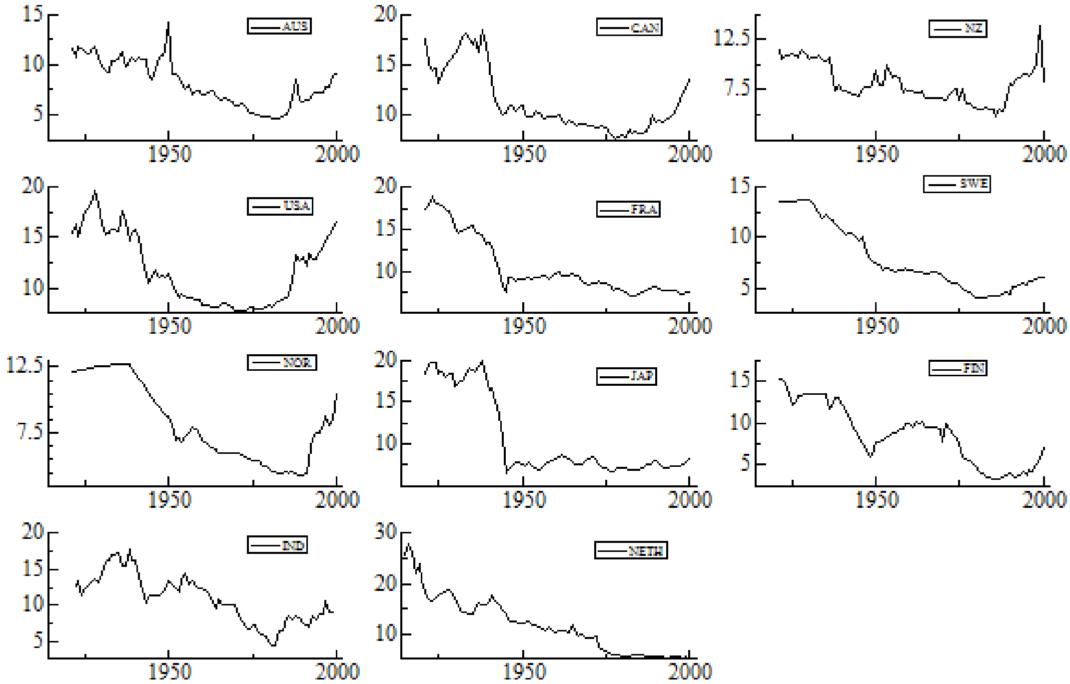

Table 1: Structural Break Test

	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		

***, ** and * denote significance at 1%, 5% and 10% levels respectively.

Table 2 Robust Trend Estimation

	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) 95% conf. int. (–0.0333, 0.0128)	N/A	N/A
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.

Table 3: Unit Root Tests

	MZ_a	MZ_t	MSB	MP_T	$DF-GLS$	P_T
	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).