Income inequality and persistence changes*

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Abstract

This paper carries out a time series analysis of the Gini coefficient for disposable income in a sample that includes both advanced and emerging economies. Our results show that, in most countries, inequality has alternated between stationary and nonstationary regimes during the period 1960-2017. These changes coincide with the implementation of structural reforms and with periods of economic and, especially, financial distress. Our findings also suggest that the persistence of income inequality seems to be related to tax progressivity, income for top earners, and working conditions.

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1 Introduction

The unequal distribution of income and wealth has attracted the attention of academics in the last four decades, especially after the Great Recession. The main reason for this is that economic inequality can lead to political, economic, and social unrest (Alvaredo, Chancel, Piketty, Saez, and Zucman 2018). In fact, one of the goals established in the 2030 Agenda for Sustainable Development, adopted by the member states of the United Nations in September 2015, is to reduce inequality both within and among countries. The importance of adopting an economic standpoint in the study of inequality has become clear as, with greater data availability, researchers have found it is related to growth (Banerjee and Duflo 2003), financial crises (Rajan 2011), happiness (Ferrer-i-Carbonell and Ramos 2014), family composition (Cornelson and Siow 2016) and terrorism (Ezcurra and Palacios 2016), among other issues.

Piketty and Saez (2014) review the long-run evolution of inequality, focusing on Europe and the United States (US). One of the main conclusions drawn by these authors is that inequality does not follow a deterministic process. This claim has fostered a recent strand of the literature that deals with the persistence of income inequality through the study of its order of integration. In particular, Islam and Madsen (2015) applied panel stationarity tests to several measures of inequality in a sample of 21 OECD countries during 1870-2011. These authors did not find evidence against the null hypothesis of trend stationarity, concluding that shocks to income inequality induce temporary effects. This result brings into question the assertion that inequality contains a stochastic trend. In a related work, Christopoulos and McAdam (2017) analyzed the persistence of the Gini coefficient for both gross and net income in an unbalanced panel covering 47 countries over the period 1975-2013. These authors were not able to reject the presence of a unit root, suggesting that the factors driving inequality are likely to have permanent effects[1].

There are several explanations for the conflicting evidence about the persistent nature of income inequality at country level. On the one hand, the studies mentioned above consider different sample compositions and time spans. On the other, two types of methods - stationarity and unit root tests - have been implemented to determine the order of

[1]This same conclusion has been recently reached by Choi (2019) and Gil-Alana, Škare, and Pržiklas-Družeta (2019) using, respectively, univariate standard unit root tests and fractional integration techniques.

integration of inequality measures. Despite this, it is worth noting that both Islam and Madsen (2015) and Christopoulos and McAdam (2017) control for the presence of multiple[2] structural breaks. Even so, they were established assuming a definite order of integration for the variables under scrutiny. In addition, structural breaks have not been taken into account under the null and alternative hypotheses simultaneously. This is equivalent to saying that, to date, the literature on income inequality persistence has not addressed the circular problem encountered when testing for whether a time series contains a broken trend and determining its order of integration, see Perron (2006).

Following previous arguments, and as a first contribution of the present paper, the suitability of a broken trend for income inequality using univariate methods that do not require prior knowledge about the order of integration of the time series analyzed will be established (Kejriwal and Perron 2010; Perron and Yabu 2009). To do so, the Gini coefficient for disposable income in a sample including advanced and emerging economies from 1960 to 2017 will be studied. The estimated number of trend breaks and their corresponding dates will be considered both under the null and the alternative hypotheses when testing for the presence of a unit root in inequality. This will be done applying the procedure developed in Carrion-i-Silvestre, Kim, and Perron (2009), which is based on the implementation of the quasi-generalized least squares (GLS) detrending method proposed by Elliott, Rothenberg, and Stock (1996).

In theory, the Gini coefficient cannot contain a unit root because it is a bounded variable. Christopoulos and McAdam (2017) suggest that the nonstationarity displayed by this measure of inequality in finite samples may be a consequence of its prolonged adjustment towards the long-run trend. It has also been established that the data generated by a persistence change process with no breaks in the deterministic component usually exhibit a simultaneous level and/or trend shift at the persistence change date (Kurozumi 2005). Following these arguments, and as a second contribution of this paper, the extent to which income inequality can be characterized as a process experiencing changes in persistence will be assessed. With this aim, the regression-based test proposed by Leybourne, Kim, and Taylor (2007), which can accommodate multiple changes and is valid regardless of

[2] See Makhlouf (2018) for a related study based on the implementation of univariate time series methods that assume a single change in the deterministic component.

their direction, will be used. This procedure is grounded on the application of doubly-recursive sequences of unit root test statistics *à la* Elliott, Rothenberg, and Stock (1996) and associated breakpoint estimators.

The rest of the paper is structured as follows. Section 2 presents the data source, the composition of the sample, and some preliminary evidence. The methods that have been implemented are explained in Section 3. Section 4 shows the results obtained in the empirical analysis. The main findings are interpreted and discussed in Section 5. Finally, Section 6 concludes.

2 Data

Income inequality will be measured using the Gini coefficient, which reflects the extent to which income deviates from a perfect equal distribution. This would imply that each unit of reference (households, for example) receives an equal share of the income. If this is the case, the coefficient takes a value of zero. On the contrary, a Gini index of one reflects extreme inequality: a single reference unit receives all the income. The data have been extracted from the Standardized World Income Inequality Database[3] (SWIID; Solt 2016) and covers the period 1960-2017. Given that we are adopting a univariate time series approach, the sample that has been analyzed includes the largest advanced (Group of Seven[4], G7) and emerging (Group of Five[5], G5) economies. Further, we have restricted our study to income after tax and transfers, i.e., disposable income.

The main virtue of the SWIID[6] is that it permits comparisons across countries. This is especially relevant when the sample is made up of countries at different levels of development, see Solt (2009). This author analyzed the most frequently used cross-national data sets to study income inequality, concluding that there was a trade-off between comparability and coverage. First, the Luxembourg Income Study Database (LIS) imposes a uniform set of assumptions and definitions to calculate similar inequality measures across countries and over time. Nevertheless, this information is limited and mainly available for richer countries. Second, the World Income Inequality Database (WIID) mixes data

[3]Version 7.1 (August 2018), available at `https://fsolt.org/swiid/`.

[4]Canada, France, Germany, Italy, Japan, United Kingdom (UK) and US.

[5]Brazil, China, India, Mexico and South Africa.

[6]See Solt (2015) and Jenkins (2015) for a debate about the pros and cons, respectively, of using this comprehensive database to study income inequality.

sources of more than 150 countries, without mitigating the above-mentioned trade-off satisfactorily. To deal with this problem, the SWIID combines several databases - taking the LIS as the baseline - to estimate the missing information on income inequality using a model-based multiple imputation approach, see Gelman and Hill (2007). Reflecting the underlying uncertainty of the methods implemented, the SWIID provides 100 imputations for each reported value of inequality measures.

[Insert Figure 1 about here]

Figure 1 plots the time series of the average imputed values of the Gini coefficient for disposable income together with their 95% confidence intervals[7], which show the variability of the multiple imputations. From this point forward, the focus is put on these yearly averages. That is to say, the present study is grounded on estimated rather than observed measures of inequality. Emerging economies display higher coefficients than developed countries, with values that tend to be over 0.4. Surprisingly, income inequality in the US is steadily moving towards this level. South Africa has been the most unequal country, achieving a maximum average imputed Gini index of 0.59 in 2016. The lowest levels of income inequality at the end of the period, with values slightly below 0.3, correspond to France and Germany. Nonetheless, the Gini coefficients in these two countries were even lower in earlier years. With the exception of France, and to a greater or lesser extent, income inequality has increased in advanced economies since the beginning of the 1980s. This upward trend can also be found in China, India and South Africa. Lastly, it is worth noting that France, Italy and, especially, emerging Latin American economies have clearly been able to reduce inequality over the sample period.

The presence of a unit root in disposable income inequality will be analysed firstly by means of standard univariate tests. The alternatives that have been implemented - because they have good size and power - are those discussed in Ng and Perron (2001). These procedures apply a local-to-unity GLS detrending method (Elliott, Rothenberg, and Stock 1996) before testing the unit root null hypothesis. Looking at the time series plotted in Figure 1, the deterministic component has been considered to be made up of both a constant and a linear trend. The number of truncation lags used to augment the auxiliary

[7] $\pm$ 1.96 times the standard deviation.

regression or to obtain the spectral density estimator at frequency zero have been chosen with the modified Akaike information criterion (AIC). These lags permit controlling for the presence of autocorrelation that induces size distortions in unit root tests. The maximum number of lags allowed has been set to four.

[Insert Table 1 about here]

Table 1 presents the results obtained from the application of the standard GLS detrending-based unit root tests to the time series of average imputed Gini coefficients for disposable income. These figures suggest that, with the exception of China, there is little evidence against the unit root null hypothesis at conventional significance levels. This would lead us to conclude that shocks to income inequality generally have permanent effects. Nonetheless, the importance of adequately modelling the deterministic component in time series analysis is well known. As can be observed in Figure 1, the Gini coefficients evolve around a more complicated trend than a linear one in most countries. In line with Christopoulos and McAdam (2017), this evolution may be determining the inability of the standard unit root tests that have been applied to reject the unit root null[8] because they have low power in the presence of changes in the deterministic components (Perron 1989). Moreover, standard unit root tests are not consistent against processes that experience persistence changes (Kim 2003). The study of these two issues in Gini coefficients for disposable income is the main aim of our empirical analysis.

The unit root tests that have been implemented do not account for the variability of the multiple imputations. Unfortunately, uncertainty regarding the estimates of the Gini index provided by the SWIID cannot be introduced in this context in a straightforward way because the imputations do not have a longitudinal dimension. Therefore, the unit root tests cannot be applied to individual time series of imputed Gini coefficients before combining, or jointly interpreting, the resulting statistics at the country level. Figure 1 shows that the confidence bands of the imputed values are wider in emerging economies. This is especially the case in Brazil, Mexico, and South Africa at the beginning of their

[8]This lack of evidence against the null hypothesis may also be related to the fact that unit root test statistics display low power when applied to short time series with a yearly frequency. This problem can be mitigated through the application of panel unit root tests that exploit both the cross-sectional and the temporal dimensions of the data. Nonetheless, the present study is confined to a time series approach because it allows us to analyze the possible presence of persistence changes in disposable income inequality.

corresponding sample periods. This implies that results obtained from a time series analysis of yearly average imputed values are less reliable in this group of countries and, hence, should be considered with more caution.

3 Methodology

3.1 Unit root and trend shift testing

Testing for unit roots is complicated by shifts in the deterministic component. In addition, testing for structural breaks in the trend function depends on the order of integration of the variable under scrutiny (Perron 2006). First, if the variable is considered to be difference-stationary when it is, in fact, stationary in levels, structural break tests will suffer from low power. Second, structural break test statistics applied to a time series in levels will have different limiting distributions depending on its order of integration. This circular problem between tests regarding the parameters of the trend function and unit root tests has motivated the appearance of procedures to detect changes of the deterministic component for univariate time series that are robust to the presence of unit roots. These methods are intended to be used before studying the integration order of a given time series. Trend breaks should be accounted for by unit root tests both under the null and the alternative hypotheses.

The robust test for the stability of the trend function that has been applied in the present paper is that developed in Perron and Yabu (2009, $Exp\!-\!W_{FS}$). In contrast to alternative procedures, it does not involve any random scaling. Moreover, and although this is not necessarily true for local-to-unity processes, this test is the best choice in both the I(0) and I(1) cases. The specification for the shift in the deterministic component of the Gini coefficient for disposable income that has been considered is that corresponding to 'Model III', allowing for a structural change in the intercept and in the slope parameters. The null hypothesis is that of no trend shift. The $Exp\!-\!W_{FS}$ test statistic is based on a feasible GLS procedure that uses a super-efficient estimate of the autoregressive parameter which relies on a truncation when it is equal to one. In addition, the finite sample properties of this method are improved using a bias correction. For an unknown break date, the exponential function of Wald tests for all possible alternatives (Andrews and Ploberger 1994) has very similar limiting distributions in the I(0) and I(1) cases. For this reason, and although this makes the test conservative, the largest critical value is taken for the sake of robustness.

With the aim of searching for the presence of further trend shifts, Kejriwal and Perron (2010) proposed implementing the $Exp{-}W_{FS}$ test in a sequential manner. Their procedure tests the null hypothesis of l breaks against the alternative of $l+1$ breaks, allowing for the consistent estimation of the number of breaks. For the model under the null, break dates are the global minimizers of the sum of squared residuals from an ordinary least squares (OLS) estimation. These dates have been obtained using the dynamic programming algorithm developed by Bai and Perron (2003). The presence of an additional break is tested for in the $l+1$ segments of the resulting partition. The highest value of the corresponding test statistics will be taken as that for the sequential test, denoted as $F_T(l+1|l)$, and compared to asymptotic critical values calculated from the relevant quantiles of the $Exp - W_{FS}$ test limiting distribution.

The majority of the studies that tackle the issue of unit root testing in the presence of shifts in the deterministic component only considered them under the alternative hypothesis of stationarity. Building on Kim and Perron (2009), an extension to GLS detrending-based tests was proposed by Carrion-i-Silvestre, Kim, and Perron (2009) to allow for multiple breaks both under the null and the alternative hypotheses. Its implementation requires obtaining the noncentrality parameter and the critical values using simulations, as they depend on the number of breaks and their location. The unit root tests in the presence of multiple breaks will be applied assuming that their dates correspond to those estimated using the dynamic programming algorithm.

3.2 Detecting multiple persistence changes

As pointed out in the introductory section, the data generated by a process that undergoes a persistence change and without breaks in the deterministic component usually exhibit a simultaneous level and/or trend shift at the break date (Kurozumi 2005). Seminal work on the detection of changes in the order of integration of linear time series can be found in Banerjee, Lumsdaine, and Stock (1992) and Kim (2000). Although the assumption of a single break may be too restrictive in practice, most techniques have been developed for this case, see Perron (2006) and Choi (2015) and the references therein.

Nonetheless, single break tests have low power in detecting processes with multiple shifts in persistence (Kejriwal, Perron, and Zhou 2013).

Following previous arguments, Leybourne, Kim, and Taylor (2007) proposed a method to test for the presence of multiple persistence changes and to consistently estimate the dates at which they take place[9]. This M test assumes that the data have been generated by a time-varying autoregressive process where coefficients and orders are regime-dependent. The null hypothesis that the time series contains a unit root during the sample period is tested against the alternative that there exists, at least, one shift between I(0) and I(1) regimes. One of the main features of this method is that it is valid regardless of the direction of the change. The M test statistic is a doubly-recursive application of ADF-GLS unit root tests (Elliott, Rothenberg, and Stock 1996) and associated break point estimators. The latter refer to the start and end dates of the first stationary regime detected over the whole sample. The lag length has been selected using the sequential method proposed by Ng and Perron (1995) with a significance level of 10%. Once the most prominent stationary regime is established, the presence of any further I(0) partition can be detected by analyzing the remaining I(1) subsamples. In doing so, a minimum temporal dimension of 20 years has been considered. The reason is that this is the lowest number of observations for which finite sample critical values have been calculated by Leybourne, Kim, and Taylor (2007). The sequential application of the M test permits the identification of I(1) regimes between the end point of a stationary partition and the start point of the next one.

4 Results

Table 2 displays the values resulting from the application of the $Exp{-}W_{FS}$ test statistic to the average imputed Gini coefficient for disposable income in the countries that make up our sample. Both at the beginning and at the end of the sample period, 15% of the observations have been disregarded. Taking into account the asymptotic critical values reported in Perron and Yabu (2009) for this trimming choice and the trend shift model that has been specified, the null hypothesis of no broken trend can be rejected at conventional significance levels in most countries. The exceptions are China, South Africa and the

[9] Alternative approaches, based on the implementation of Wald-type test statistics, are Kejriwal, Perron, and Zhou (2013) and Kejriwal (2019).

UK. When the $Exp - W_{FS}$ test is sequentially implemented to detect further shifts in the deterministic component, the null hypothesis of a single break cannot be rejected for Brazil, Canada and the US. Therefore, the linear trend of the Gini indices in these countries should be specified with one shift. Estimated break dates for Canada and the US are around 1980, when income inequality soared, see Figure 2. On the contrary, the structural break experienced by the Gini coefficient of Brazil in 1997 corresponds to a dramatic fall in inequality that has lasted two decades.

[Insert Table 2 and Figure 2 about here]

The sequential applications of the $Exp - W_{FS}$ test proposed by Kejriwal and Perron (2010) provides evidence of multiple trend shifts in India, Japan, Mexico and the European countries. That is to say, the specification of the deterministic component for the time series of Gini coefficients in these countries is more complex. The number of shifts ranges from two in India to the maximum of five that has been allowed in France, Japan and Mexico. With the exception of India, and similarly to Canada and the US, there is a set of estimated break dates around 1980 when there was a change in the decreasing path followed by income inequality at the beginning of the sample period. Our results also suggest that European countries experienced trend shifts at the beginning of the 2000s when Gini indices picked up. Estimated break dates for India correspond to the period from 1996 to 2006 when the increase in income inequality accelerated.

Previous evidence of trend shifts in the average imputed Gini coefficients for disposable income has been taken into account when testing for unit roots in the corresponding countries using the method developed in Carrion-i-Silvestre, Kim, and Perron (2009). Resulting unit root test statistics allowing for multiple breaks in the deterministic component both under the null and the alternative hypotheses are reported in Table 3. The null hypothesis of income inequality evolving around a broken linear trend in a nonstationary manner cannot be definitely rejected in any country. There is some weak evidence against the unit root null, mainly corresponding to the ADF test statistic, for Germany, India and Mexico. These findings allow us to state that, even after accounting for the presence of multiple trend shifts in Gini indices, they evolve following a nonstationary process. This implies that, with the exception of China, the factors that influence income inequality have permanent effects.

[Insert Table 3 about here]

At least initially, the Gini coefficient cannot contain a unit root because it takes values between zero and one, i.e., it is a bounded variable. In a related study, Christopoulos and McAdam (2017) attribute the nonstationary behavior displayed by income inequality time series to their slow convergence towards the long-run trend. As has been explained in the previous section, our proposal consists of testing the null hypothesis that the Gini coefficient is an I(1) process against the more reasonable alternative in the present context that this inequality measure shifts between I(0) and I(1) regimes. The results obtained from the application of the M test are shown in Table 4. The null that income inequality contains a unit root within the whole sample period can be rejected in all countries, except South Africa and the US. It is worth noting that, although the $Exp - W_{FS}$ test detects a trend shift in the Gini index for disposable income in the US at the beginning of the 1980s, this structural break does not seem to correspond to a persistence change.

[Insert Table 4 about here]

The most prominent stationary regime detected in the UK takes place at the beginning of the sample period when inequality was reasonably stable. However, and even if there is no evidence of a trend shift, the M test considers that there is a persistence change in 1980 when inequality soared in this country. On the contrary, the structural break of the average imputed Gini index in China is found at the end of the sample period. As was also the case of the UK, there is no statistical evidence of a trend shift in income inequality in China. Despite this, the persistence change test considers that the Gini index of this country became stationary in 2004. This same result is obtained for Mexico. The I(0) regimes that are initially detected in the other countries cover subperiods between the 1980s and the early 2000s. Estimated stationary regimes have been represented in Figure 2 using shaded areas.

The application of the M test to the resulting subsamples provides further evidence of shifts between I(0) and I(1) regimes in four countries. Those detected in France and

Germany suggest that shocks to income inequality had transitory effects during a similar time period that began in 1980 and lasted around two decades. These additional stationary subsamples cover periods for which trend shifts were established using the sequential implementation of the $Exp - W_{FS}$ test, see Table 2. This is also the case for the I(0) regime detected in Japan during 1974-1986. These results corroborate the argument put forward by Kurozumi (2005) according to which trend shifts are observed around the dates when persistence changes take place. The additional stationary regime found in the Gini index for disposable income in the UK covers the period 2007-2017. Therefore, this is the only advanced country where income inequality was under an I(0) regime at the end of the sample period analyzed.

5 Discussion

Income inequality remained at low levels in developed countries during the postwar era (1945-1980), mainly due to the expansion of the modern welfare state (Schustereder 2010). At the beginning of our sample period, with the exception of the UK, the Gini coefficient for disposable income was under a nonstationary regime in the developed economies that were analyzed. Given that this is also the case for the developing countries considered, it can be stated that income inequality was characterized by a high degree of persistence before the 1980s. This implies that monetary and fiscal policy shocks may have exerted permanent effects on the distribution of income within countries during that period.

It is widely acknowledged that income inequality increased in almost all global regions after 1980. The most common explanations for this 'inequality turn' (Atkinson 2018) - especially in developed countries - are skill-biased technological change, globalization, labor market institutions, and public policies (Alvaredo, Chancel, Piketty, Saez, and Zucman 2018; Battisti and Zeira 2018). The results presented in the previous section suggest that the institutional and policy changes put into place in the early 1980s affected not only the level of income inequality but probably also its degree of persistence. Furthermore, the empirical analysis carried out shows that diverse patterns of inequality persistence dynamics exist across countries, and these may have been determined by their different economic, institutional, and political contexts.

In Canada, the severe recessions that took place in the early 1980s and the early 1990s seem to be linked to shifts in the persistence of the Gini index for disposable income. The corresponding stationary period lasted until 1993, when a change in government entailed a fall in the redistributive impact of taxes and transfers as well as an increase in income among top earners (Green, Riddell, and St-Hilaire 2016). The Gini coefficients of the countries that belong to the euro area also display an intermediate I(0) regime. The structural break is dated to 1980 in both France and Italy. The latter country also underwent a short nonstationary period around the crisis of the European Monetary System in the early 1990s. The first change in the persistence of the Gini index of Germany took place just before its reunification. The presence of a European pattern in the dynamics of income inequality persistence is corroborated by the fact that these three countries experienced a shift around the introduction of the euro. The related I(1) regime at the end of the sample implies that, under the common European currency, monetary policy shocks and the 2008 global financial crisis have likely had long-lasting effects on income inequality (Colciago, Samarina, and Haan 2019). As pointed out by Alvaredo, Chancel, Piketty, Saez, and Zucman (2018), this nonstationary period was also distinguished by changes in labor market institutions and working conditions, and by an increase in the income of top earners, both from labor and capital.

Moriguchi (2017) shows that the structure of inequality and redistribution in Japan is different to that of other developed countries. This also seems to be the case for the dynamics of income inequality persistence. The first structural break detected in the average imputed Gini coefficient of this country takes place during the 1973 oil crisis. The corresponding I(0) regime lasted until 1986, when the economy began an expansion that led to the formation of an asset price bubble. Income inequality in Japan remained stable in another stationary period between the Asian and the global financial crises (1997-2008). In the UK, the Gini index for disposable income behaves as an I(0) process at the beginning of our sample period. This may be a consequence of the economic policies implemented under the "post-war consensus" (Addison 2011). The election of Margaret Thatcher as prime minister in 1979 coincided with the outset of a nonstationary regime. This period was characterized by privatization, deregulation, tax reforms, and changes to labor market institutions. Although the M test detects another I(0) regime between the elections held in 1990 and 2001, the persistence change is not statistically significant. Therefore, we can state that the nonstationary regime of the Gini index in the UK lasted until the recent global financial crisis.

Moving on to developing countries, in China and Mexico income inequality behaves as a nonstationary process during most of the sample period. However, the average imputed Gini coefficients of both countries experience a persistence change in 2004, when they consolidated a downward trend. Similar to the UK, the I(1) regime in China comes hand in hand with a process of privatization of state-owned enterprises and market liberalization. Even though income inequality steadily increases in India throughout the time period analyzed, the Gini index of this country displays an intermediate I(0) regime from 1983 to 1993, when its upward trend becomes less steep. This shift coincides with the introduction of structural reforms and trade openness. The balance of payments crisis and the financial liberalization that took place in the early 1990s concurred with the onset of another nonstationary regime. Income inequality in Brazil also behaves as a stationary process in the middle of the sample. The I(0) regime starts in 1989, when the Gini coefficient for disposable income permanently changes its upward trend. This structural break may be a consequence of the good performance of the labor market and declines in the skill wage premium (Alvaredo, Chancel, Piketty, Saez, and Zucman 2018). A second nonstationary regime begins with an acceleration in the decrease of income inequality in 2001. This period was distinguished by increases in the minimum wage and pensions, cash transfers to the poor, and tax progressivity (Alvaredo, Chancel, Piketty, Saez, and Zucman 2018; Arestis and Sawyer 2018; Engbom and Moser 2018).

To sum up, our empirical analysis shows that, in most of the countries included in the sample, income inequality has experienced persistence changes. More importantly, the estimated dates for these shifts coincide with the implementation of new policies[10] as well as with periods of economic and financial distress. The exceptions are South Africa and the US, where the Gini coefficient for disposable income follows an upward trend throughout the sample period. In the case of South Africa, this is likely a reflection of the deeply unequal socioeconomic structure created by apartheid. The nonstationary behavior of

[10] Gil-Alana, Škare, and Pržiklas-Družeta (2019) have also pointed at policies as a source of inequality persistence.

inequality in the US may have been determined by the upsurge of income for top earners, and the less progressive character of taxes over the last decades. Therefore, our results also allow us to state that, apparently, the persistence of income inequality tends to increase with lower tax progressivity, higher income for top earners, and worse working conditions.

6 Concluding remarks

This paper contributes to the recent literature on the persistence of income inequality measures at the country level. Information extracted from a database that allows for cross-country comparisons was used and a univariate time series approach was adopted. More specifically, the evolution of average imputed Gini coefficients for disposable income in both advanced and emerging economies was studied for 1960-2017. As a first contribution, this paper circumvents the circular problem of testing for the presence of breaks in the deterministic component of income inequality and of establishing its order of integration. This was done by assessing the suitability of a broken trend for Gini indices using a robust test statistic. These results have been taken into account when testing for the unit root character of income inequality both under the null and the alternative hypotheses.

Even after controlling for multiple trend shifts and with the exception of China, Gini coefficients display nonstationary behavior. In principle, the Gini index cannot contain a unit root because it is a bounded variable. Moreover, trend breaks are known to be observationally equivalent to persistence changes. These two arguments have encouraged us to test the unit root null hypothesis of income inequality during the sample period against the alternative that it has undergone at least one shift between a stationary and a nonstationary regime. The multiple persistence change test statistic that has been implemented shows that Gini indices have experienced structural breaks in all countries except South Africa and the US. These changes coincide with policy shifts and with periods of economic and, especially, financial difficulties. Furthermore, our results suggest that the persistence of income inequality seems to be related to tax progressivity, income for top earners, and working conditions.

Although we have discussed some historical events that may be linked to the structural changes detected and the persistence of income inequality, our empirical analysis does not allow us to establish causal relationships. In addition, there are other factors not considered in our discussion - such as demographics or technological change - that can also affect the persistence of income inequality. Therefore, a promising avenue of research is the study of its determinants. In light of the present work, and in an effort to overcome its limitations, there are several issues to take into account when dealing with this topic. The inclusion of a greater number of countries in the sample would make the analysis more representative. The consideration of a longer time period would also increase the sample size, making inferences more reliable. In this line, analysis of an observed, rather than imputed, measure of inequality would avoid estimation uncertainty regarding the dependent variable.

Apparently, most of the Gini coefficients that have been studied are under an I(1) regime at the end of the sample period. This implies that, at present, fiscal and monetary policy shocks, as well as downturns, will probably exert long-lasting effects on the distribution of income. This will also be the case for the fourth industrial revolution and the resulting transformation of society, which is currently taking place at a global scale. Although income inequality should be considered to be a global phenomenon, within-country inequality is more important from a political point of view (Milanovic 2016). In line with the goals established in the 2030 Agenda for Sustainable Development, tackling within-country inequality is crucial for wiping out global poverty (Alvaredo, Chancel, Piketty, Saez, and Zucman 2018). With this aim - and similar to Alvaredo, Chancel, Piketty, Saez, and Zucman (2017) - the empirical analysis carried out in the present paper leads us to conclude that educational policies, the use of progressive taxes and transfers, and the reform of labor market institutions could be useful in restraining not only the level but also the persistence of income inequality.

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Table 1: Standard univariate unit root testing: Average imputed Gini coefficients for disposable income, 1960-2017.

Country	Sample period	P_t	MP_t	ADF	Z_α	MZ_α	MSB	MZ_t
Brazil	1960-2015	17.755	16.872	-1.210	-5.388	-5.133	0.270	-1.386
Canada	1971-2016	55.603	38.681	-1.782	-3.870	-2.531	0.460	-1.082
China	1978-2015	3.620***	3.709***	-2.186	-29.746***	-29.663***	0.126***	-3.738***
France	1970-2015	9.315	9.601	-1.988	-10.004	-9.955	0.214	-2.129
Germany	1960-2015	49.827	47.498	-0.579	-1.374	-1.151	0.471	-0.543
India	1976-2012	17.500	16.485	-1.743	-6.038	-5.516	0.299	-1.649
Italy	1967-2015	15.496	13.976	-1.816	-6.979	-6.529	0.266	-1.737
Japan	1961-2014	23.432	20.066	-1.269	-4.971	-4.488	0.327	-1.469
Mexico	1963-2016	18.995	17.873	-1.633	-5.342	-5.094	0.313	-1.593
South Africa	1975-2015	15.679	14.097	-1.668	-7.123	-6.465	0.273	-1.765
United Kingdom	1961-2017	14.066	14.369	-1.361	-6.356	-6.333	0.261	-1.653
United States	1961-2016	37.581	30.520	-1.500	-3.673	-2.972	0.408	-1.214

Note: Reported statistics correspond to the GLS detrending-based unit root tests discussed in Ng and Perron (2001). The deterministic component is made up of a constant and a linear trend. Setting the maximum number of augmentation lags to four, which has been selected using the modified AIC. * $p<0.10$, ** $p<0.05$, *** $p<0.01$.

Table 2: Trend shift test statistics and estimated break dates.

Country	$Exp-W_{FS}$	$F_T(2\|1)$	$F_T(3\|2)$	$F_T(4\|3)$	$F_T(5\|4)$	TB_1	TB_2	TB_3	TB_4	TB_5
Brazil	7.115***	1.788				1997				
Canada	11.931***	2.326				1976				
China	2.473									
France	4.300**	7.710***	25.899***	32.171***	4.681**	1977	1983	1989	1995	2006
Germany	8.765***	3.460*	117.791***	44.252***	3.802	1970	1980	1996	2004	
India	3.432**	14.932***	3.253			1996	2002			
Italy	4.774***	183.104***	183.104***	183.104***	1.502	1984	1992	2003		
Japan	4.715***	14.102***	5.006**	3.997*	3.997*	1971	1979	1990	1998	2006
Mexico	2.737*	22.279***	11.579***	19.581***	19.581***	1975	1983	1991	1999	2007
South Africa	0.750									
United Kingdom	1.592									
United States	3.175**	2.379				1981				

Note: This table reports the test statistic of Perron and Yabu (2009) and its sequential implementation to detect multiple shifts in trend, see Kejriwal and Perron (2010). The trimming parameter has been set to 0.15. Break dates TB_l are the global minimizers of the sum of squared OLS residuals from the model with l trend shifts. * p<0.10, ** p<0.05, *** p<0.01.

Table 3: Unit root testing in the presence of multiple trend shifts.

Country	P_t	MP_t	ADF	Z_α	MZ_α	MSB	MZ_t
Brazil	14.235	14.134	-2.389	-10.989	-9.886	0.217	-2.149
Canada	23.253	24.122	-1.427	-5.901	-5.121	0.264	-1.353
France	27.488	27.720	-4.100	-26.167	-17.893	0.157	-2.804
Germany	16.654	15.613	-5.250***	-36.764	-24.463	0.143	-3.493
India	13.449	12.653	-4.211**	-24.774*	-16.242	0.170	-2.761
Italy	28.662	24.065	-3.252	-15.707	-11.647	0.207	-2.412
Japan	25.795	25.071	-3.785	-24.962	-18.714	0.152	-2.834
Mexico	24.004	20.439	-4.471*	-31.314	-21.763	0.151	-3.295
United States	20.748	16.516	-2.528	-11.597	-10.124	0.218	-2.206

Note: Reported unit root test statistics allow for multiple trend shifts under both the null and the alternative hypotheses, see Carrion-i-Silvestre, Kim, and Perron (2009). Setting the maximum number of augmentation lags to four, which has been selected using the modified AIC. Break dates are those reported in Table 2. * $p<0.10$, ** $p<0.05$, *** $p<0.01$.

Table 4: Testing for multiple persistence changes.

Country	Period	M test	I(0) start	I(0) end
Brazil	1960-2015	-5.579**	1989	2001
	1960-1988	-4.975	1966	1979
Canada	1971-2016	-26.329***	1983	1993
	1994-2016	-4.394	2000	2012
China	1978-2015	-33.161***	2004	2014
	1978-2003	-4.273	1991	2003
France	1970-2015	-50.536***	1992	2002
	1970-1991	-16.312***	1980	1991
Germany	1960-2015	-15.481***	1989	2001
	1960-1988	-3.257	1969	1982
India	1976-2012	-93.941***	1983	1993
	1994-2012	-4.056	1997	2007
Italy	1967-2015	-12.681***	1980	1990
	1991-2015	-18.935***	1993	2003
Japan	1961-2014	-10.712***	1997	2008
	1961-1996	-8.489***	1974	1986
Mexico	1963-2016	-9.570***	2004	2015
	1963-2003	-3.153	1985	1999
South Africa	1975-2015	-4.359	1982	1993
United Kingdom	1961-2017	-5.819**	1961	1979
	1980-2017	-6.616**	2007	2017
	1980-2006	-4.676	1990	2001
United States	1961-2016	-4.431	1977	1990

Note: This table reports the test statistic for the change in the order of integration developed by Leybourne, Kim, and Taylor (2007) and the underlying trend-stationary regime. The lag length has been selected using the sequential method proposed in Ng and Perron (1995) and a significance level of 10%. The deterministic component is made up of a constant and a linear trend. * $p<0.10$, ** $p<0.05$, *** $p<0.01$.

Figure 1: Gini coefficient for disposable income, 1960-2017: Average imputed values and 95% confidence intervals.

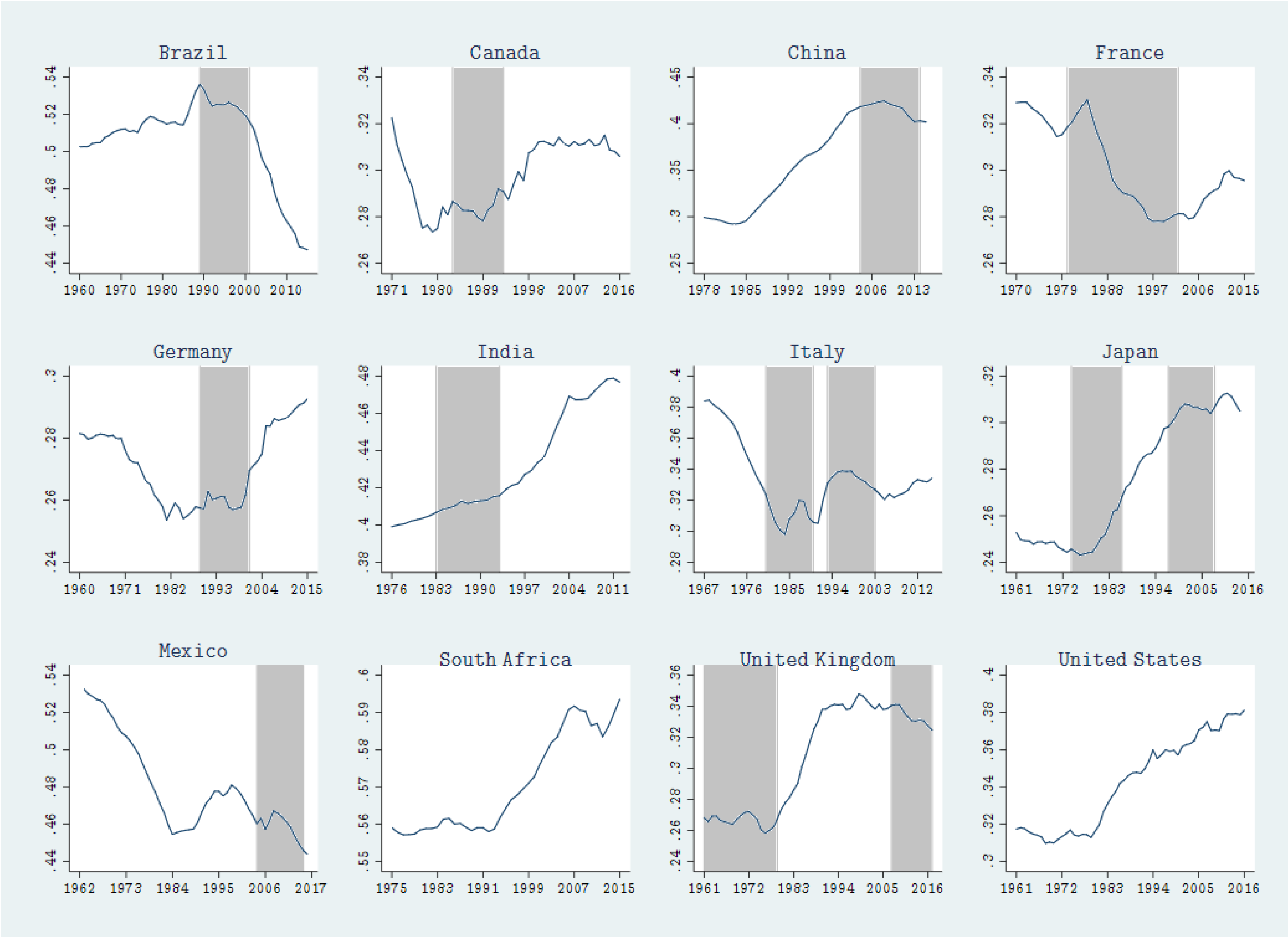

Figure 2: Income inequality and persistence changes. Grey areas represent I(0) regimes.