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ON THE PERSISTENCE OF CROSS-COUNTRY INEQUALITY MEASURES

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On the Persistence of Cross-Country Inequality Measures

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Abstract

We examine inequality persistence in a multi-country unbalanced panel framework using a battery of stationarity and long-run memory tests. Inequality is measured by the (gross and net) Gini indices of income inequality. Results suggests that we cannot reject a unit root in the inequality measures. This applies to both gross and net indices: thus whilst redistributive measures have reduced the level of inequality, they have not sufficiently modified its apparent unit root. A more likely conclusion is that inequality measures are exceptionally persistent if not strictly speaking a unit root. Thus shocks to inequality have very long-lasting effects. We also introduce a new panel stationarity test useful for series subject to unknown structural breaks.
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1 Introduction

In recent decades and across many countries, inequality – as measured by the Gini index – has risen (e.g., Guest and Swift (2008)). Accordingly, it has become a subject of growing concern for economists and policy makers (inter alia, Piketty and Saez (2003); Rajan (2010); Ostry et al. (2014)).[1] This concern reflects not only social norms of fairness but also the concern that extremes of inequality undermine growth and incentives (Atkinson, 2015).

Our contribution is to examine the persistence of inequality measures. If inequality is highly persistence then innovations (e.g., technology or financial shocks) have long-lasting effects (in the limit, permanent effects). Since the Gini index is bounded, one's prior might be that it does *not* contain a unit root. However, even if inequality is ultimately mean reverting, its adjustment path may be highly protracted.[2] This matters for policy: it may warrant pronounced redistributive or risk-sharing responses, or derail other objectives.

We examine inequality persistence in a *multi-country* heterogeneous panel framework. We do so for two reasons. First, the recognised trends in inequality across countries warrants such a perspective.[3] Second, in contrast to the low power of individual unit-root tests, panel unit-root tests more powerfully combine cross-section variation with time series information.

The paper proceeds as follows. Section 2 discusses the data. Gini indices measuring income inequality are taken from Solt (2009). We also look at both gross and net (i.e., after taxes and transfers) inequality measures. This allows us to gauge whether redistribution policies, whilst clearly reducing the *level* of inequality, have also modified its dynamics. Section 3 reviews the unit-root tests for (unbalanced) panels. We examine persistence using increasingly robust stationarity and long-run memory tests. First, we use panel Dickey-Fuller (DF) tests using the Fisher-type tests of Maddala and Wu (1999) and Choi (2001). We then use the non-linear IV method of Chang (2002), Chang and Song (2009) which allows for the presence of cross cointegration, as well as for cross-section correlation. This is followed by the Pesaran et al. (2013) test which allows a multi-factor structure of the cross-correlation.

[1]See also the symposia in, e.g., the Journal of Economic Perspectives (2015, 9, 1), Review of Economic Dynamics (2010, 13, 1).

[2]A unit root is unbounded and can wander arbitrarily far from its initial point. Nonetheless, many persistent series are often treated as (near) unit-root processes despite their apparent boundedness (e.g., nominal rates cannot be strongly negative; unemployment rate is roughly a percentage). In our case, a unit-root process would be rejected by subject-matter reasoning, in particular the presence of a time trend in its data generation process (DGP). However in "medium-run" samples there can be drifts and trends which can be nonetheless highly informative about dynamic characteristics.

[3]Papers examining Gini stationarity at the country level include Jacobsen and Giles (1998), Maestri and Roventini (2012).

We additionally implement a semi-parametric strategy in checking the existence of a unit-root process based on fractional integration (Shimotsu and Phillips (2004), Lobato and Velasco (2007)). By allowing the order of integration to take fractional values, we allow data to be mean reverting but to still have long memory in the process.

These tests are also more powerful in detecting unit-root behavior when the actual DGP is unknown, especially in the presence of incidental trends or where some breaks and threshold nonlinearity occur, see Smallwood (2015).

None of the above tests formally allows for structural breaks (SBs); testing persistence whilst ignoring breaks risks over acceptance of the unit-root null. Accordingly, building on Enders and Lee (2012), we propose a new test which admits (in form and number) unknown SBs in an unbalanced panel.

Section 4 concludes: all tests demonstrate that a unit root in inequality cannot be rejected. Thus shocks to income inequality have permanent (or at least very long-lasting) effects; inequality tends to increase (or decrease) over time in a secular manner.

2 Data

Income inequality is captured by the Gini index from the Standardized World Income Inequality Database (SWIID), Solt (2009). This provides extensive coverage of internationally comparable income inequality data (173 countries, 1960-2012/13). The SWIID standardizes data comes from multiple sources (e.g. the United Nations University's World Income Inequality Database, the OECD's Income Distribution Database and the Socio-Economic Database for Latin America and the Caribbean by CEDLAS and the World Bank, as well as data from several national statistical offices). It is currently the best suited dataset to perform cross-national research on income inequality (given its emphasis on cross-country comparability).[4]

We employ an unbalanced panel dataset for $N = 47$ countries. The countries were selected having at least 30 data points and no in-sample NAs. This gives a sample from 1975. **Figure 1** shows the gross and net Gini measures. The index lies between 0 (perfect equality) and 100 (perfect inequality).

In many cases, both series are trending upwards, and strongly so in proportionate terms for some countries (i.e., Australia, Bulgaria, China, New Zealand, the UK). In others, there have been sustained reductions in income inequality (e.g., Kenya, Mexico, Turkey). The most unequal countries (in gross terms) within our sample are South Africa, Brazil, Kenya, Colombia. The most equal tend to be Bulgaria, Indonesia, Japan, Jordan, Korea, Norway, Pakistan, Taiwan. Further, in most cases, the ratio of average gross-to-net

[4]We are constrained to use aggregate Ginis, rather than also particular percentiles (such as the top/bottom 10%) since these are not available on an internationally-comparable basis.

Gini strongly exceeds 1 (e.g. Belgium, Finland, Germany, Netherlands, Norway, Sweden) whilst in others, reflecting weak or ineffective redistributive schemes, it is around 1 (e.g., Bulgaria, Chile, China, Colombia, India, Mexico, Sri Lanka, Taiwan). Summary Statistics are provided in appendix **Table A1**.

3 Methodology

We now examine (unbalanced) panel unit-root tests following Maddala and Wu (1999), Choi (2001), Chang (2002), Chang and Song (2009) and Pesaran et al. (2013).

3.1 Panel Unit-Root tests without Breaks

To test the stationarity properties of the Gini index in a panel-data setting, we initially use the Maddala and Wu (1999) which can be employed in an unbalanced dataset contrary to other well-known (symmetric) panel unit-root tests (e.g., Harris and Tzavalis (1999), Levin et al. (2002)). The regression for each i^{th} cross-section unit (i.e., country) is,

$$\Delta Gini_{it} = \alpha_i + \beta_i t + \phi_i Gini_{it-1} + \sum_{\ell} \delta_{i\ell} \Delta Gini_{it-\ell} + e_{it}, \quad i = 1, \ldots N; t = 1, \ldots T \tag{1}$$

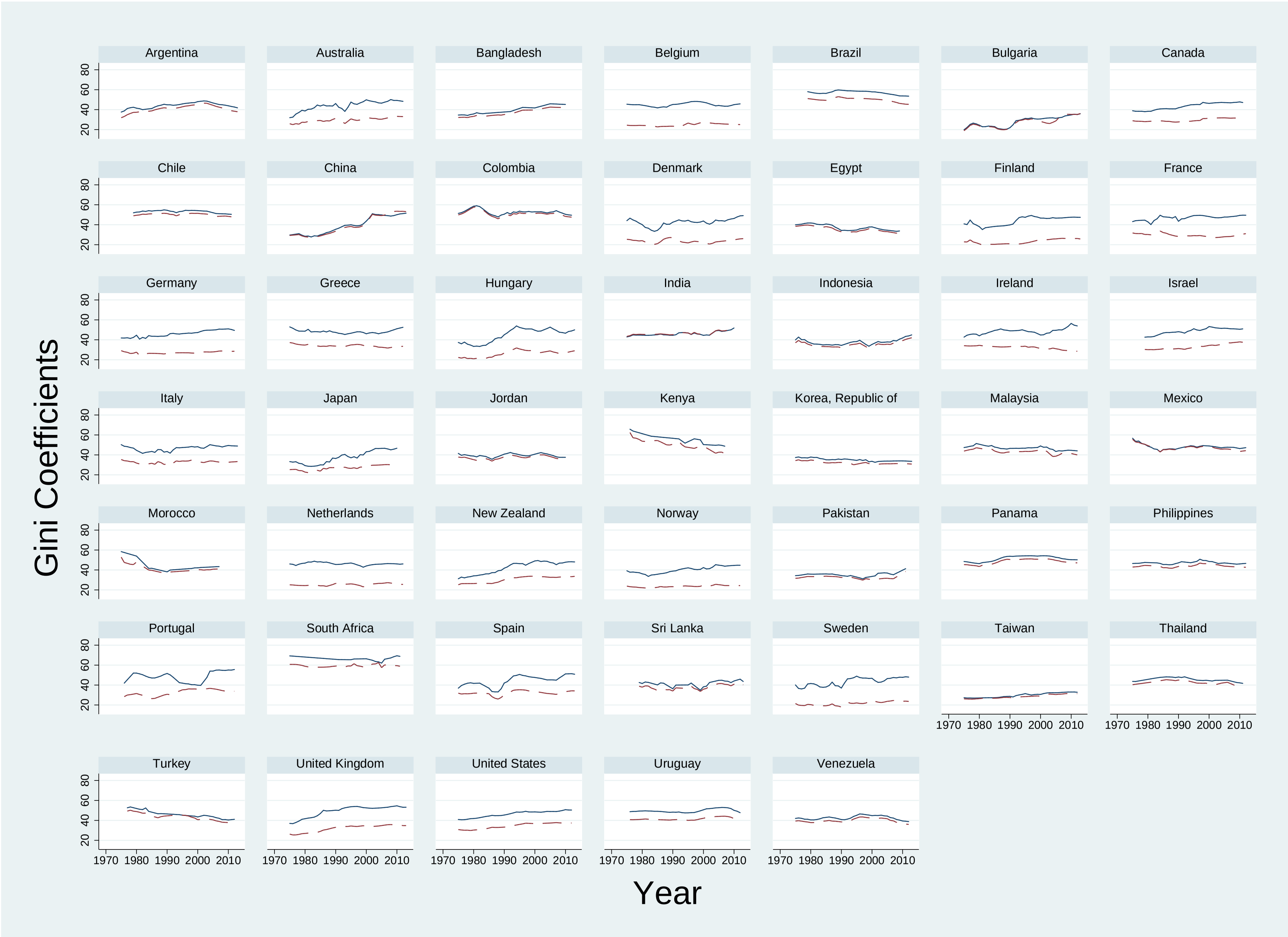
where $\ell \geq 1$ denotes the lags required to ensure white noise errors in e_{it}. The null hypothesis of a unit root ($H_0 : \phi_i = 0 \; \forall i$ against $H_1 : \phi_i < 0 \; \forall i$) can be tested using the formulae suggested by Maddala and Wu (1999) and Choi (2001):

$$P : -2 \sum_{i=1}^{N} \log(p_i) \xrightarrow{d} \chi^2(2N)$$
$$P_m : -\frac{1}{\sqrt{N}} \sum_{i=1}^{N} [\log(p_i) + 1] \xrightarrow{d} \mathcal{N}(0, 1)$$
$$Z : \frac{1}{\sqrt{N}} \sum_{i=1}^{N} \Phi^{-1} p_i \xrightarrow{d} \mathcal{N}(0, 1)$$

where p_i is the bootstrapped probability value from the t-ratio t_{ϕ_i}, and Φ is the standard normal CDF.[5]

[5] Choi (2001) suggests this modified-P test, P_m, is more appropriate for large N but that the Z test works best overall.

Figure 1: Gini Coefficients, Gross and Net



Notes: Solid (dashed) line indicates Gross (Net) Gini coefficients.
Source: Solt (2009), version 5.

3.2 Panel Unit-Root tests With Breaks

Testing for a unit root ignoring valid SBs in the DGP leads to a significant loss of power, Perron (1989). Accordingly, we develop a new panel unit-root test allowing for their presence. Assume the following DGP,[6]

$$Gini_t = \alpha(t) + \phi Gini_{t-1} + a_{time}t + \varepsilon_t \tag{2}$$

where $\alpha(t)$ is a time-varying, deterministic mean. The unit-root null corresponds to $\phi = 1$. However, since $\alpha(t)$ is unknown this null cannot be tested. Becker et al. (2004) used a Fourier series expansion to approximate $\alpha(t)$:

$$\alpha(t) = \alpha_o + \sum_{\Bbbk=1}^{\mathbb{K}} \left[\alpha_1^{\Bbbk} \sin\{\frac{2\pi\Bbbk t}{T}\} + \alpha_2^{\Bbbk} \cos\{\frac{2\pi\Bbbk t}{T}\} \right] \tag{3}$$

where $\Bbbk$ is the number of frequencies of the Fourier function, and $\pi = 3.14$.

According to Becker et al. (2004) breaks can be identified using the low-frequency components of a Fourier expansion. The main advantage of the Fourier series expansion is that it can approximate an unknown number of breaks of unknown functional form contrary to that assumed by standard break tests, e.g., Perron (1990), Bai and Perron (2003), Lee and Strazicich (2003). Note that in this specification the breaks are modelled as smooth processes rather than level shifts. $\alpha_1^{\Bbbk} = \alpha_2^{\Bbbk} = 0$ retrieves the standard DF linear model without structural change.

A problem related to specification (3) is to identify the appropriate number of frequencies to include. We follow Ludlow and Enders (2000) and Enders and Lee (2012) who showed that a single frequency, $\mathbb{K} = 1$ is sufficient to approximate the Fourier expansion in empirical applications. According to Becker et al. (2004) equation (2) under specification (3) has more power to detect several smooth breaks of unknown form in the intercept than the standard Bai and Perron (2003) multi-break tests.

Given model (2)-(3), Enders and Lee (2012) suggested a lagrange-multiplier methodology in testing the unit-root null $\phi = 1$ against $\phi < 1$. First estimate,

$$\Delta Gini_t = \alpha_o + \alpha_1 \Delta \sin\{\frac{2\pi\Bbbk t}{T}\} + \alpha_2 \Delta \cos\{\frac{2\pi\Bbbk t}{T}\} + v_t \tag{4}$$

Then construct the detrended series,

$$Z_t = Gini_t - R_t - \widehat{\alpha}_0 t - \widehat{\alpha}_1 \sin\{\frac{2\pi\Bbbk t}{T}\} + \widehat{\alpha}_2 \cos\{\frac{2\pi\Bbbk t}{T}\} \tag{5}$$

[6]To avoid the Fourier function having the same starting and ending values, the introduction of a time trend is necessary. Thus, changes in the constant and in the slope of the deterministic function are captured by the Fourier approximation.

where $R_t = Gini(1) - \widehat{\alpha}_0 - \widehat{\alpha}_1 \sin\{\frac{2\pi \Bbbk t}{T}\} + \widehat{\alpha}_2 \cos\{\frac{2\pi \Bbbk t}{T}\}$ and where $Gini(1)$ is the first observation of $Gini_t$. Finally, estimate,

$$\Delta Gini_t = \gamma_0 + \phi_Z Z_{t-1} + \gamma_1 \Delta \sin\{\frac{2\pi \Bbbk t}{T}\} + \gamma_2 \Delta \cos\{\frac{2\pi \Bbbk t}{T}\} + \sum_{j=1}^{J} \gamma_j^* \Delta Z_{t-j} + \varepsilon_t \qquad (6)$$

The null of a unit root $H_0 : \phi_Z = 0$ can be tested against $H_1 : \phi_Z < 0$. Since the distribution of this test is non-standard, critical values for various frequencies $\Bbbk$ are reported in Enders and Lee (2012).

We extend the Enders-Lee time-series test to the case where inference is based on a panel. The new test, $DCPM$, is applied to a panel setting by combining the $p-$values of the individual $t-$statistics for a unit root following a non-parametric Fisher-type test,

$$DCPM : -2 \sum_{i=1}^{N} \log(p_i) \xrightarrow{d} \chi^2(2N) \qquad (7)$$

where p_i is the probability value from the t-test $t_{\phi_{Z_i}}$. Monte Carlo simulations were used to derive the probability values of this test. The panel unit-root hypothesis can be now interpreted as:

$$H_0 : \phi_{Z_i} = 0 \ \forall i \in [1, \ldots N]$$
$$H_1 : \phi_{Z_i} < 0 \ \exists i \in [1, \ldots N]$$

This new panel data test has some key advantages: (a) it does not require a balanced dataset; (b) it is possible to allow for different lag lengths in the individual DF regressions; (c) it allows us to make inference based on a combination of series with $[\alpha_1^{\Bbbk} \neq \alpha_2^{\Bbbk} \neq 0]$ and without $[\alpha_1^{\Bbbk} = \alpha_2^{\Bbbk} = 0]$ any structural change[7]; and (d) it permits unlike standard panel unit-root tests which allow for a break (see Carrion-i-Silvestre et al. (2005) and Im et al. (2005)), that the individual DF regressions have different number of breaks.

3.3 Chang-Song Panel Unit-Root test

To construct more powerful unit-root tests appropriate for small samples, designed either for symmetric or asymmetric panels, Chang (2002) and Chang and Song (2009) suggest unit-root tests allowing for the presence of cross cointegration, as well as for cross-section correlation. Failure to account for cross-section correlation lead to large size distortions

[7]Thus the results are valid even in the case where no breaks occur under the null for a subset of units.

in panel unit-root tests, e.g., O'Connell (1998). Thus, we have the regression,

$$Gini_{it}^{\mu t} = \rho_i Gini_{it-1}^{\mu t} + \eta_{it} \tag{8}$$

where η_{it} is an error term and $Gini_{it}^{\mu t}$ represents the demeaned-detrended Gini series (μt):

$$Gini_{it}^{\mu t} = Gini_{it} - \widehat{\alpha}_i^* - \widehat{\beta}_i^* t - \widehat{\gamma}_i^* Gini_{it-1} - \sum_{l=1}^{L} \widehat{\delta}_{il}^* \Delta Gini_{it-l} \tag{9}$$

To test the unit-root null $H_0 : \rho_i = 1 \; \forall i$ against $H_1 : \rho_i < 1 \; \forall i$ in a panel setting they showed that equation (8) can be estimated by non-linear IV OLS using as instruments non-linear transformations of the lagged levels:

$$\Pi_i(Gini_{it-1}) = Gini_{it-1} e^{-\sigma_i |Gini_{it-1}|} \tag{10}$$

where $\sigma_i = K T_i^{-0.5} \psi^{-1}(\Delta Gini_{it})$, and $\psi^2(\Delta Gini_{it}) = T_i^{-1} \sum_{t=1}^{T} (\Delta Gini_{it})^2$ and where K is a constant for every $i = 1, ...N$ and then using the standardized sum of the individual $t-$ratios to generate the statistic:

$$S = N^{-1/2} \sum_{i=1}^{N} t_{\rho_i} \xrightarrow{d} \mathcal{N}(0,1) \tag{11}$$

Parameter σ_i is crucial for the properties of the test as $\Pi_i(Gini_{it-1}) \in [-(\sigma_i e)^{-1}, (\sigma_i e)^{-1}]$ with $Gini_{it-1} \in [-\frac{1}{\sigma_i}, \frac{1}{\sigma_i}]$. Accordingly σ_i must be proportional to the inverse of the standard deviation of the $\Delta Gini_{it}$. To avoid over rejection of the unit-root null when T is small we follow Chang (2002) and use a larger K to correct for the size distortions.[8] Finally this test is robust to cross-section dependence and cross cointegration.

3.4 Pesaran-Smith-Yamagata test

Although the Chang and Song (2009) test allows for both the presence of cross-sectional correlation and cross cointegration it does not assume a multi factor structure of the cross-correlation. This leads to size distortions and potentially misguided inference. In addition this test is appropriate for small panels with relatively large T.

To this end, Pesaran et al. (2013) have proposed a panel unit-root test, $CIPS$, adapted to take into account the multifactor structure of the errors. In doing this, they utilize the information contained in a number of additional covariates that together are assumed

[8] To make our results invariant to the initial choice of the K value we used values between $1 - 8$. The results reminded qualitatively similar independent from the value of K. In light of this, we set $K = 5$.

to share the common factors of the series of interest. Thus, the resulting augmented DF regression is augmented with the cross-sectional averages of the series of interest and the additional covariates. In particular they specify the following ADF regression:

$$\Delta Gini_{it} = \alpha_i^* + \beta_i^* t + \phi_i^* Gini_{it-1} + \sum_{\ell} \delta_{i\ell}^* \Delta Gini_{it-\ell} + \nu_{it} \tag{12}$$

where,

$$\nu_{it} = \boldsymbol{\gamma'_{i,Gini}} \boldsymbol{\zeta_t} + w_{i,Gini,t} \tag{13}$$

where $\boldsymbol{\zeta}$ is a $q \times 1$ vector of unobserved common effects following a covariance stationary process; $\gamma'_{i,Gini}$, is a vector of factor loadings; and $w_{i,Gini,t}$ is an idiosyncratic component.[9]

Substituting (13) into (12) yields:

$$\Delta Gini_{it} = \alpha_i^* + \beta_i^* t + \phi_i^* Gini_{it-1} + \sum_{\ell} \delta_{i\ell}^* \Delta Gini_{it-\ell} + \boldsymbol{\gamma'_{i,Gini}} \boldsymbol{\zeta_{it}} + w_{i,Gini,t} \tag{14}$$

The null of a unit root,

$$H_0 : \phi_i^* \ = \ 0 \ \forall i \in [1, \dots N]$$

is tested against,

$$\begin{aligned} H_1 : \phi_i^* &< 0 \ \forall i \in [1, \dots N_1] \\ \phi_i^* &= 0 \ \forall i \in [1, \dots N_1 + 1, \dots N] \end{aligned}$$

where $N_1/N \to c \in (0, 1]$ as $N \to \infty$. A test of the panel unit-root hypothesis can be based on the individual t-ratio t_{ϕ_i} derived from OLS estimation of,

$$\Delta Gini_{it} = \alpha_i^* + \beta_i^* t + \phi_i^* Gini_{it-1} + \sum_{l=1}^{L} \delta_{il}^* \Delta Gini_{it-l} + \zeta_i' \overline{Gini}_{it-1} + g_i' \Delta \overline{Gini}_{it} + \varpi_{it} \tag{15}$$

where $\overline{Gini}$ is the cross-section average. The resulting panel unit-root test is the average of the average ratios:

$$CIPS = \frac{1}{N} \sum_{i=1}^{N} t_{\phi_i} (N, T) \tag{16}$$

This test by allowing the case of a multi-factor error structure is shown to have the stable size for all combinations of cross-section units and time series dimensions considered.

[9] To allow for unobserved common effects in a panel context (i.e., common factors affecting all the sampled countries) we employ some stationary covariates; in our case these are the cross-section averages.

3.5 A Fractional Approach

Standard panel unit-root tests have lower power in detecting the null unit-root hypothesis when the true DGP is unknown, or some breaks and threshold non-linearity occur, Smallwood (2015). To overcome this we slightly extend the unit-root test of Lobato and Velasco (2007) in a panel setting. In particular, we check for the existence of a unit root based on the long-memory approach in fractional integrated series advanced by Lobato and Velasco (2007).

The fractional integration approach is more general than the standard parametric unit-root tests in the sense that the integration parameter, d can be any real number. In particular the fractional process is defined by,

$$y_t = (1-L)^{-d} v_t = \sum_{k=0}^{t-1} \frac{\Gamma(k+d)}{\Gamma(d)\Gamma(k+1)} v_{t-k} \tag{17}$$

where $\Gamma(\cdot)$ is the generalized fractional function, L is the lag operator and v_t is a short-memory process. Next we describe a three step strategy in order to test for a fractional unit root process.

(1) We use the Exact Local Whittle (ELW) estimator of Shimotsu and Phillips (2004) to estimate d. The ELW estimator is consistent, asymptotically Normal and robust even for a non-stationary process. The estimate of d can be obtained by using the Whittle likelihood function:

$$Q_m(G,d) = -\frac{1}{m} \sum_{j=1}^{m} [\log(G\omega_j^{-2d}) + \frac{1}{G} I_{(1-L)^d Gini_t}(\omega_j)], \; j = 1, \ldots m. \tag{18}$$

where $I_{(1-L)^d Gini_t}(\omega_j)$ is the periodogram of the fractional difference of the Gini; ω_j are the set of Fourier frequencies, $2\pi j/T$; and where G is $f(\omega)$ the spectral density; as $\omega \to 0$ and $m < T$ is the bandwidth parameter which must satisfy $\frac{1}{m} + \frac{m}{T} \to 0$ as $T \to \infty$ (and thereby focuses attention on the long run).

The estimated integration parameter is then given by,

$$\widehat{d} = \arg \min_{d \in [\Delta_1, \Delta_2]} \Re(d) \tag{19}$$

where $-0.5 < \Delta_1 < \Delta_2 < \infty$, and

$$\mathfrak{R}(d) = \log(\widehat{G}(d)) - \frac{2d}{m}\sum_{j=1}^{m}\log(\omega_j)$$
$$\widehat{G}(d) = \frac{1}{m}\sum_{j=1}^{m}\omega_j^{2d} I_{Gini}(\omega_j)$$

(2) Following Lobato and Velasco (2007) we run the DF regression:

$$\Delta Gini_t = \xi \frac{(1-L)^{\hat{d}-1} - 1}{1-\hat{d}} Gini_{t-1} + \sum_{\ell} \xi_\ell^* \Delta Gini_{t-\ell} + v_t \tag{20}$$

where $\hat{d}$ is an estimate of the long-run parameter d from (1) whilst v_t is the disturbance. The unit-root null $H_0 : \xi = 0$ vs. $H_1 : \xi < 0$ can be tested using the t_ξ statistics derived from OLS. Lobato and Velasco (2007) have shown this test is distributed as standard normal.

(3) Given the individual t_ξ-ratios are distributed as standard normal we can straightforwardly construct a *panel version* of this test by taking the average value of the cross section units: $\overrightarrow{t_\xi} = N^{-1}\sum_{i=1}^{N} t_{\xi i}$.

3.6 Results

Tables 1-3 reports results based on panel unit-root tests with and without a trend (presented for all countries, OECD, Non-OECD). Applying the P, P_m and Z panel unit-root tests which combine the p_i-values of individual DF regressions, Table 1, we conclude that the null hypothesis of a unit root cannot be rejected. For more powerful panel unit-root tests such as the S panel unit-root test – allowing for cross-sectionally correlated errors as well as for cross cointegration (between the different country Ginis for example) – the same story holds.

The $CIPS$ test (which takes into account the multi-factor structure of the cross correlation)[10] and is more appropriate for all combinations of cross section units and time series dimension provides some evidence that the net Gini rejects a unit root. However, it should be noted that the $CIPS$ test rejects the null of a unit root at 5% for the Net index for all countries as well for the OECD when a trend is included in the fitted regression. Given that the power of the panel root tests is quite low in the presence of a linear trend

[10] All these assumptions run contrary to P, P_m and Z panel unit-root tests to tests with less size distortion.

(see Moon et al. (2007) while the $CIPS$ test is less powerful for small values of T (Pesaran et al. (2013)) it could reasonable to adopt as a working hypothesis that both inequality indices follow a unit root process.

Table 1: Panel Unit-Root Tests without Structural Breaks

	All Countries		OECD		Non-OECD	
	Gross	Net	Gross	Net	Gross	Net
			Constant, no trend†			
P	81.70	84.62	48.08	51.51	33.61	33.11
P_m	0.89	0.68	0.09	0.36	1.26	1.31
Z	0.75	1.36	0.12	0.49	1.20	1.45
S	−0.40	−0.62	−0.24	−0.41	−0.56	−0.85
$CIPS$	−2.12	−1.99	−2.13*	−2.02	−2.10*	−1.96
			Constant plus trend‡			
P	89.52	107.74	46.47	62.63	43.05	45.11
P_m	0.35	1.02	0.16	1.49	0.30	0.10
Z	0.10	0.39	0.04	0.99	0.18	0.45
S	−1.04	−1.22	−0.85	−0.87	−0.94	−0.83
$CIPS$	−2.41	−2.69**	−2.37	−2.89**	−2.45	−2.48

Notes: The null in each case is of a unit root. †The critical values for P tests at 1%, 5% and 10% statistical level are 135.87, 124.34 and 118.49 respectively [All Countries], 76.15, 67.51 and 63.17 [OECD, non-OECD, where $N_{\text{OECD}} = 24 \approx N_{\text{Non-OECD}} = 47 - 24 = 23$]. The critical values for the $CIPS$ test at 1%, 5% and 10% statistical levels are [All Countries]: -2.74, -2.61 and -2.53, [OECD, non-OECD]: -2.85, -2.68 and -2.59. ‡The critical values for the $CIPS$ test at 1%, 5% and 10% are -2.22 -2.09 -2.01 (All Countries), OECD/non-OECD: -2.34 -2.16 -2.07. The number of lags in the dynamic term was optimally selected using the BIC.

Table 2: Panel Unit-Root Tests with Structural Breaks

	All Countries		OECD		Non-OECD	
	Gross	Net	Gross	Net	Gross	Net
$DCPM$	84.04	111.76	40.44	53.20	47.99	59.00

Notes: The null in each case is of a unit root. The critical values for the $DCPM$ tests are the same as the P test.

Table 3: Panel Fractional Unit-Root Tests

	All Countries		OECD		Non-OECD	
	Gross	Net	Gross	Net	Gross	Net
			Constant, no trend			
$\overrightarrow{t_{\xi}}$	0.81	0.71	0.80	0.69	0.83	0.72
			Constant plus trend			
$\overrightarrow{t_{\xi}}$	0.63	0.52	0.60	0.50	0.66	0.54

Notes: The null in each case is of a unit root. For the constant plus trend case we implemented the test of Dolado et al. (2008) which takes into account the existence of a deterministic trend in the fitted ADF regression.

Next, to avoid erroneous inference due to possible SBs, we fitted a Fourier model (see section 3.1), Table 2. We set $\max(\mathbb{k}) = 5$ then used unit increments to search for the optimal frequency ($\mathbb{k}$) using the BIC. Then an $F-$statistic was employed to test the null hypothesis $\alpha_1^{\mathbb{k}} = \alpha_2^{\mathbb{k}} = 0$ (i.e., no structural change) against $\alpha_1^{\mathbb{k}} \neq \alpha_2^{\mathbb{k}} \neq 0$ (some structural changes).[11] Then, having identified the number of breaks for each country we estimate equation (6) for each country and combine the individual p_i-values to construct the $DCPM$ panel unit-root test. The results of this test strongly support the conclusion that both inequality indices follow a unit-root process. This is a striking conclusion: namely that even controlling for breaks, the presence of a unit root across country inequality measures continues to hold. Finally, Table 3, we present the fractional panel unit-root results. Consistent with previous findings, the fractional tests confirm a unit root in the two Gini series.

4 Conclusions

We examined the persistence of international inequality measures. Across a battery of heterogeneous panel tests, we tend to find that a unit root in income inequality cannot be rejected within our sample.[12] This is robust to the presence of SBs. This applies to both gross and net indices: thus whilst redistributive measures have reduced the level of inequality, they have not sufficiently modified its dynamic. This non-stationarity corroborates the qualitative evidence regarding the growing inequality in recent decades across many countries. However, given subject-matter reasoning and our limited sample (1975-2012), a more plausible conclusion is that inequality measures are exceptionally persistent if not strictly speaking a unit root.

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[11] For the Net index, for 38/47 countries the optimal frequency equalled 1 while for India and South Africa it was zero. These findings suggest for most countries there was only one major event (SB) that affected the DGP. For the remaining seven countries [Colombia, Denmark, Greece, Mexico, Netherlands, Spain, Malaysia] $\mathbb{k} = 2$, suggesting that more than one SB was present in the DGP. The Gross Gini shows a similar picture. For 37 countries we were able to identify only one major event ($\mathbb{k} = 1$), for 7 [Australia, Chile, China, Colombia, Japan, Korea, Panama] $\mathbb{k} = 2$, for 2 [Finland, Indonesia] $\mathbb{k} = 3$, while $\mathbb{k} = 0$ for India. (See appendix C). For critical values of this $F-$ statistic see Enders and Lee (2012)

[12] Maestri and Roventini (2012) cannot reject a unit root for several developed countries' inequality measures, albeit in a single-country (non-panel) framework with less-rigorous testing strategies.

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A Summary Statistics

Table A1: Gini Indices

Country	Variable	Obs	Mean	St. Dev	Min	Max	Country	Variable	Obs	Mean	St. Dev	Min	Max
Argentina	Gross	46	43.07	3.62	35.55	48.76	Kenya	Gross	37	57.19	5.42	48.73	66.79
	Net	46	39.85	4.09	31.42	46.54		Net	37	52.48	7.47	41.26	67.82
Australia	Gross	53	42.12	5.66	27.68	50.12	Korea	Gross	51	34.62	2.46	28.06	39.96
	Net	53	28.91	2.23	24.99	33.26		Net	51	31.69	2.21	25.55	36.57
Bangladesh	Gross	40	39.34	4.02	33.19	45.91	Malaysia	Gross	48	47.48	2.09	43.46	51.43
	Net	40	36.70	3.80	31.03	42.59		Net	48	43.69	2.11	38.39	47.20
Belgium	Gross	47	45.45	1.85	41.77	48.27	Mexico	Gross	46	48.84	2.92	43.09	56.70
	Net	47	24.96	1.31	22.18	27.87		Net	46	47.99	3.09	42.95	55.65
Brazil	Gross	35	57.23	2.22	53.56	64.46	Morocco	Gross	35	45.61	6.64	38.13	58.42
	Net	35	50.07	2.26	45.37	56.43		Net	35	41.94	4.43	36.53	52.87
Bulgaria	Gross	53	25.86	5.74	17.34	35.92	Netherlands	Gross	48	46.79	2.32	42.80	56.06
	Net	53	24.90	5.83	16.53	36.09		Net	48	25.78	1.61	23.09	31.20
Canada	Gross	52	41.10	5.04	29.89	47.82	New Zealand	Gross	53	40.25	6.50	30.34	49.46
	Net	52	28.55	2.60	20.32	31.79		Net	53	29.41	3.32	24.81	33.81
Chile	Gross	39	52.71	1.80	46.94	54.94	Norway	Gross	53	39.40	3.21	33.60	45.31
	Net	39	49.79	1.65	44.32	51.52		Net	53	23.33	1.14	20.33	25.67
China	Gross	43	38.70	8.38	27.78	51.61	Pakistan	Gross	49	35.44	1.91	31.10	41.33
	Net	43	38.34	9.33	27.32	53.56		Net	49	32.61	1.63	29.83	38.54
Colombia	Gross	49	53.85	3.69	47.87	65.05	Panama	Gross	50	51.42	2.56	46.37	56.24
	Net	49	52.39	3.89	45.92	63.51		Net	50	48.52	2.50	43.50	54.02
Denmark	Gross	51	41.80	3.55	33.42	49.17	Philippines	Gross	46	47.37	1.34	45.21	51.05
	Net	51	23.70	1.77	20.39	27.26		Net	46	44.06	1.27	41.74	47.00
Egypt	Gross	45	38.63	3.54	33.41	44.75	Portugal	Gross	40	47.52	5.73	34.94	55.61
	Net	45	36.48	3.09	31.56	41.08		Net	40	31.86	3.95	21.72	36.72
Finland	Gross	51	43.20	3.93	35.20	49.32	South Africa	Gross	43	66.29	2.16	58.54	69.36
	Net	51	23.35	2.17	19.64	26.38		Net	43	59.23	1.77	51.19	62.59
France	Gross	51	47.06	2.38	40.20	52.52	Spain	Gross	46	41.86	7.07	28.07	51.41
	Net	51	30.58	2.56	27.18	40.37		Net	46	31.04	3.43	20.39	35.28
Germany	Gross	49	46.03	3.44	37.80	51.63	Sri Lanka	Gross	38	41.24	2.97	34.60	45.94
	Net	49	27.65	1.33	24.41	30.30		Net	38	37.81	2.74	32.47	42.54
Greece	Gross	50	48.64	2.10	45.37	55.17	Sweden	Gross	53	43.04	3.55	36.10	48.88
	Net	50	34.19	1.36	31.75	38.52		Net	53	22.76	2.49	17.96	27.01
Hungary	Gross	52	44.00	6.17	33.39	53.90	Taiwan	Gross	44	29.70	2.18	26.97	33.08
	Net	52	25.73	2.99	21.02	31.88		Net	44	28.42	1.88	25.90	31.83
India	Gross	51	45.42	2.11	41.42	51.89	Thailand	Gross	37	45.44	1.83	41.75	48.30
	Net	51	45.75	2.00	42.08	51.36		Net	37	42.30	1.96	38.10	45.41
Indonesia	Gross	46	37.73	2.64	33.61	44.97	Turkey	Gross	39	46.30	3.65	40.43	53.48
	Net	46	35.40	2.38	31.44	42.07		Net	39	43.67	3.46	37.82	50.23
Ireland	Gross	50	47.50	3.38	38.84	56.43	United Kingdom	Gross	53	45.70	7.66	34.78	54.63
	Net	50	32.67	1.64	28.52	34.66		Net	53	30.58	3.88	25.44	35.81
Israel	Gross	33	48.80	3.04	42.67	53.29	United States	Gross	53	45.23	3.34	39.41	50.74
	Net	33	33.14	2.70	30.11	37.95		Net	53	33.99	2.66	29.88	37.82
Italy	Gross	47	47.76	3.36	41.59	53.77	Uruguay	Gross	37	49.65	1.72	47.55	52.99
	Net	47	33.44	2.12	29.10	37.61		Net	37	41.50	1.33	39.81	44.18
Japan	Gross	51	36.72	5.55	28.54	46.69	Venezuela	Gross	49	42.51	1.87	38.79	46.49
	Net	51	26.63	2.34	22.27	30.88		Net	49	39.68	1.92	35.98	43.49
Jordan	Gross	42	39.56	1.63	35.63	42.44							
	Net	42	37.38	1.65	33.76	40.45							

Note: Derived from Solt (2009).

B Country Results on Persistence

We have so far concentrated on the unit-root hypothesis for the *multi-country* heterogeneous panel as a whole. This, reflects, as we said in the Introduction, that the recognised trends in inequality across countries warrants such a perspective, and in contrast to the low power of individual unit-root tests, that panel unit-root tests more powerfully combine cross-section variation with time series information. Nonetheless for completeness, we can show the unit root parameters for the individual estimation of equation (1) which then make up the various fisher-transform tests. These are listed below in the tables for the constant and constant plus trend case.

Moreover, according to Stock (1991) unit-root tests and point estimates might be not a good approximation to describe the DGP as they fail to provide us with critical information about the range of values of $1 + \varphi_i$ that are consistent with the observed data. Thus confidence intervals for $1 + \varphi_i$ can shed additional light about the degree of persistence. The calculation of confidence intervals in practice is not a trivial task since the asymptotic distribution $1 + \varphi_i$ is not standard when $1 + \varphi_i \approx 1$.

To this end we follow the method suggested by Stock (1991) to construct consistent confidence intervals for $1 + \varphi_i$. The following table shows 95% confidence intervals for the $1 + \varphi_i$ based on the Dickey Fuller distribution. For both the gross and net Ginis, the confidence intervals almost always traverse unity (either directly in the central estimate or in the upper bound for $1 + \varphi_i$). Occasionally (e.g., the UK in the constant plus trend case) the parameter does not exceed unity but clearly as indicated by the upper half life (of 81 years in tat case) it is effectively a unit root process and insufficiently away from a unit-root process to affect inference for the panel as a whole.

Table B1: The estimated $\widehat{\phi}_i$ based on equation (1), with constant

	Net				Gross			
country	$1+\phi_i$	$1+\overline{\phi_i}$	$1+\underline{\phi_i}$	hl^{up}	$1+\phi_i$	$1+\overline{\phi_i}$	$1+\underline{\phi_i}$	hl^{up}
Argentina	0.955	1.025	0.884	∞	0.925	1.037	0.813	∞
Australia	0.847	1.092	0.602	∞	0.814	1.009	0.620	∞
Bangladesh	0.988	1.046	0.929	∞	0.986	1.053	0.919	∞
Belgium	0.867	1.085	0.649	∞	0.906	1.050	0.763	∞
Brazil	0.965	1.083	0.846	∞	0.965	1.077	0.853	∞
Bulgaria	0.948	1.055	0.841	∞	0.958	1.065	0.851	∞
Canada	0.975	1.105	0.844	∞	0.971	1.052	0.891	∞
China	0.994	1.051	0.938	∞	0.991	1.056	0.926	∞
Chile	0.915	1.206	0.624	∞	0.959	1.226	0.692	∞
Colombia	0.848	1.059	0.637	∞	0.829	1.058	0.600	∞
Denmark	0.750	0.943	0.557	12	0.871	1.098	0.643	∞
Egypt	0.937	1.043	0.832	∞	0.938	1.045	0.830	∞
Finland	0.980	1.131	0.829	∞	0.937	1.113	0.762	∞
France	0.877	1.085	0.668	∞	0.770	1.116	0.423	∞
Germany	0.817	1.180	0.455	∞	0.955	1.114	0.797	∞
Greece	0.772	1.008	0.537	∞	0.839	1.153	0.526	∞
Hungary	0.941	1.060	0.822	∞	0.955	1.053	0.857	∞
India	0.945	1.245	0.646	∞	0.976	1.259	0.693	∞
Indonesia	0.888	1.155	0.621	∞	0.861	1.128	0.594	∞
Ireland	1.038	1.172	0.904	∞	0.912	1.117	0.707	∞
%Israel	1.008	1.106	0.910	∞	0.898	1.042	0.754	∞
Italy	0.678	1.053	0.303	∞	0.860	1.114	0.606	∞
Japan	1.000	1.179	0.820	∞	0.998	1.126	0.871	∞
Jordan	0.828	1.076	0.580	∞	0.798	1.069	0.526	∞
Kenya	0.974	1.137	0.811	∞	0.945	1.09	0.801	∞
Korea	0.855	1.041	0.668	∞	0.898	1.085	0.712	∞
Malaysia	0.843	1.097	0.588	∞	0.930	1.166	0.694	∞
Mexico	0.851	1.040	0.662	∞	0.795	1.006	0.585	∞
Morocco	0.893	1.061	0.725	∞	0.924	1.009	0.839	∞
Netherlands	0.767	1.042	0.492	∞	0.806	1.066	0.547	∞
New Zealand	0.982	1.040	0.924	∞	0.964	1.044	0.884	∞
Norway	0.873	1.142	0.605	∞	0.975	1.111	0.839	∞
Pakistan	0.977	1.282	0.672	∞	0.953	1.210	0.696	∞
Panama	0.953	1.042	0.864	∞	0.957	1.020	0.895	∞
Philippines	0.836	1.090	0.582	∞	0.807	1.088	0.526	∞
Portugal	0.961	1.051	0.870	∞	0.895	1.037	0.752	∞
South Africa	0.480	1.017	-0.057	∞	0.828	1.097	0.560	∞
Sri Lanka	0.894	1.231	0.557	∞	0.798	1.165	0.431	∞
Spain	0.844	1.060	0.628	∞	0.929	1.072	0.785	∞
Sweden	0.863	1.119	0.606	∞	0.850	1.072	0.629	∞
Taiwan	0.986	1.067	0.905	∞	0.981	1.075	0.886	∞
Thailand	0.957	1.159	0.756	∞	0.968	1.131	0.805	∞
Turkey	0.942	1.054	0.830	∞	0.934	1.044	0.823	∞
United Kingdom	0.956	1.000	0.913	∞	0.923	0.991	0.854	81
United States	0.977	1.025	0.929	∞	0.965	1.024	0.907	∞
Uruguay	0.913	1.044	0.783	∞	0.921	1.036	0.806	∞
Venezuela	0.927	1.069	0.785	∞	0.907	1.058	0.755	∞

Table B2: The estimated $\widehat{\phi}_i$ based on equation (1), with constant plus trend

		Net				**Gross**		
country	$1+\phi_i$	$1+\overline{\phi_i}$	$1+\underline{\phi_i}$	hl^{up}	$1+\phi_i$	$1+\overline{\phi_i}$	$1+\underline{\phi_i}$	hl^{up}
Argentina	0.977	1.101	0.852	∞	0.927	1.132	0.722	∞
Australia	0.341	0.929	-0.247	∞	0.551	1.002	0.100	∞
Bangladesh	0.794	0.989	0.599	63	0.799	1.024	0.574	∞
Belgium	0.781	1.123	0.440	∞	0.897	1.075	0.719	∞
Brazil	0.918	1.073	0.762	∞	0.928	1.068	0.787	∞
Bulgaria	0.844	1.052	0.636	∞	0.837	1.047	0.627	∞
Canada	0.869	1.145	0.592	∞	0.858	1.274	0.441	∞
China	0.914	1.122	0.706	∞	0.924	1.133	0.716	∞
Chile	0.822	1.176	0.469	∞	0.847	1.183	0.510	∞
Colombia	0.805	1.074	0.537	∞	0.791	1.078	0.505	∞
Denmark	0.753	0.989	0.516	62	0.753	1.024	0.482	∞
Egypt	0.838	1.120	0.556	∞	0.813	1.091	0.535	∞
Finland	0.845	1.104	0.586	∞	0.828	1.167	0.489	∞
France	0.865	1.219	0.511	∞	0.560	1.128	-0.008	∞
Germany	0.461	0.931	-0.009	10	0.294	1.135	-0.547	∞
Greece	0.706	1.098	0.315	∞	0.903	1.268	0.538	∞
Hungary	0.802	1.145	0.460	∞	0.802	1.140	0.464	∞
India	0.749	1.258	0.241	∞	0.745	1.264	0.227	∞
Indonesia	0.480	1.071	-0.111	∞	0.457	0.978	-0.064	31
Ireland	0.696	1.108	0.284	∞	0.744	1.290	0.198	∞
Israel	0.786	1.149	0.423	∞	0.789	1.189	0.390	∞
Italy	0.627	1.086	0.169	∞	0.709	1.057	0.362	∞
Japan	0.676	1.135	0.217	∞	0.736	1.109	0.364	∞
Jordan	0.639	1.038	0.240	∞	0.704	1.132	0.275	∞
Kenya	0.923	1.154	0.692	∞	0.937	1.060	0.814	∞
Korea	1.058	1.457	0.660	∞	0.943	1.247	0.639	∞
Malaysia	0.969	1.108	0.830	∞	0.961	1.060	0.862	∞
Mexico	0.838	1.106	0.570	∞	0.810	1.065	0.555	∞
Morocco	0.841	1.161	0.521	∞	0.808	1.154	0.462	∞
Netherlands	0.620	1.026	0.215	∞	0.763	1.098	0.429	∞
New Zealand	0.944	1.104	0.784	∞	0.944	1.198	0.690	∞
Norway	0.516	1.035	-0.003	∞	0.740	1.087	0.394	∞
Pakistan	0.442	1.097	-0.213	∞	0.869	1.283	0.454	∞
Panama	0.731	1.196	0.266	∞	0.742	1.195	0.289	∞
Philippines	0.636	1.063	0.210	∞	0.602	1.069	0.136	∞
Portugal	0.888	1.054	0.722	∞	0.886	1.058	0.713	∞
South Africa	0.856	1.105	0.606	∞	0.852	1.037	0.667	∞
Sri Lanka	0.799	1.127	0.471	∞	0.716	1.066	0.366	∞
Spain	0.820	1.095	0.544	∞	0.841	1.063	0.619	∞
Sweden	0.429	0.977	-0.119	30	0.617	1.057	0.177	∞
Taiwan	0.822	1.019	0.626	∞	0.791	1.027	0.556	∞
Thailand	0.854	1.214	0.495	∞	0.842	1.148	0.537	∞
Turkey	0.774	1.117	0.432	∞	0.711	1.156	0.265	∞
United Kingdom	0.958	1.104	0.812	∞	0.922	1.086	0.757	∞
United States	0.922	1.141	0.703	∞	0.880	1.138	0.622	∞
Uruguay	0.917	1.137	0.696	∞	0.916	1.093	0.738	∞
Venezuela	0.933	1.117	0.748	∞	0.910	1.104	0.716	∞

C Structural Break Information

Table C1: Estimates of frequency $\Bbbk$ in the Fourier model, plus F-test for $\alpha_1^{\Bbbk} = \alpha_2^{\Bbbk} = 0$

	Net		Gross			Net		Gross	
Country	$\hat{\Bbbk}$	$F[\hat{\Bbbk}]$	$\hat{\Bbbk}$	$F[\hat{\Bbbk}]$	Country	$\hat{\Bbbk}$	$F[\hat{\Bbbk}]$	$\hat{\Bbbk}$	$F[\hat{\Bbbk}]$
Argentina	1	66.33	1	77.88	United Kingdom	1	33.29	1	26.20
Australia	1	8.61	2	10.96	United States	1	62.02	1	29.80
Belgium	1	53.72	1	20.51	Bulgaria	1	17.92	1	38.73
Brazil	1	29.36	1	50.57	Egypt	1	12.77	1	16.04
Canada	1	71.87	1	56.37	Hungary	1	61.31	1	94.55
Chile	1	12.88	2	17.03	Indonesia	1	20.05	3	18.15
Colombia	2	18.67	2	15.97	Jordan	1	19.70	1	11.55
Denmark	2	15.31	1	16.97	Kenya	1	10.95	1	12.52
Finland	1	160.33	3	64.72	Korea	1	24.87	2	29.51
France	1	39.08	1	8.71	Malaysia	2	16.11	1	25.04
Germany	1	28.56	1	37.58	Morocco	1	25.68	1	25.18
Greece	2	23.67	1	24.96	Pakistan	1	11.28	1	12.52
India	0	5.12	0	5.38	Panama	1	182.50	2	204.14
Israel	1	34.01	1	30.77	Philippines	1	12.65	1	12.14
Italy	1	8.47	1	43.11	South Africa	0	5.20	1	18.85
Japan	1	27.15	2	53.07	Sri Lanka	1	26.30	1	19.91
Mexico	2	17.25	1	14.52	Taiwan	1	34.39	1	48.28
Netherlands	2	7.13	1	14.42	Thailand	1	48.76	1	70.63
New Zealand	1	79.33	1	59.99	Bangladesh	1	45.56	1	37.85
Norway	1	39.48	1	55.18	China	1	48.86	2	58.01
Portugal	1	135.17	1	10.39	Ireland	1	20.27	1	20.50
Spain	2	12.17	1	18.23	Uruguay	1	21.33	1	21.74
Sweden	1	49.12	1	19.48	Venezuela	1	25.77	1	20.05
Turkey	1	10.16	1	13.69					

Note: The critical values are taken from Table 1 of Enders and Lee (2012).