Revisiting the bi-directional causality between debt and growth: Evidence from linear and nonlinear tests

Glauco De Vita [1], Emmanouil Trachanas [2], Yun Luo [3,*]

[1] *Centre for Business in Society, Coventry University, Coventry CV1 5FB, UK. E-mail: glauco.devita@coventry.ac.uk.*

[2] *Department of Accounting, Finance and Economics, Oxford Brookes University, Oxford OX3 0BP, UK. E-mail: etrachanas@brookes.ac.uk.*

[3] *Centre for Business in Society, Coventry University, Coventry CV1 5FB, UK. E-mail: yun.luo@coventry.ac.uk *Corresponding author.*

Journal of International Money and Finance, 2018
https://doi.org/10.1016/j.jimonfin.2018.02.004

ABSTRACT

We revisit the bi-directional causality between public debt and the rate of GDP growth for 10 EMU countries alongside the US, UK and Japan, over sample periods spanning from 1970 up to 2014 whilst accounting for the nonlinear properties of both the individual time series, and their relation in both directions. Our results indicate that the causal relation between debt and growth, in either direction, is weak at best. For most of the countries in our sample, we find no robust evidence of a long-run causal effect using bi-variate Granger causality tests. Bi-directional causality is detected only for Austria, while for France, Luxembourg and Portugal, causality runs solely from debt to growth, but the estimated effects are very small. In Finland, Spain and Italy, Granger causality (from growth to debt in the former two and debt to growth in Italy) appears to be present in the short-run. Our findings cannot be taken as evidence that a high level of public indebtedness is not risky for the economy or as invalidating hypotheses postulating effects in either direction in the relation between debt and growth.

JEL: C22; F34; H63; O40; O57

Keywords: Public debt; economic growth; Granger causality; nonlinear causality; SYS-GMM

1. Introduction

At a time when many EU governments have implemented harsh austerity budgets in the face of the euro area sovereign debt crisis in spite of a slow recovery in the aftermath of the 2007/08 financial crisis and high, if not growing, unemployment, the significance of the relation between public debt and economic growth has assumed particular policy relevance. It is not surprising, therefore, that the debate on the debt-growth nexus has flourished recently.

The debate has mainly revolved around whether there exists a public indebtedness threshold beyond which countries carrying high levels of debt experience significant costs in terms of GDP growth. At its core, is the seminal work by Reinhart and Rogoff (2010) suggesting that countries experience a sharp GDP growth penalty if they allow their public debt (defined as all debt that a government owes) to rise above 90% of GDP.[1] The significance of this finding cannot be overstated given the historically high levels of public debt in most advanced economies. Since its publication, a spurt of studies focusing on whether there exists, in fact, a universally valid tipping point in the debt/GDP ratio above which growth prospects are dramatically compromised, has ensued, causing much controversy (see, among others, Herndon et al., 2013; Pescatori et al., 2014; Égert, 2015).

Adjudication of the debate notwithstanding, even the absence of a specific common threshold beyond which public debt penalizes economic growth, of course, would not mean that countries do not invite trouble by running irresponsible budgets, or that debt levels and economic performance are not statistically correlated. Yet, by focusing on the corroboration or refutation of Reinhart and Rogoff's findings, much of the subsequent literature has missed the opportunity to redirect the debate towards the critical question of the direction by which the implicit causality runs, namely, does high public indebtedness cause slower GDP growth or does low GDP growth increase public debt?

Arguably, countries pursuing irresponsible budget policies end up compromising future growth. However, it is equally reasonable to expect that an economic downturn produces higher levels of public debt even under an unchanged policy regime (as a result of lower tax revenue and higher government expenditure in public benefits), and a regime switch to policies aimed at alleviating the impact of a recession is likely to lead, at least in the

[1] Reinhart and Rogoff (2010) present this result as a stylized fact, stressing the relevance of this relation to a range of times and places. They examine 44 countries covering almost two centuries of data on central government debt, inflation and growth. They organize the country-years data units in four groups by public debt/GDP ratios: 0 to 30%, 30 to 60%, 60 to 90%, and greater than 90%. They then compare average real GDP growth rates across the debt/GDP groupings. The non-parametric method used unearths a nonlinear relation, with a debt threshold effect becoming significant at around 90% of GDP.

short term, to even higher debt. Significantly, most of related literature follows the assumption that causality runs from public debt to GDP growth. The few studies considering the possibility of reverse causality can be counted on one hand (Ferreira, 2009; Di Sanzo and Bella, 2015; Gómez-Puig and Sosvilla-Rivero, 2015; and Puente-Ajovín and Sanso-Navarro, 2015). This is striking given the existence of other work positing that a country's level of public indebtedness is likely to grow when the growth rate of GDP is low (see, e.g., Bell et al., 2015).

With the sole exception of Di Sanzo and Bella (2015), the few studies investigating bi-directional causality between debt and growth have done so by applying the standard, linear Granger causality approach, thereby ignoring the possibility of a nonlinear relation. This approach is problematic not only because the conventional Granger causality test has low power to detect nonlinear causal relations, but also because many studies suggest the existence of a nonlinear relation between debt and growth, either in the form of an inverse U-shaped curve or threshold-type behavior.

Di Sanzo and Bella (2015) is the only study investigating the links between debt and growth using both linear parametric and nonlinear nonparametric Granger causality tests, finding mixed evidence. Despite their contribution in being the first to apply nonlinear causality tests to the debt-growth relation, several caveats must be noted about their study, which provide us with plenty of scope to advance on their work. Although Di Sanzo and Bella (2015) argue that overlooking nonlinearities may result in misleading conclusions about Granger causality and that linear as well as nonlinear causality tests are only valid under the hypothesis that the observed data are stationary, they test for unit roots using solely basic, linear unit root tests. Moreover, though they test for the presence of nonlinear causality using also the Diks and Panchenko (2006) test that we too employ, they do not consider the existence of nonlinearities in the individual series nor do they employ any nonlinear unit root tests to accurately establish the stationarity of the variables. We use the Harvey et al. (2008) test to verify whether there may exist any nonlinearities in the evolution of the individual time series, and for the series found to exhibit nonlinearities, we perform the nonlinear unit root test proposed by Kruse (2011). Furthermore, though Di Sanzo and Bella (2015) acknowledge that the short length of the time series they use might affect the power of the tests they perform, we verify the robustness of our results by means of SYS-GMM panel estimations based on sample sizes of up to 930 observations on an extended panel of 37 countries.

We contribute to this literature by revisiting the bi-directional causality between debt and growth for the original EMU-11 countries (with the exclusion of Germany, for which consistent data over the sample period is not available)[2] along with the US, UK and Japan, over sample periods spanning from 1970 up to 2014. The analysis accounts for the nonlinear properties of both public debt and GDP growth, and their causal relation in both directions. In addition to conventional linear Granger causality tests and the adoption of a relevant cointegration method to address endogeneity concerns,[3] we employ state-of-the-art nonlinear techniques on the individual countries' time series dimension so as to best approximate the true data generation process in each country considered and, unlike any previous study using time series estimation methods, we perform robustness causality tests within a panel SYS-GMM framework that can satisfactorily deal with potential problems stemming from serial correlation, small-sample bias, measurement error and endogeneity. We find no robust evidence of causality in most of the countries in our sample, linearly or nonlinearly.

2. A synthesis of the literature on the debt-growth nexus

2.1. The effect of debt on growth

Theoretically, both neoclassical and endogenous growth models (Modigliani, 1961; Diamond, 1965; Saint-Paul, 1992; Aizenman et al., 2007) suggest that high levels of public debt will always reduce the rate of economic growth. Additional channels in support of a negative effect of public debt on long-term growth, include: (i) the 'debt overhang' hypothesis (Krugman, 1988; Sachs, 1989); (ii) the 'liquidity constraint' hypothesis (Moss and Chiang, 2003); (iii) the 'crowding out' effect (Hansen, 2004); and (iv) the 'uncertainty' channel (Codogno et al., 2003; Cochrane, 2011). Another channel through which high debt can have a negative impact on economic growth is that of long-term interest rates (e.g., Elmendorf and Mankiw, 1999; Tanzi and Chalk, 2000). Finally, some effects associated with financial liberalization ranging from greater bank risk-taking to the accumulation of a large external debt, can make a country vulnerable to economic shocks which often have severe recessionary consequences (see Eichengreen and Leblang, 2003; Nyambuu and Bernard, 2015).

[2] The original 11 members are: Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain. Germany had to be excluded from our sample since German debt data were only available from 1991 in the IMF WEO and AMECO databases, and from 1992 in the OECD database.

[3] We also test for Granger causality within a cointegration framework using the associated error correction model (Granger, 1988).

Given the theoretical predictions highlighted above, it is somewhat surprising that Reinhart and Rogoff's conclusion of a debt threshold of 90% of GDP above which countries experience a substantial decline in their GDP growth rate, prompted such controversy. Herndon et al. (2013) provided a strong critique, pointing to coding errors and less than adequate methodological choices. In their replication, they found that over 1946-2009, countries with public debt/GDP ratios above 90% averaged 2.2% real annual GDP growth, not -0.1%, as indicated by Reinhart and Rogoff. Other critiques ensued (see, e.g., Irons and Bivens, 2010). While admitting the data coding error, Reinhart and Rogoff showed in subsequent papers (see Reinhart et al., 2012) that the growth rate in regimes of high debt, while positive, was still much lower than in regimes of low debt. Using data comprising up to 17 OECD advanced economies, Panizza and Presbitero (2014) employ a new instrument for public debt (a variable capturing valuation effects brought about by the interaction between foreign currency debt and exchange rate volatility) and, after controlling for endogeneity, are unable to reject the null hypothesis that debt has no impact on growth. Pescatori et al. (2014) found no evidence of any particular debt threshold above which growth prospects are seriously compromised and showed that countries with high but declining debt appear to grow equally as fast as countries with lower debt.

While this literature generally concludes that there is no universally valid tipping point after which the ratio of debt to GDP will necessarily lead to slower economic growth, there are also several studies which support the essence of Reinhart and Rogoff's findings (e.g., Cecchetti et al., 2011; Baum et al., 2013; Checherita-Westphal and Rother, 2012). Several other studies, albeit failing to find evidence for a universally valid debt threshold, do find evidence that there is a significant negative effect of public debt build-up on output growth (e.g., Chudik et al., 2017). Eberhardt and Presbitero (2015) also find some evidence for a negative relation between public debt and growth, but not for a common debt threshold.

Whilst the literature is rather consensual in the view that high public debt poses significant economic challenges and makes the economy less resilient to shocks, some channels postulating a positive debt-growth relationship have also been hypothesized. For example, although Abbas and Christiansen (2010) find that above a ratio of 35% of bank deposits, domestic debt begins to undermine growth, they also observe that moderate levels of noninflationary domestic debt have a positive overall impact on economic growth via a range of channels, including improved monetary policy, strengthened institutions/accountability, and enhanced private savings and financial intermediation. DeLong and Summers (2012) demonstrate that, under certain conditions, expansionary fiscal policies that create public debt accumulation but avoid a protracted recession, result in a positive effect on both short- and long-term growth. For a comprehensive survey of what theory tells us about the relation between public debt and growth, see Panizza and Presbitero (2013).

Earlier empirical analyses had focused on the existence of an inverse U-shaped curve in the debt-growth relation. Patillo et al. (2002) found that at low levels of total external debt, in developing countries, the impact on the rate of GDP growth is positive, but this relation becomes negative at high levels of debt. Schclarek (2004) investigated the debt-growth relation for a number of developing (59) and industrial (24) economies. In contrast to Patillo et al. (2002), he found that, for developing countries, lower total external debt levels are associated with higher growth rates. He did not find any support for an inverted-U shape relation. For industrial countries, he did not find any significant relation. Presbitero (2005) found no evidence of an inverse U-shaped curve. His results suggest that indebtedness generally reduces income growth, and that this effect is much larger in low income and highly indebted poor countries. Puente-Ajovín and Sanso-Navarro (2015) test for Granger causality between debt (government as well as non-financial corporate and household debt) and growth in 16 OECD countries over 1980-2009. Their results barely provide evidence against the null hypothesis that government debt does not cause real GDP growth while finding evidence against the absence of causality from non-financial private debt to growth.

2.2. Bi-directional causality

The theoretical literature also postulates that causality may run in the other direction, from growth to debt. It has been hypothesized that low growth causes higher levels of debt (Reinhart et al., 2012; Bell et al., 2015) because tax revenues go down and spending goes up in a recession, thus increasing debt levels to maintain the welfare state. Yet, relatively few empirical studies have investigated bi-directional causality, and their findings are mixed.

Using OECD data for 20 countries over 1988-2001, Ferreira (2009) concludes in favor of Granger causality between the growth of real GDP per capita and public debt, and that this causality is always bi-directional. Lof and Malinen (2014) analyze the dynamic relation between sovereign debt and economic growth for 20 developed countries over 1954-2008. They find no robust evidence for an effect of debt on growth but do find that growth has a statistically significant impact on debt.

Gómez-Puig and Sosvilla-Rivero (2015) investigate bi-directional causality before and after endogenously detected breaks for 11 EMU countries over 1980-2013. They follow

Ferreira (2009) in pre-testing for unit roots using fairly basic, linear tests and then use the conventional Granger causality test but allowing for endogenously determined structural breaks, thus permitting the causality tests to be performed before and after such breakpoints. They find no evidence of a causal relationship from growth to debt for four out of the eleven countries considered (Austria, Greece, Italy and Portugal) and mixed evidence in the remaining cases. Their evidence on the relation from public debt to growth unveils a negative relationship in five of the eleven countries after the detected break dates (which in most cases take place between 2007 and 2009), and that the 'debt threshold' above which the relationship becomes negative differs by country, with the estimated tipping point ranging from 53% in the case of Spain to 103% for Italy. However, akin to most papers in this literature, this study neglects the possibility of nonlinearities, which cannot be captured, if present, simply by accounting for breaks.

Di Sanzo and Bella (2015) investigate the links between debt and growth for 12 euro countries over 1970-2012 using both linear parametric and nonlinear nonparametric causality tests. They find a unidirectional causality from debt to growth for Spain and Portugal and bi-directional causality for Belgium, Germany, Greece, Ireland and Italy. No causality in either direction is identified for Austria, Finland, Luxembourg and the Netherlands. Finally, for France, their tests suggest a unidirectional causality from GDP growth to debt-to-GDP ratio though this evidence is based on their shortest series (from 1977 to 2012).

3. Data and empirical strategy

3.1. Data sample, sources and measures

To ensure the quality and long time span of the time series required, country data were collected from a range of databases, and alternative measures of the debt and economic growth variables used for robustness tests. For the main analysis, the two measures are the general government consolidated gross debt (excessive deficit procedure and former definitions, based on the European System of Accounts 2010, linked series) as a percentage of GDP at current prices (D), and real GDP growth (Y). The growth data are collected from the OECD Economic Outlook database for all countries except Ireland. Given that for Ireland ready-made GDP growth data from OECD were only available from 1991, we chose to take the Irish GDP growth ready-made series from the IMF World Economic Outlook database (1980-2013). For all countries, the debt-to-GDP ratios used are from AMECO, the annual macroeconomic database of the European Commission's Directorate General for Economic and Financial Affairs. Following Eberhardt and Presbitero (2015), we too concentrate on government debt thereby choosing to ignore the complex ways in which private and public debt may interact across countries, which is beyond the scope of our study.

We consider both central and peripheral countries of the European Economic and Monetary Union (Austria, Belgium, Finland, France, Ireland, Italy, Luxembourg, Netherlands, Portugal, and Spain)[4], to which we add, for comparative purposes and comprehensiveness, the US, UK and Japan. The samples are based on annual data from 1970 to 2014 for Belgium, Finland, Italy, Luxembourg, Spain, UK, USA and Japan, 1971 to 2014 for Austria, 1977 to 2014 for France, 1975 to 2014 for the Netherlands, 1974 to 2014 for Portugal, and 1980 to 2013 for Ireland.[5] Descriptive statistics are in Appendix A (all the data and codes used are available on request).

3.2. Unit root and linearity tests

With few exceptions, prior empirical studies ignore investigating the times series properties of the individual public debt and GDP growth series, a choice seemingly based on the unwarranted assumption that both variables are linear and level stationary. Gómez-Puig and Sosvilla-Rivero (2015) and Di Sanzo and Bella (2015) do check the presence of unit roots but they do so using solely fairly basic linear unit root tests. The former uses the Augmented Dickey-Fuller (ADF) and the Kwiatkowski et al. (1992) unit root tests, the latter uses the ADF test and the Zivot and Andrews' (1992) test that allows for an endogenous break under the null. We employ the Lee and Strazicich (2003) unit root test, which is considerably more advanced as it allows for two breaks in level and trend under both the null and the alternative hypothesis. Next, we use the Harvey et al. (2008) test to establish whether any nonlinearities may exist in the evolution of the individual time series (for details, see De Vita and Trachanas, 2016). For the individual series found to be nonlinear, we then perform

[4] Following the distinction made by the European Commission (1995) between those countries whose currencies continuously participated in the ERM of the EMS from its inception and those which joined later or suspended their participation, we take the central countries to be Austria, Belgium, Finland, Luxembourg and the Netherlands, and the peripheral countries to be Ireland, Italy, Portugal, and Spain.

[5] Our GDP growth data for Ireland end at 2013 so as not to include the controversial 26% spike recorded in 2015, which in July 2016 prompted Krugman's sarcastic tweet about 'Leprechaun economics'. The Irish inflated economic growth figures for 2015 were in fact due to the concomitance of one-off factors including activity in the airline leasing sector and restructuring by multinationals involving the movement of patents.

the nonlinear unit root test proposed by Kruse (2011), which is a powerful extension of the well-known Kapetanios et al. (2003) test.[6]

3.3. Granger causality tests and ARDL cointegration

Once the most likely nature (linear or nonlinear) and order of integration of the *Y* and *D* series for each country is established, we apply standard linear Granger causality tests for the countries where both the *Y* and *D* series are level stationary.[7] For countries where some uncertainty may remain as to the order of integration of the underlying series, we proceed to test for a long-run relationship between debt and growth as well as causality using the ARDL bounds testing approach (Pesaran and Shin, 1999; Pesaran et al., 2001), a methodology which allows for the analysis of cointegration when it is not known with certainty whether the regressors are $I(1)$ or $I(0)$.[8]

Although the Granger representation theorem states that if two time series are cointegrated then there must be Granger causality between them (at least in one direction), the reverse is not necessarily true. This means that any finding of Granger causality in the absence of cointegration should be taken to be either spurious, or to reflect - at best - a short-term phenomenon that does not persist in the long-run. As Eberhardt notes (2016, p. 3), in the context of the long-run growth-debt relation "the presence of a long-run equilibrium is a pre-requisite for the existence of any long-run causal relationship in the data". Another advantage that makes the ARDL approach particularly suited to our analysis lies in the rich set of dynamics of the ARDL specification, which allows the ARDL-based estimator to satisfactorily address potential endogeneity problems (Pesaran and Shin, 1999).

To investigate the direction of causality, we test for ARDL cointegration in both directions. The ARDL(p,q) models estimated for each country are:

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[6] The Kruse (2011) modified Wald statistic $\tau = t^2_{\delta_2^{\perp}=0} + 1\left(\hat{\delta}_1 < 0\right)t^2_{\delta_1=0}$ is used to test the null of a unit root $H_0:\ \delta_1 = \delta_2 = 0$ against the alternative of a stationary ESTAR process $H_1:\ \delta_1 < 0,\ \delta_2 \neq 0$.

[7] Evidently, even for two $I(0)$ variables, OLS regressions are susceptible to endogeneity bias. Given that for the regressions $Y = \beta_0 + \beta_1 D + \nu$ and $D = a_0 + a_1 Y + \varepsilon$, the OLS estimator of β_1 is given by $\left[\left(\beta_1\sigma_\varepsilon^2\right) + \left(\alpha_1\sigma_\nu^2\right)\big/\left(\sigma_\nu^2 + \alpha_1^2\sigma_\nu^2\right)\right]$ and the bias of β_1 can be expressed as $E\left(\beta_1\right) - \beta_1 = \left[a_1\left(1 - \beta_1 a_1\right)/\sigma_\varepsilon^2/\left(\sigma_\nu^2 + a_1^2\right)\right]$, OLS estimations are biased if, and only if, a_1 is negative. If $a_1 = 0$, then debt is not endogenous and OLS estimations are unbiased.

[8] Since there is always a degree of uncertainty as to the accuracy of unit root tests, it is prudent to apply this methodology to verify the existence of any cointegrating properties.

$$\Delta D_t = \mu + \rho D_{t-1} + \theta Y_{t-1} + \sum_{j=1}^{p-1} \alpha_j \Delta D_{t-j} + \sum_{j=0}^{q-1} \pi_j \Delta Y_{t-j} + e_t \tag{1}$$

$$\Delta Y_t = \mu + \rho Y_{t-1} + \theta D_{t-1} + \sum_{j=1}^{p-1} \alpha_j \Delta Y_{t-j} + \sum_{j=0}^{q-1} \pi_j \Delta D_{t-j} + \varepsilon_t \tag{2}$$

where e_t and ε_t are *i.i.d.* stochastic processes. The existence of a stable long-run relationship is tested by means of two statistics: the modified *F*-test (F_{PSS}) advanced by Pesaran et al. (2001), which tests the joint null hypothesis of no cointegration $\rho = \theta = 0$; and a *t*-test (t_{BDM}) proposed by Banerjee et al. (1998), which tests the null of no cointegration $\rho = 0$ against $\rho < 0$.[9]

3.4. Nonlinear causality test

Finally, we perform the Diks and Panchenko (2006) nonlinear causality test.[10] Its application on the delinearized series ensures that any causality identified is solely nonlinear in nature. We apply the test by utilizing properly specified VAR models or, in the case of confirmed robust cointegration in both directions, the residuals of the ARDL models, thus accounting for the actual integration properties of the individual series and also the cointegration findings, a process that ensures the stationarity of the residuals.

4. Empirical analysis and results

We begin our analysis by inspecting the plots of the individual series for each country in our sample, and the respective scatter diagrams of the relationship between growth rates (*Y*) and debt-to-GDP ratios (*D*). Fig. 1 illustrates the plots of the *Y* and *D* series. The former series show a consistent pattern across countries, generally displaying a mean reverting tendency. The same cannot be said for the evolution of the *D* series, which exhibit a mixture of upward and downward sloping trends in addition to the presence of possible structural

[9] There are two critical bounds: upper and lower. If the values of the F_{PSS} and t_{BDM} statistics exceed the upper bound, the null hypothesis is rejected. If they lie below the lower critical bound, the null cannot be rejected, and if they lie between the critical bounds, the test is inconclusive.

[10] Diks and Panchenko (2006) introduced a new and more powerful nonparametric test for Granger non causality which avoids the severe over-rejection of the null hypothesis frequently observed in the use of the Hiemstra and Jones' (1994) test. They showed that the reason for over-rejection in the Hiemstra-Jones' test is that the test statistic ignores possible variations in the conditional distribution that may occur under the null hypothesis. As the size of the test approaches unity, the test statistic almost always rejects Granger non-causality, when in fact no such causality exists. The Diks and Panchenko's test offers a solution to the distortions of the actual size of the Hiemstra-Jones' test and was therefore preferred in our application despite the relatively short time series for some countries in the sample.

breaks and/or nonlinearities. That said, although visual inspection of the plots of time series can provide valuable information about their evolution over time, determination of the actual order of integration of each time series is a task best left to unit root tests. The scatter plots of the relationship between *Y* and *D* presented in Fig. 2 are, for most countries, suggestive of a possible negative correlation between the variables though there is no clear cut pattern of such a relationship for some of the countries in question.

[Fig. 1 and 2 here]

Panel A of Table 1 presents the results of the Lee and Strazicich (2003) unit root test for the GDP growth rates series, which are found to be level stationary for all countries. Panels B and C of Table 1 refer to the debt-to-GDP ratio, in levels and, where needed, in first and second differences. The results suggest that the debt-to-GDP ratio of Finland, France, Ireland, Italy, Portugal, Spain and the UK are level stationary, while that of Austria, Luxembourg, the Netherlands, Japan and the USA are first difference stationary. The results for the debt-to-GDP series of Belgium suggest that it is integrated of order two and thus, should this result be confirmed from nonlinear tests, this country would need to be excluded from the cointegration and causality analyses that follow.

[Table 1 here]

Table 2 reports the results of the W_{λ} linearity test statistic of Harvey et al. (2008). Regarding the growth series, the W_{λ} statistic rejects the null of linearity at the customary significance levels for Finland, France (7%), Italy, Netherlands, Spain and the UK. The test results further suggest that the debt-to-GDP series of Austria, Belgium, France, Ireland, Portugal and Japan are also nonlinear.

[Tables 2 and 3 here]

Armed with the above findings, we perform nonlinear unit root tests for the variables that exhibit nonlinearities. Table 3 presents the results of the Kruse (2011) test applied to the raw, demeaned and detrended growth (*Y*) and debt (*D*) series. The findings suggest that the growth series of Finland, France, Italy, the Netherlands, Spain and the UK are level stationary (Panel A). In addition, the debt-to-GDP ratio of Ireland is level stationary while that of Austria, France, Portugal and Japan would appear to be first difference stationary. For the debt-to-GDP ratios of France and Portugal, given the nonlinearity of the series, we follow the indication of the presence of a unit root detected by the Kruse (2011) test. The results for the debt-to-GDP ratio of Belgium confirm that this series is integrated of order two and thus this country is excluded from the analyses that follow.

Having established the most likely nature (linear or nonlinear) and order of integration of the Y and D series for each country, we apply standard linear Granger causality tests for the five countries where both the growth and debt series are level stationary, i.e. Finland, Ireland, Italy, Spain and the UK. The results are presented in Table 4.[11] They suggest that growth Granger-causes debt in Finland and Spain, while debt Granger-causes growth in Italy. Aiming to establish the existence of a possible long-run relationship between growth and debt for these countries, we additionally employed OLS. The estimated parameters and diagnostics are presented in Table 5. The long-run coefficients, for both the growth to debt (β_Y) and the debt to growth (β_D) regressions, are statistically significant only for Ireland and Italy, indicating a negative relationship, with much larger magnitudes for the β_Y coefficients (-2.96 and -6.31, respectively). However, these models fail a mixture of diagnostic tests at customary significance levels (including serial correlation) leading us to conclude that the findings are spurious. Failure to unveil robust evidence in support of a long-run relation between the variables for Finland, Ireland, Italy, Spain and the UK also signifies that the evidence of Granger causality running from growth to debt for Finland and Spain, and debt to growth for Italy reported in Table 4, is – at best – likely to apply solely to the short-term.

[Tables 4 and 5 here]

For the remaining countries, i.e. Austria, France, Luxembourg, the Netherlands, Portugal, Japan, and the USA, we have established that their growth rates are level stationary and their debt-to-GDP ratios are, most likely, first difference stationary. Thus, in order to investigate the existence of a possible long-run relationship, we proceed by using the ARDL bounds testing approach (Pesaran and Shin, 1999; Pesaran et al., 2001). Table 6, presents the ARDL bounds test results, the estimated long-run coefficients (where applicable), and some diagnostics. From the F_{PSS} and t_{BDM} statistics, at the 1% and 5% significance levels, we find evidence in support of cointegration for Austria (in both directions), France (from debt to growth), Luxembourg (from debt to growth), the Netherlands (from debt to growth), Portugal (from debt to growth), Japan (from growth to debt and debt to growth), and the USA (from debt to growth). For Austria, France, Luxembourg, Portugal and Japan, the estimated long-run coefficients (β_Y and/or β_D) are statistically significant and negative. On the other hand, for the Netherlands and the USA the long-run coefficients are not significant, suggesting the

[11] Since Granger causality tests can be sensitive to lag length, the optimal VAR lag length was chosen using the Schwarz criterion, starting with max four lags. This starting lag choice was driven by the inherent trade off between a risk of bias stemming from the smaller variance at lower lag lengths and the loss of efficiency accruing at higher lag lengths.

absence of a long-run relationship between the variables, and hence any long-run path to causality. The diagnostics of the models that exhibit cointegration for Austria, France, Luxembourg, Portugal and Japan (from growth to debt model only) are quite satisfactory. However, for Japan (debt to growth model) the diagnostics reveal serial correlation. As shown in Fig. 3, according to the cumulative sum of recursive residuals (CUSUM) and the CUSUM sum of squares (CUSUMSQ) tests to assess parameter constancy, all the models appear to be stable with the exception of the Netherlands (hence corroborating the finding of an insignificant long-run coefficient) and Japan (for the growth to debt model), which fail to pass the more powerful CUSUMSQ test.

Overall, and taking into account the inferences about the statistical significance of the long-run coefficients and the conclusions to be drawn from our battery of diagnostics tests, we detect evidence of robust, linear level relationships, and thus long-run causality, only for Austria, in both directions, and for France, Luxembourg and Portugal in the direction running from debt to growth.

[Table 6 and Fig. 3 here]

The results of the Diks and Panchenko (2006) nonlinear causality test are in Table 7. For Finland, Ireland, Italy, Spain and the UK, where both growth and debt are $I(0)$, we apply the test on the residuals of a VAR model in levels (see Table 7, Panel A). For France, Luxembourg, the Netherlands, Portugal, Japan and the USA, where the growth series are $I(0)$ and the debt series are $I(1)$, and a long-run level relation has been detected in one direction only, we apply the test on the residuals of a VAR model with the growth series in levels and the debt series in first differences (Table 7, Panel B). For Austria, where ARDL cointegration has been detected in both directions, we apply the test on the residuals of the respective ARDL model (Table 7, Panel C). The last two columns of Table 7 also presents the Ljung–Box Q-test (applied to the residuals from the VAR/ARDL models) in order to determine whether any linear dependency remains in the residuals. With the exception of Luxembourg (growth causes debt case), the results of this test suggest that the VAR/ARDL models successfully account for linear dependency.

[Table 7 here]

According to the results presented in Table 7, at the 1% and 5% levels, significant nonlinear causal effects are only detected for Finland (growth causes debt) only when two and three lags are considered, for Italy (growth causes debt) only when one lag is considered, and for Spain, from growth to debt and debt to growth, only when one and two lags are considered, respectively. Thus, with the exception of very sporadic and short-lived episodes, there is no evidence of a consistent and systematic pattern of nonlinear causality present in the relationship between public debt and economic growth across the countries in our sample.

In conclusion, we unveil evidence of (linear) bi-directional Granger causality only for Austria, with a much larger negative coefficient in the regression running from growth to debt (-20.62) while for France, Luxembourg and Portugal, linear causality runs solely in the direction of debt to growth, but the estimated long-run effects are very small, with magnitudes ranging from -0.01 for France to -0.03 for Luxemburg and Portugal. Less conclusive evidence emerges with respect to Finland, Italy and Spain, where Granger causality (from growth to debt for Finland and Spain, and debt to growth for Italy) would only appear to be present, at best, in the short-run. For all the other countries in our sample there is no reliable evidence of causality in either direction. Significantly, we also find no robust or consistent evidence in support of the existence of nonlinear causality between debt and growth across the countries in our sample.

As highlighted in our literature review, previous studies present very mixed results. These include the absence of any evidence that debt causes growth (as concluded by Panizza and Presbitero, 2014), the presence of a higher number of statistically significant relationships running from growth to debt (see, e.g., Puente-Ajovín and Sanso-Navarro, 2015), and "clear evidence" of causality and that "it is always bi-directional" (Ferreira, 2009, p. 12). Our results sit comfortably between this vastly conflicting range of prior findings, though they lean towards the rejection of the hypothesis of any causal relationship between the variables given that our battery of linear and nonlinear tests only uncover robust evidence of long-run causality in four out of the twelve countries examined (Austria, in both directions, and France, Luxembourg and Portugal from debt to growth). Di Sanzo and Bella (2015) find greater evidence in favor of causality than we do. However, lack of investigation of the cointegrating properties of the debt and growth series, and their failure to properly establish the linearity or otherwise of the individual series and to carry out nonlinear unit root tests to accurately identify the actual order of integration of the debt and growth series, raises doubts as to the reliability of their Diks and Panchenko (2006) nonlinear causality test results since, as they themselves acknowledge, the validity of the linear and nonlinear Granger causality tests they perform hinges on "the hypothesis that the observed data are stationary" (p. 640).[12]

[12] Unlike Di Sanzo and Bella (2015), as per our discussion in section 3.4, we ensured the proper application of the Diks and Panchenko (2006) test on *stationary* series by utilizing properly specified VAR models or, in the case of confirmed robust cointegration in both directions, the residuals of the ARDL models. For an analogous implementation of this test to ensure accurate results, see De Vita and Trachanas (2016).

Lack of a uniform result across all countries in our sample may be due to the fact that the effect of debt on growth depends either on its composition (Dell'Erba et al., 2013) or the specific production technologies in a given country (Temple, 1999). Any divergence with previous work also lends itself to a straightforward rationalization. For example, whilst Ferreira's (2009) sample period is rather short (14 years) and dated (ending at 2001), we cover a time frame spanning from 1970 to 2014. Moreover, Ferreira (2009) used real GDP per capita to measure economic growth, and the values she obtains for the overall "R-squared" and "R-squared between" measures of her panel estimations, reveal considerable differences in the behavior of the 20 OECD countries that make up the sample she used, which suggests that such countries constitute anything but a homogenous set.

Our results align more closely to those by Pescatori et al. (2014) who, using a novel empirical approach and an extensive dataset, conclude that "the relationship between debt and growth is relatively muted and the magnitude is much smaller than the dramatic figures suggested in earlier studies." (p. 15). However, by their own admission, while their methodology may attenuate problems of reverse causality, their analysis is still subject to potential endogeneity concerns.

5. Robustness tests

Our analysis would be incomplete without offering some reassurances as to the reliability of our time series test results. With this aim in mind, the first question we address is, how sensitive are our results to the methodologies employed? In particular, having established the behavior of the individual countries in our sample, especially in light of the relatively short time series particularly for some countries, we seek to verify the extent to which a reliable panel estimation method would generate results that broadly corroborate our time series findings. To this end, we employ a powerful panel methodology particularly suited to deal with econometric issues arising from estimation of the relationship at hand, namely SYS-GMM (Arellano and Bover, 1995; Blundell and Bond, 1998). In addition to accounting for the underlying dynamics and individual country-specific effects,[13] SYS-GMM corrects for potential problems stemming from the correlation between the regressors and the error term, measurement error and endogeneity. The SYS-GMM approach also allows for a panel data test of Granger causality. For instance, taking the regression

[13] The first-difference transformation embedded in the standard GMM estimator, the first step of SYS-GMM, effectively eliminates the heterogeneity of individual country-specific effects included in the baseline equation.

$Y_{it} = a_0 + a_1 Y_{t-1} + a_2 Y_{t-2} + \beta_1 D_{t-1} + \beta_2 D_{t-2} + u_{it}$, a Wald test distributed as a χ^2 with two degrees of freedom for the joint null $\beta_1 = \beta_2 = 0$, can be employed to test for the absence of Granger causality running from debt to growth (for a similar application, see Luo et al., 2016).

Another issue that warrants further investigation is the potential time-varying nature of the debt-growth relation (as recently highlighted by Gómez-Puig and Sosvilla-Rivero, 2017). In particular, we are interested in testing for the possibility that the generally higher debt levels or harsh austerity programs implemented in advanced economies in the aftermath of the 2007/08 financial crisis may have themselves affected the debt-growth relation. For example, Gómez-Puig and Sosvilla-Rivero (2015) do not find causality up to the breaks detected between 2007 and 2009, but for the sub-periods from the detected breaks to the end of their sample period (at 2013) they do. Within our individual-country time series estimations based on annual data, we were unable to run separate regressions for shorter sub-periods (too few observations post 2009), but we can now test this hypothesis using SYS-GMM, which is particularly well suited to address the potential 'small T' sample bias stemming from regressions with few time periods.

We then wish to establish how measure-dependent our results are by testing their sensitivity to alternative debt and growth measures. Since 'gross debt' data in particular may not be immune to measurement issues (Dippelsman et al., 2012), we start by using the alternative gross government debt-to-GDP ratios downloaded from http://www.carmenreinhart.com/data/browse-by-topic/topis/9/, which are the data updates discussed in Reinhart and Rogoff (2011), also used by Chudik et al. (2017). Further, we re-estimate using two alternative data series for growth. The first is 'real GDP per capita' from the Maddison database of the Groningen Growth and Development Centre (http://www.ggdc.net/maddison/maddison-project/home.htm), also used by Lof and Malinen (2014). The second is a real GDP growth measure we computed ourselves by taking the real GDP series available from AMECO, the same database from which we sourced the original debt-to-GDP series.

Finally, given that the gains from SYS-GMM versus the traditional GMM estimator are more pronounced when the size of T (time periods) is small relative to N (country units), we check for any possible small sample bias by expanding our cross-sectional units of the panel. The extended sample includes the EU-28 (while still excluding Germany) alongside other OECD countries, namely the US, Japan, Canada, Switzerland, Norway, Iceland, Mexico, Korea, Australia and New Zealand.

[Tables 8 and 9 here]

Tables 8 and 9 report the SYS-GMM estimations in the direction of growth to debt and debt to growth, respectively. In both tables, column (1) reports the estimations for the full sample (1980-2013), based on a balanced panel that draws from the same dataset as our time series estimations. Columns (2) and (3) refer to the sub-samples 1980-2006 and 2009-2014, respectively (to ensure an end date at 2014, the latter sub-sample is based on an unbalanced panel). Column (4) refers to estimations employing the real GDP growth measure we constructed with data sourced from AMECO, column (5) uses the alternative debt measure discussed above, and column (6) additionally uses the alternative growth data series from the Maddison database. Since the alternative measure of debt is drawn from a dataset whose coverage ends at 2010, the balanced panel in columns (5) and (6) refers to the period 1980-2010. Also, since this dataset does not provide information for Luxemburg, only eleven country units are used in estimation. Column (7) re-estimates the specification and measures used in column (1) but for a much larger sample of (37) countries, as detailed above.

As shown in Table 8 and 9, in all cases we are unable to reject the null of no Granger causality, in either direction, results which – time series exceptions notwithstanding - broadly support the general conclusions drawn from our country-level estimations. For all permutations, both the Hansen (1982) test for the validity of the over-identifying instrument restrictions and the Arellano-Bond (1991) AR(2) serial correlation test provide reassuring diagnostics, thus confirming that the proposed specifications are adequate for valid inference.[14]

6. Conclusions

Following the global financial crisis, an assumed yet substantially untested assumption concerning the relationship between public debt and economic growth has sat at the heart of justifications of national austerity responses within and beyond Europe. This paper interrogates this relationship, moving beyond the usual ‘threshold analysis’ to test for the bi-directional causality between debt and growth for 10 EMU countries along with the US,

[14] Although SYS-GMM deals with time invariant country-specific effects, it assumes the absence of cross-sectional dependence, which can arise due to spatial effects or unobserved common factors. Hence, we also re-estimated our baseline panel model of column 1 in Table (8) and (9), using the Dynamic Common Correlated Effects Estimator (DCCE) proposed by Chudik and Pesaran (2015); a method that is robust to different types of cross-sectional dependence. In the DCCE specification, the first and second lag of growth rate (*Y*) and debt-to-GDP ratio (*D*) were instrumented by their fourth and fifth lag, with three lags used for the cross section averages. Also the results of these estimations (available on request) produced results consistent with statistically insignificant effects for both the growth to debt and the debt to growth regressions.

UK and Japan, over the period 1970-2014, using state-of-the-art linear and nonlinear techniques. Our analysis addresses the still unsettled question of the effect of public debt accumulation on economic growth and whether some reverse causality may be at work, with low GDP growth causing increases in public indebtedness.

Our results indicate that the causal relationship between debt and growth, in either direction, is weak at best. In 8 out of the 12 countries in our sample (Belgium had to be excluded from the estimations due to the higher order of integration detected for the debt-to-GDP series), our data are unable to uncover robust evidence of long-run causality from either debt to growth or growth to debt. Only for Austria bi-directional causality is found (with a much stronger effect from growth to debt) while for France, Luxembourg and Portugal, causality runs solely in the direction of debt to growth, but with estimated long-run coefficients of very small magnitudes ranging from -0.01 for France, to -0.03 for Luxemburg and Portugal. Our results survive a wide range of robustness checks.

Our findings of effects varying according to country suggest that different policy recommendations would be called for depending on the country in question, whether any evidence of causality is in fact present and, if so, its direction and associated sign and magnitude of the long-run coefficient. The mixed findings at country level most likely reflect the possibility that the many positive and negative effects postulated theoretically with respect to the influence of debt on growth as well as hypotheses advanced in relation to the impact of growth on debt levels, assume relative significance and possibly cancel each other out depending on country-specific characteristics. The existence or otherwise of postulated effects may also be contingent on possible differences characterizing the structure and composition of the debt across countries (which one cannot control for when taking gross debt levels as aggregate quantities as a proportion of GDP), specific production technologies in each country, and the way in which different policy-driven regimes in different countries may themselves affect the response of one variable to changes in the other.

Our findings, therefore, cannot be taken as definitive ones and do not, by themselves, invalidate hypotheses postulating effects in either direction in the relationship between debt and growth. Not finding Granger causality in the data does not necessarily rule out the existence of an economic relation between the variables. Indeed, several caveats must be acknowledged when interpreting our results. First, unlike structural models, our Granger-type causality analysis focuses exclusively on a simple bivariate relationship and hence fails to account for the impact of other macroeconomic and institutional variables that can influence debt levels and economic growth and the causality patterns between them. Second, failure to reject the null hypothesis of ‘no causality’ at aggregate level does not mean that the two variables may not be related or that there may not be countries in which, at very high levels of debt (above 80% of GDP), such indebtedness does affect growth, as found in several studies that employ a structural nonlinear approach and specifically test for a threshold level effect (see, among others, Kumar and Woo, 2010; Baum et al., 2013). After all, many countries in our sample have, over much of the period considered in our estimations, recorded levels of indebtedness below 80% of GDP, where the correlation with economic growth is supposedly constant according to nonlinear structural models which assume a threshold effect near 80%-100%.

Appendix A. Descriptive statistics

Table A.1

Descriptive statistics.

Panel A: Growth rates

	Mean	Maximum	Minimum	Std. Dev.	Obs.
YAUS	2.39	6.21	-3.66	1.89	44
YBEL	2.25	6.19	-2.62	1.91	45
YFIN	2.55	7.60	-8.27	3.22	45
YFRA	1.95	4.62	-2.86	1.47	38
YIRE	4.12	11.18	-5.64	3.84	34
YITA	1.81	6.55	-5.50	2.37	45
YLUX	3.64	9.98	-6.57	3.44	45
YNET	2.05	5.12	-3.29	1.89	40
YPOR	2.08	7.49	-4.35	2.97	41
YSPA	2.61	8.15	-3.57	2.41	45
YJPN	2.75	10.28	-5.53	2.94	45
YUK	2.25	6.54	-4.31	2.20	45
YUSA	2.79	7.26	-2.78	2.09	45

Panel B: Debt to GDP ratios

	Mean	Maximum	Minimum	Std. Dev.	Obs.
DAUS	54.37	84.25	16.66	19.78	44
DBEL	98.67	134.44	54.43	25.17	45
DFIN	29.61	59.33	6.08	18.19	45
DFRA	51.33	95.35	20.12	22.29	38
DIRE	72.05	120.10	23.63	30.62	34
DITA	87.98	132.53	35.72	27.35	45
DLUX	11.21	23.33	4.21	5.43	45
DNET	58.29	75.46	38.40	12.35	40
DPOR	58.30	130.17	13.30	27.30	41
DSPA	41.94	99.29	11.50	22.61	45
DJPN	106.55	246.17	11.46	71.70	45
DUK	52.55	88.17	31.30	15.21	45
DUSA	60.97	104.79	40.24	18.07	45

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Table 1
Lee and Strazicich (2003) unit root tests.

Panel A: Growth rates levels

	Model A		Model C	
Variables	LM test statistic	Break dates (level)	LM test statistic	Break dates (level; slope)
YAUS	-6.832*** (0)	1980^{b}, 1987	-7.220*** (1)	1978^{a}, 1984^{c}; 1978^{a}, 1984^{a}
YBEL	-6.260*** (0)	1986, 1996^{c}	-7.587*** (0)	1984, 1990; 1984^{a}, 1990
YFIN	-5.062*** (1)	1981, 1993	-5.956*** (1)	1988^{b}, 1994^{b}; 1988, 1994^{a}
YFRA	-4.538** (0)	1988, 1997^{c}	-4.970 (1)	1986, 1993^{b}; 1986^{b}, 1993
YIRE	-3.206 (0)	2002, 2007^{a}	-7.509*** (3)	1997^{a}, 2010^{a}; 1997^{a}, 2010^{a}
YITA	-7.344*** (0)	1988, 2007	-7.333*** (0)	1979, 2007; 1979^{b}, 2007
YLUX	-6.120*** (0)	1983, 2002	-7.064*** (0)	1988^{c}, 1997; 1988, 1997^{b}
YNET	-5.006*** (1)	2000^{c}, 2007	-5.773** (1)	2000, 2006^{c}; 2000^{c}, 2006^{b}
YPOR	-4.499*** (3)	2002, 2009^{a}	-7.021*** (3)	2002, 2009^{a}; 2002, 2009^{a}
YSPA	-5.292*** (4)	1983, 2008^{a}	-5.630** (4)	1984, 2006^{c}; 1984^{b}, 2006^{a}
YJPN	-5.456*** (0)	1984, 1999	-6.458*** (0)	1982, 1991; 1982^{a}, 1991
YUK	-5.888*** (1)	1990^{b}, 2008^{a}	-5.872*** (1)	1984, 2006; 1984, 2006^{c}
YUSA	-4.575*** (0)	1989, 2000^{c}	-5.854*** (1)	1981^{a}, 2009^{a}; 1981^{a}, 2009^{a}

Panel B: Debt to GDP ratios levels

	Model A		Model C	
Variables	LM test statistic	Break dates (level)	LM test statistic	Break dates (level; slope)
DAUS	-2.815 (1)	1996^{b}, 1999	-5.244* (1)	1984, 2003; 1984^{b}, 2003^{c}
DBEL	-3.385 (3)	1991, 1997	-4.758 (3)	1988, 2001; 1988, 2001^{b}
DFIN	-3.874** (1)	1995, 1999	-5.701** (3)	1990, 2002^{b}; 1990^{a}, 2002^{a}
DFRA	-3.661* (1)	2003, 2009^{a}	-5.254*** (2)	1993^{b}, 2007; 1993^{a}, 2007^{a}
DIRE	-3.156 (2)	1990^{b}, 2010	-6.917*** (2)	1987, 2001; 1987^{a}, 2001
DITA	-3.225 (3)	1997, 2009	-5.562** (3)	1990, 2001; 1990^{a}, 2001^{a}
DLUX	-0.734 (2)	1993, 2003	-4.465 (3)	1995, 2006^{a}; 1995^{a}, 2006^{a}
DNET	-3.019 (3)	2005^{a}, 2010^{c}	-4.273 (4)	1995, 2006^{a}; 1995^{a}, 2006^{a}
DPOR	-2.004 (1)	1987, 1991^{c}	-6.877*** (3)	1985, 2003; 1985^{a}, 2003^{a}
DSPA	-4.794*** (3)	1998, 2009	-7.508*** (3)	1988, 2002; 1988^{a}, 2002^{a}
DJPN	-2.314 (3)	2003^{b}, 2010^{a}	-4.521 (3)	1988, 2003; 1988, 2003
DUK	-3.081 (1)	1993^{a}, 2004	-7.391*** (2)	1993^{a}, 2006^{a}; 1993^{a}, 2006^{a}
DUSA	-3.330 (1)	1994, 2004	-4.532 (1)	1987, 2003; 1987^{b}, 2003^{b}

Panel C: Debt to GDP ratios first and second differences

	Model A		Model C	
Variables	LM test statistic	Break dates (level)	LM test statistic	Break dates (level; slope)
Δ*DAUS*	-5.424*** (1)	1987, 1996^{a}	-6.122*** (1)	1996^{a}, 2011; 1996^{b}, 2011^{c}
Δ*DBEL*	-3.581 (0)	1987^{c}, 1993^{b}	-4.916 (0)	1979^{c}, 1997; 1979, 1997;
ΔΔ*DBEL*	7.821*** (0)	1980, 1996	-8.168*** (0)	1992^{a}, 2008; 1992^{a}, 2008^{a}
Δ*DLUX*	-4.202** (0)	1994^{c}, 2009^{b}	-9.044*** (0)	1990, 2006^{a}; 1990, 2006^{a}
Δ*DNET*	-4.381** (0)	1993, 2006	-5.523** (0)	1983, 2000; 1983, 2000^{b}
Δ*DJPN*	-4.565*** (0)	1983, 1991	-4.820 (0)	1988, 1996; 1988, 1996
Δ*DUSA*	-4.119** (1)	1981^{a}, 2003^{b}	-5.119* (2)	1983, 1999^{a}; 1983^{b}, 1999^{a}

Notes: Model A allows for a change in the level of the series; Model C allows for changes in the level and slope of the trend of the series. The optimal lag structure is chosen following a *general-to-specific* approach, as suggested by Lee and Strazicich (2004), starting with max 4 lags, and is displayed in parentheses. The critical values are from Lee and Strazicich (2003, p. 1084). ***, ** and * denote rejection of the null of a unit root at the 1, 5 and 10% significance level, respectively. ^{a}, ^{b} and ^{c} denote significance of the break dates at the 1, 5 and 10% significance level, respectively.

Table 2
Harvey et al. (2008) linearity tests.

Growth rates	W_λ	Debt to GDP ratios	W_λ
YAUS	3.14	*DAUS*	6.05**
YBEL	2.83	*DBEL*	6.58**
YFIN	8.85**	*DFIN*	0.39
YFRA	5.25*	*DFRA*	8.25**
YIRE	3.12	*DIRE*	12.04***
YITA	15.36***	*DITA*	2.66
YLUX	2.42	*DLUX*	0.06
YNET	7.22**	*DNET*	3.22
YPOR	1.30	*DPOR*	9.47***
YSPA	8.87**	*DSPA*	3.30
YJPN	1.57	*DJPN*	13.14***
YUK	8.66**	*DUK*	1.83
YUSA	3.69	*DUSA*	0.56

Notes: The W_λ statistic follows the χ_2^2 distribution and the relevant critical values are 9.21 (1%), 5.99 (5%) and 4.60 (10%). ***, ** and * denote the rejection of the null of linearity at the 1, 5 and 10% significance level, respectively.

Table 3
Kruse (2011) nonlinear unit root test.

Panel A: Growth rates levels

Variable	Level series	Demeaned series	Detrended series	*k*
YFIN	26.75***	26.19***	21.66***	0
YFRA	10.28**	17.69***	24.09***	0
YITA	21.95***	22.97***	46.27***	0
YNET	9.53**	13.07**	15.88**	0
YSPA	3.46	8.87*	13.62**	0
YUK	36.15***	36.63***	38.88***	1

Panel B: Debt to GDP ratios levels

Variable	Level series	Demeaned series	Detrended series	*k*
DAUS	8.00*	4.81	9.38	2
DBEL	5.06	2.53	5.66	1
DFRA	4.96	1.39	11.12*	1
DIRE	9.48*	12.91**	9.14	1
DPOR	3.85	2.82	2.71	1
DJPN	9.24*	1.82	1.97	1

Panel C: Debt to GDP ratios first and second differences

Variable	Level series	Demeaned series	Detrended series	*k*
$\Delta DAUS$	11.68**	17.09***	15.80**	1
$\Delta DBEL$	4.72	5.83	8.33	0
$\Delta\Delta DBEL$	18.54***	18.60***	18.37***	0
$\Delta DFRA$	15.42***	17.42***	18.87***	0
$\Delta DPOR$	4.85	14.64***	12.20*	0
$\Delta DJPN$	16.22***	22.52***	26.03***	0

Critical values

1%	13.15	13.75	17.10
5%	9.53	10.17	12.82
10%	7.85	8.60	11.10

Notes: *k* denotes the optimal lag length chosen on the basis of the Schwarz information criterion. The critical values are from Kruse (2011, p. 77). The estimation and tests were conducted using the program code written in 'R' that was produced and provided by Kruse. ***, ** and * denote the rejection of the null of a unit root at the 1, 5 and 10% significance level, respectively.

Table 4
Linear Granger causality tests.

	F-statistic	k
YFIN ≠> DFIN	9.969*** [0.000]	2
DFIN ≠> YFIN	0.458 [0.635]	2
YIRE ≠> DIRE	1.248 [0.303]	2
DIRE ≠> YIRE	1.846 [0.177]	2
YITA ≠> DITA	0.845 [0.437]	2
DITA ≠> YITA	5.111*** [0.010]	2
YSPA ≠> DSPA	28.743*** [0.000]	1
DSPA ≠> YSPA	0.377 [0.542]	1
YUK ≠> DUK	0.867 [0.428]	2
DUK ≠> YUK	0.185 [0.831]	2

Notes: The symbol ≠> means that A does not Granger cause B. k denotes the optimal VAR lag length, chosen on the basis of the Schwarz information criterion starting with max 4 lags. p-values are reported in square brackets. ***, ** and * denote rejection of the null at the 1, 5 and 10% significance level, respectively.

Table 5
OLS estimates.

	β_D	β_Y	*SC*	*FF*	*NOR*	*HET*
YFIN → DFIN	-	-1.03 [0.312]	230.583*** [0.000]	0.187 [0.667]	4.568 [0.101]	0.037 [0.847]
DFIN → YFIN	-0.03 [0.297]	-	5.173*** [0.009]	0.220 [0.641]	16.861*** [0.000]	0.174 [0.678]
YIRE → DIRE	-	-2.96*** [0.001]	109.022*** [0.000]	1.231 [0.275]	3.100 [0.212]	1.849 [0.183]
DIRE → YIRE	-0.04** [0.041]	-	13.880*** [0.000]	0.452 [0.505]	2.361 [0.307]	0.965 [0.333]
YITA → DITA	-	-6.31*** [0.000]	28.126*** [0.000]	3.020* [0.089]	2.899 [0.234]	1.186 [0.282]
DITA → YITA	-0.04*** [0.000]	-	0.527 [0.594]	0.563 [0.457]	11.357*** [0.003]	0.243 [0.624]
YSPA → DSPA	-	-3.21 [0.491]	161.269*** [0.000]	0.003 [0.956]	0.528 [0.767]	3.577* [0.065]
DSPA → YSPA	-0.03 [0.279]	-	19.479*** [0.000]	0.877 [0.354]	0.848 [0.654]	0.456 [0.502]
YUK → DUK	-	-0.88 [0.222]	143.122*** [0.000]	0.016 [0.899]	5.002* [0.081]	0.298 [0.587]
DUK → YUK	-0.01 [0.274]	-	5.109*** [0.010]	0.029 [0.863]	7.170** [0.027]	0.240 [0.623]

Notes: The symbol → indicates the causal direction tested for. β denotes the coefficient estimated from an OLS regression. The Newey-West heteroscedasticity and autocorrelation consistent standard errors are used. *SC*, *FF*, *NOR* and *HET* denote tests for serial correlation, functional form, normality and homoskedasticity, respectively. p-values are reported in square brackets. ***, ** and * denote rejection of the null hypothesis at the 1, 5 and 10% significance level, respectively.

Table 6
ARDL cointegration tests and long-run coefficients.

ARDL model	Specification	F_{PSS}	t_{BDM}	β_D	β_Y	*SC*	*FF*	*NOR*	*HET*
YAUS $\rightarrow$ *DAUS*	(3, 2)	10.61 [a]	-3.75 [b]	-	-20.82*** [0.000]	0.236 [0.791]	0.013 [0.908]	0.066 [0.967]	1.288 [0.288]
DAUS $\rightarrow$ *YAUS*	(2, 1)	27.95 [a]	-7.45 [a]	-0.03*** [0.000]	-	1.109 [0.341]	0.923 [0.343]	1.638 [0.440]	1.310 [0.283]
YFRA $\rightarrow$ *DFRA*	(2, 0)	1.02	0.12	-	-	0.089 [0.915]	0.803 [0.377]	1.230 [0.540]	0.173 [0.913]
DFRA $\rightarrow$ *YFRA*	(2, 1)	30.22 [a]	-7.64 [a]	-0.01*** [0.000]	-	1.165 [0.326]	0.337 [0.565]	1.321 [0.516]	2.367* [0.074]
YLUX $\rightarrow$ *DLUX*	(3, 0)	1.59	-1.70	-	-	1.234 [0.303]	1.071 [0.307]	204.491*** [0.000]	0.254 [0.905]
DLUX $\rightarrow$ *YLUX*	(1, 4)	21.54 [a]	-6.53 [a]	-0.03*** [0.002]	-	0.888 [0.421]	1.628 [0.210]	0.910 [0.634]	1.946 [0.101]
YNET $\rightarrow$ *DNET*	(4, 0)	3.83	-2.58	-	-	0.521 [0.599]	2.836 [0.102]	129.373*** [0.000]	1.078 [0.392]
DNET $\rightarrow$ *YNET*	(2, 3)	6.75 [b]	-3.64 [b]	0.01 [0.598]	-	0.089 [0.914]	0.850 [0.364]	1.869 [0.392]	0.810 [0.570]
YPOR $\rightarrow$ *DPOR*	(2, 0)	1.66	-0.21	-	-	0.198 [0.820]	0.107 [0.744]	0.384 [0.825]	1.469 [0.239]
DPOR $\rightarrow$ *YPOR*	(4, 1)	12.86 [a]	-4.81 [a]	-0.03** [0.026]	-	0.440 [0.648]	1.627 [0.212]	1.553 [0.459]	1.459 [0.225]
YJPN $\rightarrow$ *DJPN*	(1, 4)	44.76 [a]	-4.81 [a]	-	-64.09*** [0.000]	1.946 [0.159]	3.413* [0.073]	1.500 [0.472]	0.689 [0.659]
DJPN $\rightarrow$ *YJPN*	(4, 1)	33.76 [a]	-7.88 [a]	-0.01*** [0.000]	-	4.635** [0.017]	0.050 [0.823]	0.492 [0.781]	0.477 [0.820]
YUSA $\rightarrow$ *DUSA*	(2, 0)	1.44	-0.69	-	-	0.369 [0.693]	0.932 [0.340]	9.012** [0.011]	0.554 [0.648]
DUSA $\rightarrow$ *YUSA*	(1, 2)	17.66 [a]	-5.70 [a]	-0.03 [0.131]	-	0.262 [0.770]	0.735 [0.396]	1.844 [0.397]	2.755** [0.041]

Notes: The symbol $\rightarrow$ indicates the causal direction tested for. The choice of the optimal ARDL specifications is based on the SBC, starting with max q = max p = 4. At the 1% (5%) significance level, the pair of critical values (bounds) for the F_{PSS} and t_{BDM} statistics are 7.87 to 8.96 (5.29 to 6.17) and -3.43 to -3.82 (-2.86 to -3.22), respectively. The critical values for the F_{PSS} and t_{BDM} statistics are from Narayan (2005, p. 1988) and Pesaran et al. (2001, p. 303). β denotes the long-run coefficient. *SC*, *FF*, *NOR* and *HET* denote tests for serial correlation, functional form, normality and homoskedasticity, respectively. p-values are reported in square brackets. [a] and [b] denote rejection of the null hypothesis of 'no cointegration' at the 1 and 5% significance level, respectively. ***, ** and * denote statistical significance at the 1, 5 and 10% level, respectively.

Table 7
Diks and Panchenko (2006) nonlinear Granger causality tests.

Panel A: with VAR filtered residuals for the I(0) pairs - both growth and debt series are I(0)

$Lx = Ly$	1	2	3	4	5	6	7	$LB(1)$	$LB(2)$
YFIN ≠> DFIN	1.255 [0.104]	1.684** [0.046]	1.650** [0.049]	1.447* [0.073]	0.898 [0.184]	0.987 [0.161]	1.07 [0.140]	0.108 [0.742]	1.979 [0.372]
DFIN ≠> YFIN	0.480 [0.315]	-0.956 [0.830]	-0.552 [0.709]	-0.475 [0.682]	-0.784 [0.783]	-1.196 [0.884]	-0.941 [0.826]	0.239 [0.624]	1.312 [0.519]
YIRE ≠> DIRE	-1.112 [0.866]	-1.198 [0.884]	-0.330 [0.629]	-0.770 [0.779]	-0.928 [0.823]	-0.996 [0.840]	-0.777 [0.781]	0.004 [0.946]	0.015 [0.992]
DIRE ≠> YIRE	-0.573 [0.716]	0.033 [0.486]	-0.322 [0.626]	0.000 [0.500]	0.000 [0.500]	-0.314 [0.623]	0.000 [0.500]	0.007 [0.932]	0.129 [0.937]
YITA ≠> DITA	1.801** [0.035]	0.994 [0.159]	1.008 [0.156]	0.940 [0.173]	0.105 [0.457]	0.100 [0.459]	0.748 [0.227]	0.000 [0.980]	0.310 [0.856]
DITA ≠> YITA	0.740 [0.229]	0.445 [0.327]	0.331 [0.370]	1.389* [0.082]	1.149 [0.125]	0.075 [0.469]	0.709 [0.238]	0.010 [0.917]	0.066 [0.967]
YSPA ≠> DSPA	2.494*** [0.006]	1.562* [0.059]	0.585 [0.279]	-0.110 [0.544]	0.178 [0.429]	-0.163 [0.565]	0.000 [0.500]	3.227* [0.072]	3.433 [0.180]
DSPA ≠> YSPA	1.288* [0.098]	1.925** [0.027]	0.873 [0.191]	1.026 [0.152]	0.769 [0.220]	0.736 [0.230]	1.259 [0.103]	0.258 [0.611]	0.917 [0.632]
YUK ≠> DUK	0.696 [0.242]	0.576 [0.282]	-0.227 [0.589]	0.096 [0.461]	0.178 [0.429]	-0.508 [0.694]	-0.644 [0.740]	0.274 [0.600]	0.534 [0.765]
DUK ≠> YUK	-0.418 [0.662]	-0.720 [0.764]	-0.183 [0.572]	-0.096 [0.538]	-0.733 [0.768]	-0.944 [0.827]	-0.951 [0.829]	0.035 [0.851]	0.183 [0.912]

Panel B: with VAR filtered residuals for the mixed order pairs - growth series are I(0) and debt series are I(1)

$Lx = Ly$	1	2	3	4	5	6	7	$LB(1)$	$LB(2)$
YFRA ≠> DFRA v	0.437 [0.330]	-0.146 [0.558]	-0.671 [0.748]	-0.536 [0.704]	0.188 [0.425]	0.513 [0.303]	0.058 [0.476]	0.059 [0.808]	0.112 [0.945]
DFRA ≠> YFRA v	-0.554 [0.710]	0.371 [0.355]	1.007 [0.156]	0.063 [0.474]	0.681 [0.247]	0.571 [0.283]	0.764 [0.222]	0.000 [0.976]	0.516 [0.772]
YLUX ≠> DLUX v	0.874 [0.191]	-0.216 [0.585]	-0.782 [0.783]	-1.192 [0.883]	-1.275 [0.898]	-1.262 [0.896]	-0.755 [0.775]	0.047 [0.828]	6.194** [0.045]
DLUX ≠> YLUX v	-0.252 [0.599]	0.837 [0.201]	1.054 [0.145]	0.554 [0.289]	0.426 [0.334]	-0.073 [0.529]	-0.721 [0.764]	0.147 [0.701]	0.233 [0.890]
YNET ≠> DNET	-0.229 [0.590]	-0.454 [0.675]	-1.485 [0.931]	-0.712 [0.761]	-0.826 [0.795]	-1.097 [0.863]	0.648 [0.258]	0.011 [0.914]	0.490 [0.783]
DNET ≠> YNET	0.286 [0.387]	-0.561 [0.712]	-0.315 [0.623]	1.057 [0.145]	-0.339 [0.632]	-1.223 [0.889]	-0.729 [0.767]	0.071 [0.789]	0.357 [0.836]
YPOR ≠> DPOR	-1.074 [0.858]	-0.671 [0.749]	-1.491 [0.932]	-0.493 [0.689]	0.164 [0.434]	-0.491 [0.688]	-0.567 [0.714]	0.054 [0.816]	0.055 [0.973]
DPOR ≠> YPOR	0.156 [0.437]	-0.323 [0.626]	-0.275 [0.608]	0.766 [0.221]	1.093 [0.136]	1.238 [0.107]	1.252 [0.105]	1.883 [0.170]	1.912 [0.384]
YJPN ≠> DJPN	0.381 [0.351]	-1.058 [0.855]	-0.762 [0.777]	-0.452 [0.674]	0.068 [0.472]	0.704 [0.240]	0.058 [0.476]	0.488 [0.485]	0.494 [0.781]
DJPN ≠> YJPN	0.638 [0.261]	0.218 [0.413]	0.205 [0.418]	0.398 [0.345]	0.457 [0.323]	0.593 [0.276]	1.087 [0.138]	0.295 [0.587]	0.326 [0.850]
YUSA ≠> DUSA	-0.424 [0.664]	0.020 [0.491]	-0.919 [0.821]	-0.771 [0.779]	-0.748 [0.772]	-0.629 [0.735]	0.152 [0.439]	0.414 [0.520]	0.888 [0.641]
DUSA ≠> YUSA	-0.416 [0.661]	-0.842 [0.800]	-0.482 [0.685]	-0.525 [0.700]	-0.995 [0.840]	-0.287 [0.613]	-0.166 [0.566]	0.065 [0.798]	0.459 [0.795]

Panel C: with ARDL filtered residuals for the mixed order pairs - growth series are I(0) and debt series are I(1) and cointegrated (both directions)

$Lx = Ly$	1	2	3	4	5	6	7	$LB(1)$	$LB(2)$

YAUS ≠> DAUS	-0.974 [0.835]	-1.867 [0.969]	-1.343 [0.910]	-0.631 [0.736]	0.175 [0.430]	-0.343 [0.634]	-0.850 [0.802]	0.065 [0.797]	0.226 [0.893]
DAUS ≠> YAUS	-0.832 [0.797]	-0.746 [0.772]	-0.800 [0.788]	0.359 [0.359]	0.439 [0.330]	0.598 [0.274]	-0.102 [0.540]	0.279 [0.597]	0.785 [0.675]

Notes: The symbol ≠> means that A does not Granger cause B. Parameter *C* for the bandwidth is 8, the theoretical optimal rate β is 2/7, and the optimal bandwidth ε_n is 1.5. The optimal VAR lag length has been chosen following the Schwarz information criterion, starting with max four lags. The estimation and tests were conducted using a program code written in C language provided by Cees Diks. *LB*(1) and *LB*(2) are the Ljung-Box statistics based on the VAR and ARDL residual series of the dependent variable up to the 1st and 2nd lag. ^{v} denotes cases in which the SIC suggested VAR structure ('constant only model') was overridden with a VAR(1) to allow estimation. *p*-values are reported in square brackets. ***, ** and * denote rejection of the null hypothesis at the 1, 5 and 10% significance level, respectively.

Table 8

SYS-GMM Granger causality test: growth Granger causes debt.

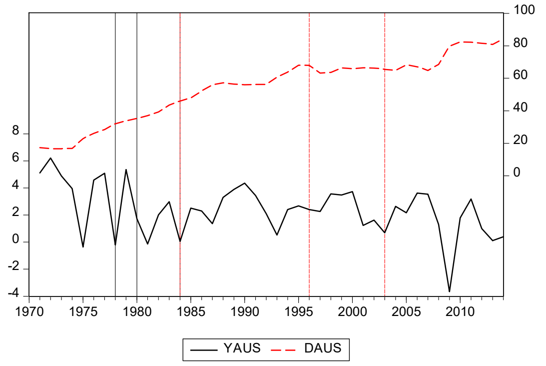
Dependent Variable: D	(1) 1980-2013	(2) 1980-2006	(3) 2009-2014	(4) 1980-2013	(5) 1980-2010	(6) 1980-2010	(7) 1980-2014
D_{t-1}	1.391***	2.306***	1.621***	1.493***	1.162***	1.020***	1.508***
	[0.342]	[0.291]	[0.123]	[0.279]	[0.192]	[0.296]	[0.148]
D_{t-2}	-0.388	-1.366***	-0.721***	-0.497*	-0.132	0.030	-0.548***
	[0.363]	[0.305]	[0.104]	[0.284]	[0.219]	[0.332]	[0.151]
Y_{t-1}	0.083	0.047	0.773*	0.098	0.727	0.655	-0.427
	[0.434]	[0.809]	[0.454]	[0.359]	[1.047]	[0.733]	[0.370]
Y_{t-2}	-0.443	0.879	-0.105	-0.321	-1.823	-1.954	-0.055
	[0.513]	[0.571]	[0.227]	[0.464]	[1.561]	[1.347]	[0.130]
$\Sigma(Y)$	-0.360	0.926	0.668	-0.214	-1.096	-1.299	-0.482
Wald test p-value > χ^2	0.604	0.170	0.233	0.774	0.456	0.349	0.311
AR(2)	0.317	-0.554	0.531	0.321	0.188	0.182	0.306
Hansen p-value	0.146	0.126	0.479	0.260	0.293	0.197	0.208
Constant	1.856	0.090	9.953	1.705	3.211*	2.296*	3.627**
	[1.366]	[2.993]	[7.129]	[1.207]	[1.914]	[1.208]	[1.824]
Number of instruments	11	11	11	11	11	11	11
Observations	384	300	47	384	319	319	930
Number of countries	12	12	12	12	11	11	37

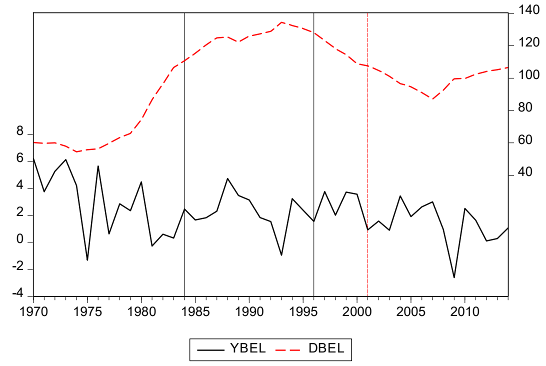
Notes: Instruments for the different equations include lags four to seven of the growth rate (Y) and debt-to-GDP ratio (D) in columns (1), (3) ,(4) and (7), lags three to six (Y and D) in column (2), and lags five to eight (Y and D) in columns (5) and (6), all used as GMM-style instruments. Columns (1), (2), (4), (5) and (6) are balanced data while (3) and (7) are unbalanced data. In column (4), the independent variable is the GDP growth rate measure based on the authors' calculations using the AMECO real GDP series (code: OVGD). In column (5) and (6), the dependent variable is gross Government debt-to-GDP ratio, using the series downloaded from http://www.carmenreinhart.com/data/browse-by-topic/topis/9/. In column (6), the independent variable is real GDP per capita growth, based on the series downloaded from http://www.ggdc.net/maddison/maddison-project/data.htm. In column (7), we extend the country units to the EU-28 (with the exclusion of Germany, for which consistent data over the sample period is not available) plus the US, Japan, Canada, Switzerland, Norway, Iceland, Mexico, Korea, Australia and New Zealand. Real GDP growth (Y) data for Bulgaria, Croatia, Cyprus, Lithuania, Malta, Romania and Slovakia, and debt (D) data for Switzerland Mexico, Korea, Australia and New Zealand, are collected from the IMF World Economic Outlook database. The 'collapse option' of xtabond2 was chosen to limit instrument proliferation. Windmeijer-corrected standard errors are reported in square brackets. ***, ** and * denote statistical significance at the 1, 5 and 10% level, respectively.

Table 9
SYS-GMM Granger causality test: debt Granger causes growth.

Dependent Variable: Y	(1) 1980-2013	(2) 1980-2006	(3) 2009-2014	(4) 1980-2013	(5) 1980-2010	(6) 1980-2010	(7) 1980-2014
Y_{t-1}	0.795**	0.646***	-0.186	0.586	0.207	0.266***	0.238
	[0.311]	[0.240]	[0.160]	[0.380]	[0.255]	[0.081]	[0.570]
Y_{t-2}	0.183	-0.124	-0.307***	0.246	-0.064	-0.031	0.791
	[0.394]	[0.277]	[0.080]	[0.477]	[1.027]	[0.239]	[0.510]
D_{t-1}	0.118	-0.075	-0.082	0.074	0.065	0.131*	0.268
	[0.252]	[0.088]	[0.191]	[0.217]	[0.148]	[0.078]	[0.205]
D_{t-2}	-0.132	0.085	0.089	-0.086	-0.073	-0.139*	-0.263
	[0.270]	[0.091]	[0.190]	[0.227]	[0.166]	[0.082]	[0.219]
$\Sigma(D)$	-0.014	0.01	0.007	-0.012	-0.008	-0.008	0.005
Wald test p-value > χ^2	0.757	0.287	0.847	0.510	0.908	0.214	0.121
AR(2)	0.424	0.635	0.388	0.479	0.754	0.167	0.139
Hansen p-value	0.177	0.459	0.243	0.154	0.144	0.815	0.429
Constant	0.568	0.896*	0.450	0.924	2.545*	1.646***	-0.859
	[0.860]	[0.519]	[2.179]	[1.042]	[1.514]	[0.594]	[1.291]
Number of instruments	11	11	11	11	11	11	11
Observations	384	300	47	384	319	319	930
Number of countries	12	12	12	12	11	11	37

Notes: Instruments for the different equations include lags four to seven of the growth rate (*Y*) and debt-to-GDP ratio (*D*) in columns (1), (2), (4) and (6), lags three to six (*Y* and *D*) in column (3), and lags five to eight (*Y* and *D*) in columns (5) and (7), all used as GMM-style instruments. Columns (1), (2), (4), (5) and (6) are balanced data while (3) and (7) are unbalanced data**.** In column (4), the dependent variable is the GDP growth rate measure based on the authors' calculations using the AMECO real GDP series (code: OVGD). In columns (5) and (6), the independent variable is gross Government debt-to-GDP ratio, using the series downloaded from http://www.carmenreinhart.com/data/browse-by-topic/topis/9/. In column (6), the dependent variable is real GDP per capita growth, based on the series downloaded from http://www.ggdc.net/maddison/maddison-project/data.htm. In column (7), we extend the country units to the EU-28 (with the exclusion of Germany, for which consistent data over the sample period is not available) plus the US, Japan, Canada, Switzerland, Norway, Iceland, Mexico, Korea, Australia and New Zealand. Real GDP growth (*Y*) data for Bulgaria, Croatia, Cyprus, Lithuania, Malta, Romania and Slovakia, and debt (*D*) data for Switzerland Mexico, Korea, Australia and New Zealand, are collected from IMF World Economic Outlook database. The 'collapse option' of xtabond2 was chosen to limit instrument proliferation. Windmeijer-corrected standard errors are reported in square brackets. ***, ** and * denote statistical significance at the 1, 5 and 10% level, respectively.


YAUS DAUS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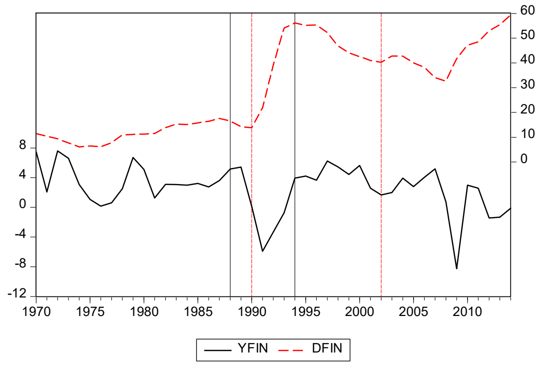
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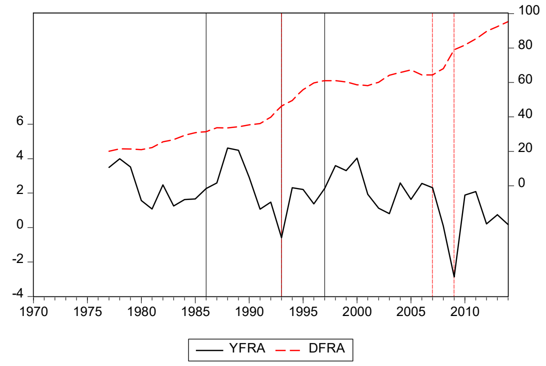
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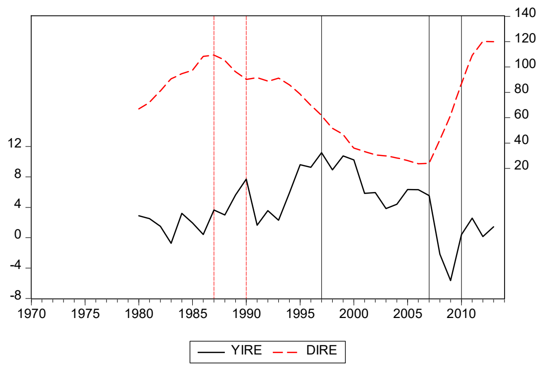
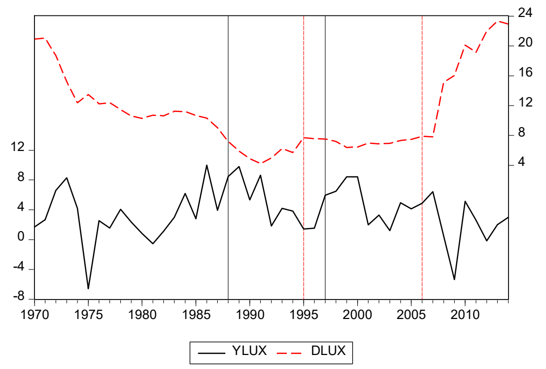
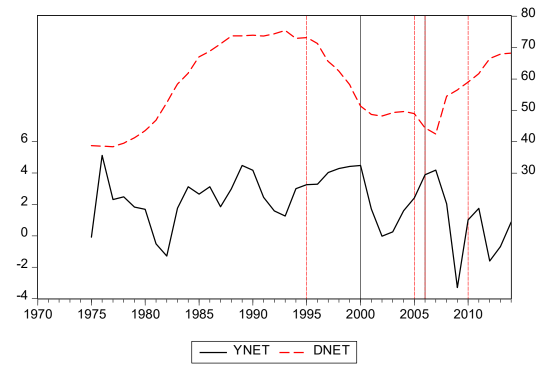

YIRE DIRE


YITA DITA


YLUX DLUX


YNET DNET

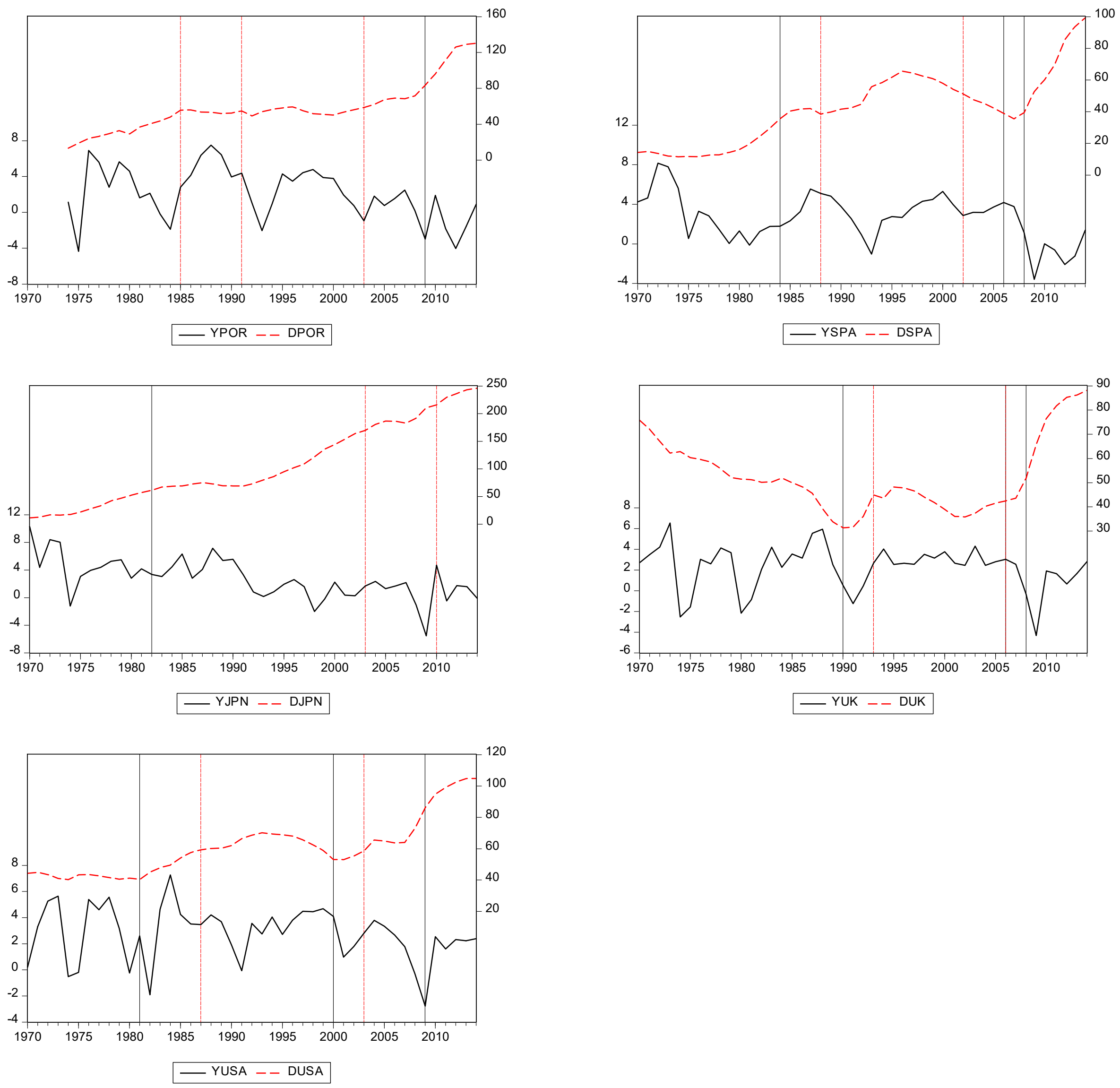
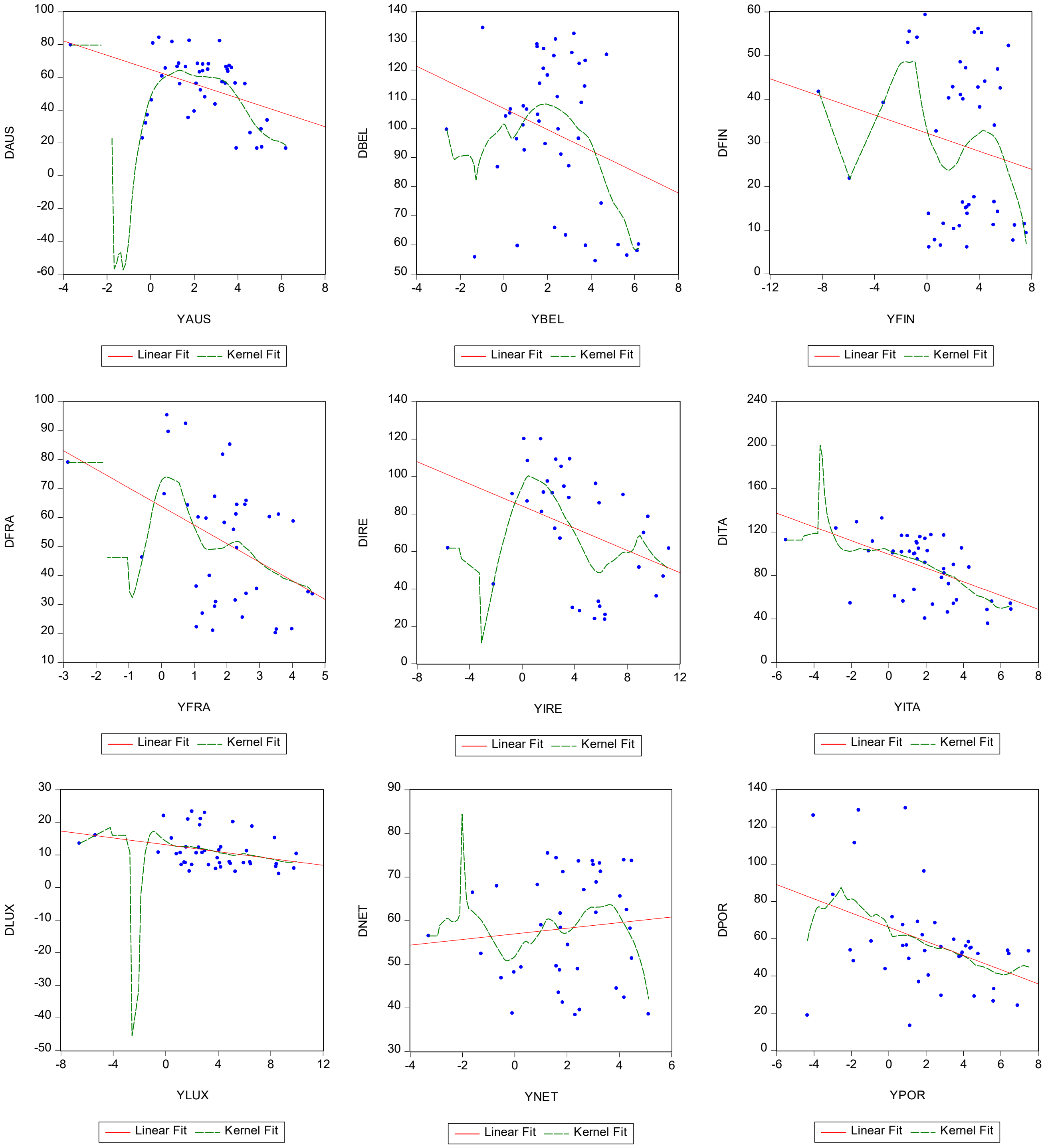


Fig. 1. The evolution of the growth rates (*Y*) and the debt-to-GDP ratios (*D*). The vertical lines correspond to statistically significant structural breaks reported in Table 1 (continuous line for *Y* and dotted line for *D* – some overlap also due to same break date for *Y* and *D*)


DAUS
YAUS
Linear Fit
Kernel Fit
DBEL
YBEL
Linear Fit
Kernel Fit
DFIN
YFIN
Linear Fit
Kernel Fit
DFRA
YFRA
Linear Fit
Kernel Fit
DIRE
YIRE
Linear Fit
Kernel Fit
DITA
YITA
Linear Fit
Kernel Fit
DLUX
YLUX
Linear Fit
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DNET
YNET
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DPOR
YPOR
Linear Fit
Kernel Fit

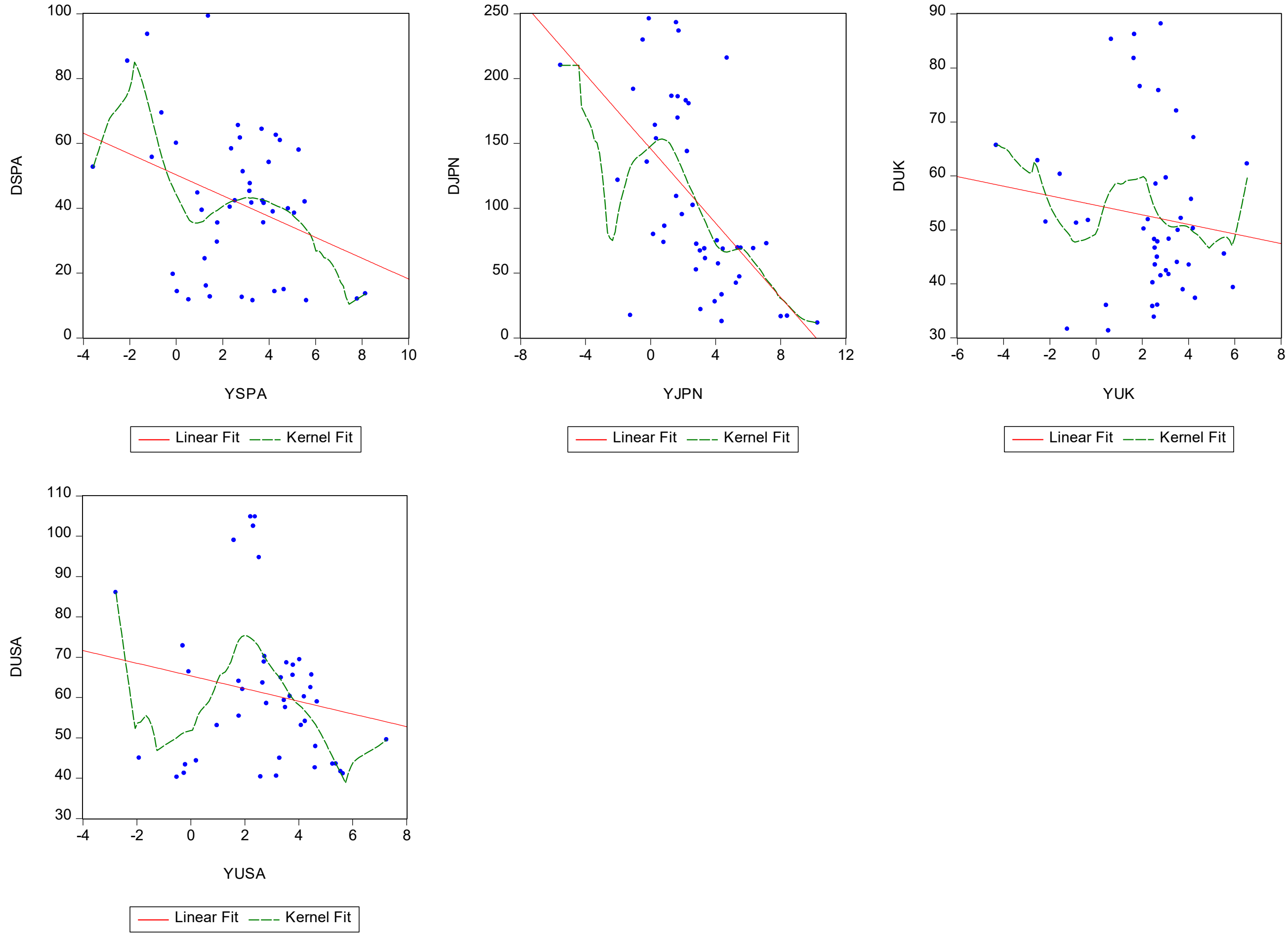
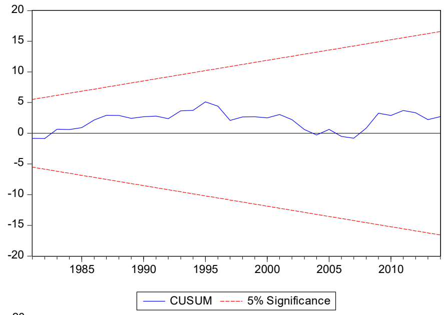
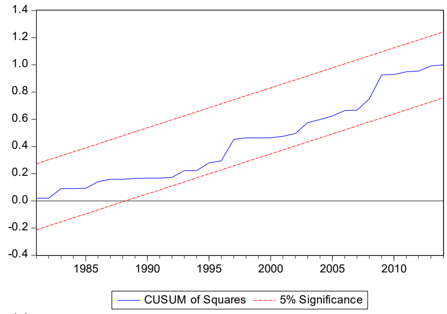


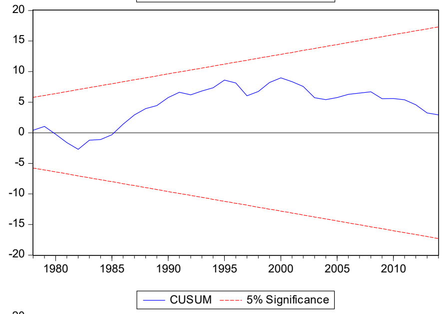
Fig. 2. Scatter plots of the relationship between growth rates (*Y*) and debt to GDP ratios (*D*).

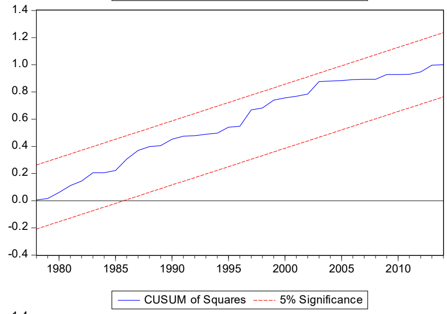
$YAUS \rightarrow DAUS$



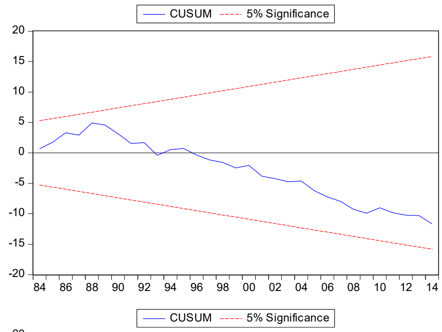


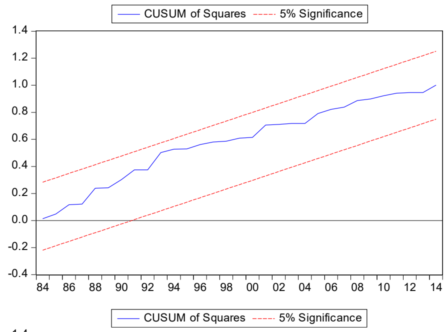
$DAUS \rightarrow YAUS$



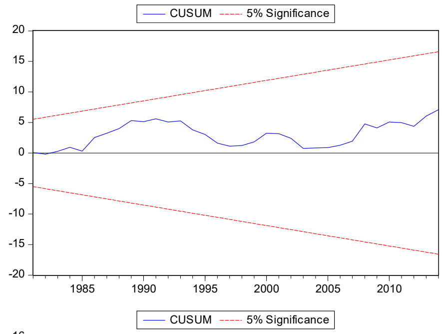


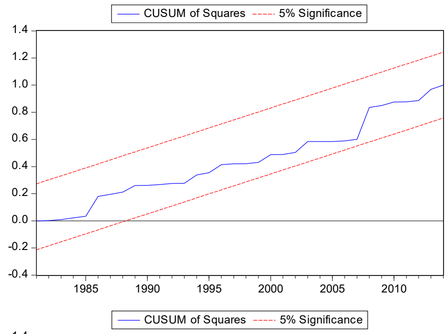
$DFRA \rightarrow YFRA$



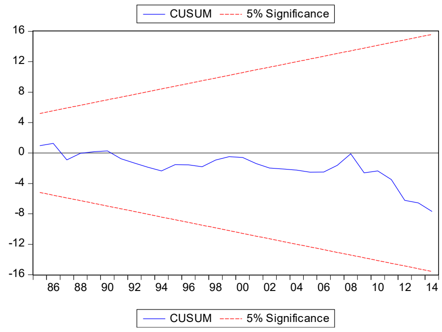


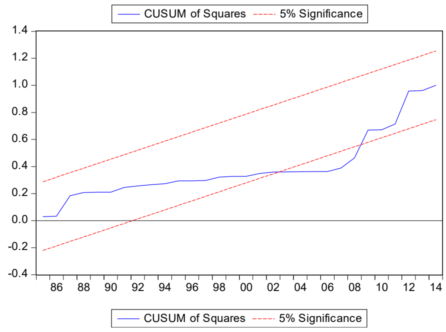
$DLUX \rightarrow YLUX$





$DNET \rightarrow YNET$





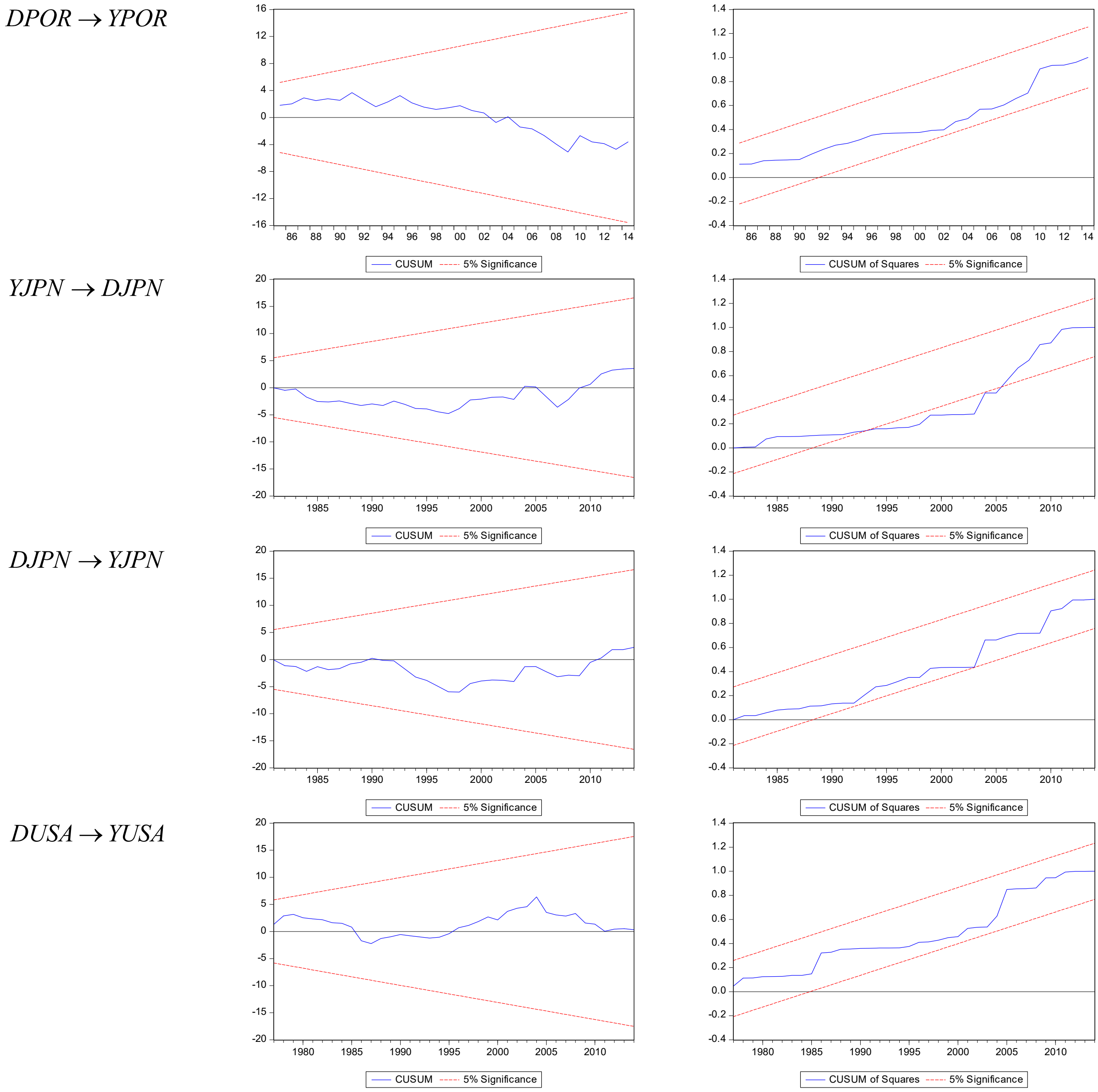


Fig. 3. CUSUM and CUSUMSQ tests on the ARDL models.