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On the bi-directional causal relationship between public debt and economic growth in EMU countries

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Abstract

New evidence is presented on the possible existence of bi-directional causal relationships between public debt and economic growth in both central and peripheral countries of the European Economic and Monetary Union. We test for heterogeneity in the bi-directional Granger-causality across both time and space during the period between 1980 and 2013. The results suggest evidence of a "diabolic loop" between low economic growth and high public debt levels in Spain after 2009. For Belgium, Greece, Italy and the Netherlands debt has a negative effect over growth from an endogenously determined breakpoint and above a debt threshold ranging from 56% to 103% depending on the country.

Keywords: Public debt, economic growth, Granger-causality, euro area, peripheral EMU countries, central EMU countries.

JEL Classification Codes: C22, F33, H63, O40, O52

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1. Introduction

Shambaugh (2012) pointed out that the European Economic and Monetary Union (EMU) countries have faced three interlocking crises (banking, sovereign debt, and economic growth) which together challenged the viability of the currency union. According to this line of thought, these crises connected with one another in several ways: the problems of weak banks and high sovereign debt were mutually reinforcing, and both were exacerbated by weak, constrained growth.

Whilst Gómez-Puig, Sosvilla-Rivero and Singh (2015) focused on the interconnection between banking and sovereign debt crises in EMU countries, in this paper we will pay attention to the interrelationship between public debt and economic growth. The empirical literature on this topic not only presents ambiguous results, but is centred mainly on the possible impact of high debt levels on economic growth, disregarding the possibility of reverse causality running from growth to debt [Ferreira (2009) and Puente-Ajovín and Sanso-Navarro (2015) are some of the few exceptions]. However, there are some theoretical reasons for thinking that public debt is likely to accumulate when growth is low (see Bell *et al.*, 2015). In this regard, since low growth means lower government revenue, governments may be forced to increase their debt levels to maintain the welfare state.

Therefore, some kind of "diabolic loop" (see Brunnemeier *et al.*, 2011)[1] between high levels of debt and low levels of economic growth may have arisen with the outbreak of the financial crisis in most economies in 2008 (EMU countries among them). An increase in government debt levels both raises government deficits and changes investors' expectations. On the one hand, the rise in the government deficit increases the need to borrow and pushes up interest rates; this, through a "crowding-out" effect on private investment, may undermine economic growth. On the other, the increase in investors' risk

[1] These authors described the development of a "diabolic loop" between sovereign and banking risk in euro area countries.

aversion may induce a "flight-to-safety" that may favour bonds of countries generally regarded to have a low default risk. In turn, as economic growth decreases, since government revenues also decline, governments are forced to increase their debt levels in order to maintain the welfare state (the economy's "automatic stabilizers" may start to work, meaning an increase in public expenditure in social security transfers, for instance).

The recent European debt crisis indicates open academic questions that policy makers might need answer to. Events of the last few years have led to increasing concern about the possibly adverse consequences of the substantial accumulation of public debt of EMU countries. The debate is currently so hotly contested, because pundits draw widely different conclusions for macroeconomic policy, and in particular pro and contra economic austerity policies.

In this context, the main objective of this paper is to shed some additional light on this challenging avenue of research by applying the Granger-causality approach and endogenous breakpoint tests in order to analyse the evolution of bi-directional causality between sovereign debt and economic growth in 11 EMU countries (both central and peripheral) during a time period that spans from 1980 to 2013. Therefore, we will consider the potential heterogeneity in the bi-directional causality across both time and space (periods and countries).

Although some authors have applied the Granger methodology to examine causality between these two variables for a group of countries member of the Organization for Economic Co-operation and Development (OECD) by means of panel techniques[2], to our knowledge the bi-directional causal relationship between public debt and economic growth and its evolution over time has not been examined individually for each EMU country. In

[2] Panizza and Presbitero (2014) also investigate causation (not correlation) between sovereign debt levels and economic growth, adopting an instrumental variable approach that captures valuation effects caused by the interaction between foreign currency debt and exchange rate volatility.

particular, in order to compare our results with those of the existing literature, our paper is closely related to those of Ferreira (2009) and Puente-Ajovín and Sanso-Navarro (2015). However, in contrast to those studies, we do not make use of panel Granger-causality tests to combine the power of cross section averaging with all the subtleties of temporal dependence; rather, we explore the time series dimension of the issue. Our methodology is data-driven and allows us to select the statistical model that best approximates the relationship between the variables under study for any particular country. We follow Ferreira (2009) regarding the need to pre-test for unit roots and cointegration (not required in the approach used by Puente-Ajovín and Sanso-Navarro, 2015), but we rely on direct estimation of individual relationships rather than on bootstrapping country-specific effects, as in Puente-Ajovín and Sanso-Navarro (2015). Moreover, our approach allows the possible Granger-causal relationships to change during the sample period (see Gómez-Puig and Sosvilla-Rivero, 2014) and to determine endogenously the breakpoints in the evolution of those relationships, thus permitting us to evaluate them before and after such breakpoints.

In our opinion, the proposed analysis is particularly relevant in an environment in which deleveraging policies need to be implemented in some euro area countries, but their beneficial effects are still a matter for debate. The rest of the paper is organized as follows. Section 2 presents a short literature review. Section 3 explains the econometric methodology, and the data and empirical results are reported in Section 4. Finally, some concluding remarks and policy implications are provided in Section 5.

2. Literature review

The results from the empirical literature on the relationship between public debt and economic growth are far from being conclusive (see Panizza and Presbitero (2013) for a survey). While the first studies [see, for example, Modigliani (1961), Diamond (1965) and

Saint-Paul (1992)] sustained that a public debt increase always contributed to economic growth, more recent work has presented totally different results. Patillo *et al.*, (2004) conclude that whilst low levels of public debt affect economic growth positively, high levels have a negative impact; Schclarek (2005) does not find any significant relation between public debt and economic growth in industrial countries, whereas Kumar and Woo (2010), controlling for other factors that also influence growth, detected an inverse relationship between the two variables.

In their seminal work, Reinhart and Rogoff (2010) studied economic growth for different thresholds of public debt using a database of 44 countries over a time period spanning 200 years. Their results suggest that the relationship is weak for public debt ratios below 90% of GDP, but that, above this threshold, on average, growth rates decrease substantially. However, since the publication of their paper, the 90% threshold has not only been questioned but has also been the focus of much of the debate in the literature: see Cecchetti *et al.* (2011), Minea and Paren (2012), Presbitero (2012), Baum *et al.* (2012), Checherita-Westphal and Rother (2012), Herdon *et al.* (2013) and Égert (2013), to name a few.

Moreover, the recent global recession and sovereign debt crisis in Europe have stimulated an intense debate both on the effectiveness of fiscal policies and on the consequences of public debt increases, in a situation in which leverage is already very high in European economies. In this regard, Gómez-Puig (2013) attempts to quantify the total level of indebtedness (public and private) in all euro area countries, using a database created with the statistics provided by the European Central Bank. According to her calculations, total leverage (public and private) over GDP recorded levels of 710%, 487%, 413%, 360% and 353%, in September 2012, in Ireland, Portugal, Spain, Italy and Greece respectively.

There is currently no consensus among economists in this area. While some suggest that now is precisely the time to apply the lessons learnt during the Great Depression and that policymakers should implement expansionary fiscal policies [see, among others, Krugman (2011), Berg and Ostry (2011) or DeLong and Summers (2012)][3], others argue that, since the high level of public sector leverage has a negative effect on economic growth, fiscal consolidation is fundamental to restoring confidence and improving expectations about the future evolution of the economy (Cochrane, 2011). The latter approach, which supports austerity measures, has been highly influential among the EMU authorities and has the support of the empirical evidence presented in some influential papers (Reinhart and Rogoff, 2010, among them).

Nevertheless, in our reading of the empirical evidence, no strong case has yet been made for a causal bi-directional relationship between public debt and economic growth in EMU countries, taking into account not only cross-country heterogeneity in the euro area, but also its time-varying nature. This paper aims to fill this gap in the literature.

3. Econometric methodology

3.1 Testing for causality

The concept of Granger-causality was introduced by Granger (1969) and Sims (1972) and is widely used to ascertain the importance of the interaction between two series. The central notion is one of predictability (Hoover, 2001): one variable Granger-causes some other variable, given an information set, if past information about the former can improve the forecast of the latter based only on its own past information. Therefore, knowledge of the evolution of one series reduces the forecast errors of the other, suggesting that the latter does not evolve independently of the former.

[3] These authors state that deleveraging policies can even prove to be detrimental, depending on the fundamental variables of the economy. Their argument currently receives support from some politicians in Southern Europe.

Testing Granger causality typically uses the same lags for all variables. This poses a potential problem, since Granger-causality tests are sensitive to lag length. Therefore, it is important that the lengths selected should be the right ones so as to avoid inconsistently estimating the model and drawing misleading inferences (see Thornton and Batten, 1985). In determining the optimal lag structure for each variable, we follow Hsiao's (1981) sequential method to test for causality, which combines Akaike (1974)'s final predictive error (FPE, from now on) and the definition of Granger-causality[4]. Essentially, the FPE criterion trades off the bias that arises from under-parameterization of a model against a loss in efficiency resulting from its over-parameterization, thus avoiding the ambiguities of the conventional procedure.

Consider the following models

$$X_{\mathrm{t}} = \alpha_0 + \sum_{i=1}^{m} \delta_i X_{t-i} + \varepsilon_t \qquad (1)$$

$$X_t = \alpha_0 + \sum_{i=1}^{m} \delta_i X_{t-i} + \sum_{j=1}^{n} \gamma_j Y_{t-j} + \varepsilon_t \qquad (2)$$

where X_t and Y_t are stationary variables [i.e., they are I(0) variables]. The following steps are used to apply Hsiao's procedure for testing Granger-causality:

i) Treat X_t as a one-dimensional autoregressive process (1), and compute its FPE with the order of lags *m* varying from 1 to *m*[5]. Choose the order which yields the smallest FPE, say *m*, and denote the corresponding FPE as FPE_X (m, 0).

ii) Treat X_t as a controlled variable with *m* number of lags, and treat Y_t as a manipulated variable as in (2). Compute again the FPE of (2) by varying the order

[4] Thornton and Batten (1985) show that Akaike (1974)'s FPE criterion performs well relative to other statistical techniques.

[5] FPE_X(m,0) is computed using the formula: $FPE_X(m,0) = \frac{T+m+1}{T-m-1} \cdot \frac{SSR}{T}$, where T is the total number of observations and SSR is the sum of squared residuals of OLS regression (1)

of lags of Y_t from 1 to *n*, and determine the order which gives the smallest FPE, say *n*, and denote the corresponding FPE as FPE_X (m,n)[6].

iii) Compare FPE_X (m, 0) with FPE_X (m,n) [i.e., compare the smallest FPE in step (i) with the smallest FPE in step (ii)]. If FPE_X (m,0) > FPE_X (m,n), then Y_t is said to cause X_t. If FPE_X (m,0) < FPE_X (m,n), then X_t is an independent process.

iv) Repeat steps i) to iii) for the Y_t variable, treating X_t as the manipulated variable.

When X_t and Y_t are not stationary variables but are first-difference stationary [i.e., they are I(1) variables] and cointegrated (see Dolado *et* al., 1990), it is possible to investigate the existence of a Granger-causal relationships from ΔX_t to ΔY_t and from ΔY_t to ΔX_t, using the following error correction models:

$$\Delta X_t = \alpha_0 + \sum_{i=1}^{m} \delta_i \Delta X_{t-i} + \varepsilon_t \tag{3}$$

$$\Delta X_t = \alpha_0 + \beta Z_{t-1} + \sum_{i=1}^{m} \delta_i \Delta X_{t-i} + \sum_{j=1}^{n} \gamma_j \Delta Y_{t-j} + \varepsilon_t \tag{4}$$

where Z_t is the ordinary least square (OLS) residual of the cointegrating regression ($X_t = \mu + \lambda Y_t$), known as the error-correction term. Note that, if X_t and Y_t are I (1) variables, but they are not cointegrated, then β in (4) is assumed to be equal to zero.

In both cases [i.e., X_t and Y_t are I(1) variables, and they are or are not cointegrated], we can use Hsiao's sequential procedure substituting X_t with ΔX_t and Y_t with ΔY_t in steps i) to iv), as well as substituting expressions (1) and (2) with equations (3) and (4). Proceeding in this way, we ensure efficiency since the system is congruent and encompassing (Hendry and Mizon, 1999).

[6] FPE_X(m,n) is computed using the formula: $FPE_X(m,n) = \frac{T+m+n+1}{T-m-n-1} \cdot \frac{SSR}{T}$, where T is the total number of observations and SSR is the sum of squared residuals of OLS regression (2)

3.2 Stability Diagnostics

In conventional Granger-causality analysis, the relationship between two variables is assumed to exist at all times. However, in a time period that spans more than 30 years, parameter non-constancy may occur (with the advent of either a financial or economic crisis, or due to some other kind of shock) and may generate misleading inferences if left undetected (see Bai and Perron, 1998, 2003; Perron, 1989; Zivot and Andrews, 1992). Furthermore, the pre-testing issue in early studies may have induced a size distortion of the resulting test procedures (Bai, 1997). Thus, it is desirable to let the data select when and where regime shifts occur (i. e., we need to test for the null hypothesis of no structural change *versus* the alternative hypothesis that changes are present). To this end, we first identify a single structural change using the Quandt–Andrews one-time unknown structural break test. We then use the procedure suggested by Bai (1997) and Bai and Perron (1998, 2003) to detect multiple unknown breakpoints in order to obtain further evidence of the existence of the breakpoints previously detected endogenously.

3.2.1 Quandt-Andrews Breakpoint Test

A particular challenge in empirical time series analysis is the determination of the appropriate timing of a potential structural break. In a traditional Chow (1960) test[7], we have to set a specific breakpoint based on *a priori* knowledge about the potential break date. In our analysis, however, we do not assume any prior knowledge of potential break dates, but use a data-based procedure to determine the most likely location of a break. In particular, we use the Quandt–Andrews unknown breakpoint test, originally introduced by Quandt (1960) and later developed by Andrews (1993) and Andrews and Ploberger (1994). The idea behind the Quandt-Andrews test is that a single Chow breakpoint test is

[7] The basic idea of the breakpoint Chow test is to fit the equation separately for each subsample and to see whether there are significant differences in the estimated equations. A significant difference indicates a structural change in the relationship.

performed at each observation between two dates, or observations (τ_1 and τ_2). The k test statistics from those Chow tests are then summarized into one test statistic for a test against the null hypothesis of no breakpoints between τ_1 and τ_2.

For the unknown break date, Quandt (1960) proposed likelihood ratio test statistics for an unknown change point, called the Supremum test, while Andrews (1993) supplied analogous Wald and Lagrange Multiplier test statistics for it. Andrews and Ploberger (1994) then developed Exponential (LR, Wald and LM) and Average (LR, Wald and LM) tests, calculated by using individual Chow statistics for each date in the sample, except for a portion trimmed from both ends. While the Supremum test finds the date that maximizes Chow statistics, i.e., the most likely break point, the Average and Exponential tests use all the Chow statistic values and provide information about the existence of the break, not about its date[8].

We set a search interval $\tau \in [0.15, 0.85]$ for the full sample T to allow a minimum of 15% of effective observations contained in both pre- and post-break periods. These tests allow us to determine a structural change with unknown timing endogenously from the data after examining each date in the data except for a portion trimmed from both ends.

3.2.2 Multiple Breakpoint Tests

Bai and Perron (1998) develop tests for multiple structural changes. Their methodology can be divided in two separate and independent parts. First, they propose a sequential method to identify any number of breaks in a time series, regardless of their statistical significance. Second, once the breaks have been identified, they propose a series of statistics to test for the statistical significance of these breaks, using asymptotic critical values.

[8] Andrews (1993) and Andrews and Ploberger (1994) provide tables of critical values, and Hansen (1997) provides a method for calculating p-values.

The sequential procedure is as follows:

i. Begin with the full sample and perform a test of parameter constancy with unknown break.

ii. If the test rejects the null hypothesis of constancy, determine the break date, divide the sample into two samples and perform single unknown breakpoint tests in each subsample. Add a breakpoint whenever the null is rejected in a subsample.

iii. Repeat the procedure until none of the subsamples reject the null hypothesis, or until reaching the maximum number of breakpoints allowed or the maximum number of subsample intervals to test.

For a specific set of unknown breakpoints $(T_1,...,T_p)$, we use the following set of tests developed by Bai and Perron (1998, 2003) to detect multiple structural breaks: the sup *F* type test, the double maximum tests, and the test for ℓ versus $\ell+1$ breaks. First, we consider the sup *F* type test with no structural breaks ($p=0$) versus the alternative hypothesis that there are $p=k$ breaks. Second, we use the double maximum tests, *UD*max and *WD*max, testing the null hypothesis of no structural breaks against an unknown number of breaks given some upper bound *m**. Finally, we study the sup $F_T\left(\ell+1/\ell\right)$ test, which is a sequential test of the null hypothesis of ℓ breaks against the alternative of $\ell+1$ breaks. The test is applied to each segment containing the observations $\hat{T}_{i-1}$ to $\hat{T}_i$ $i=1,\ldots,(\ell+1)$. To run these tests it is necessary to decide the minimum distance between two consecutive breaks, *h*, which is obtained as the integer part of a trimming parameter, ε, multiplied by the number of observations *T* (we use $\varepsilon=0.15$ and allow up to four breaks).

4. Data and empirical results

4.1. Data

We use annual data for the variables government debt-to-Gross Domestic Product (GDP) (denoted *DEBT*) and real economic growth (measured as the percentage change of GDP product at constant prices, denoted by *GROWTH*) from 1980 to 2013 collected from the International Monetary Fund's World Economic Outlook Database for both EMU-11 peripheral (Greece, Ireland, Italy, Portugal and Spain) and central countries (Austria, Belgium, Finland, France, Germany and the Netherlands) [9].

4.2. Preliminary results

As a first step, we tested for the order of integration of the *DEBT* and *GROWTH* by means of the Augmented Dickey-Fuller (ADF) tests. The results, shown in Table 1, decisively reject the null hypothesis of non-stationarity for *DEBT* but not for *GROWTH*, suggesting that our indicator of public debt can be treated as first-difference stationary and real growth as stationary[10].

[Insert Table 1 here]

Following Cheung and Chinn's (1997) suggestion, we confirm these results using the Kwiatkowski et al. (1992) (KPSS) tests, where the null is a stationary process against the

[9] This distinction between central and peripheral countries has been extensively used in the empirical literature. The two groups we consider roughly correspond to the distinction made by the European Commission (1995) between those countries whose currencies continuously participated in the European Exchange Rate Mechanism (ERM) from its inception maintaining broadly stable bilateral exchange rates among themselves over the sample period, and those countries whose currencies either entered the ERM later or suspended its participation in the ERM, as well as fluctuating in value to a great extent relative to the Deutschmark. These two groups are also roughly the same found in Jacquemin and Sapir (1996), applying multivariate analysis techniques to a wide set of structural and macroeconomic indicators, to form a homogeneous group of countries. Moreover, these two groups are basically the same that those found in Ledesma-Rodríguez et al. (2005) according to the perception of economic agents with respect to the commitment to maintain the exchange rate around a central parity in the ERM and those identifying by Sosvilla-Rivero and Morales-Zumaquero (2012) using cluster analysis when analysing permanent and transitory volatilities of EMU sovereign yields.

[10] These results were confirmed using Phillips-Perron (1998) unit root tests controlling for serial correlation and the Elliott, Rothenberg, and Stock (1996) Point Optimal and Ng and Perron (2001) unit root tests for testing non-stationarity against the alternative of high persistence. These additional results are not shown here in order to save space, but they are available from the authors upon request.

alternative of a unit root. As can be seen in Table 2, the results fail to reject the null hypothesis of stationarity in first differences for both variables and for growth in levels, but reject it strongly for public debt in levels, thus giving further support to our previous conclusions.

[Insert Table 2 here]

Therefore, our preliminary results indicate that, on the one hand, the relevant model for testing for Granger-causality from (percentage change in) government debt to growth is the following one:

$$GROWTH_t = \alpha_0 + \sum_{i=1}^{m} \delta_i GROWH_{t-i} + \varepsilon_t \tag{5}$$

$$GROWTH_t = \alpha_0 + \sum_{i=1}^{m} \delta_i GROWTH_{t-i} + \sum_{j=1}^{n} \gamma_j \Delta DEBT_{t-j} + \varepsilon_t \tag{6}$$

where Δ denotes first difference (i. e., $\Delta DEBT_t = DEBT_t - DEBT_{t-1}$).

On the other hand, the relevant model for testing for Granger-causality from growth to (percentage change in) government debt is the following one[11]:

$$\Delta DEBT_t = \alpha_0 + \sum_{i=1}^{m} \delta_i \Delta DEBT_{t-i} + \varepsilon_t \tag{7}$$

$$\Delta DEBT_t = \alpha_0 + \sum_{i=1}^{m} \delta_i \Delta DEBT_{t-i} + \sum_{j=1}^{n} \gamma_j GROWTH_{t-j} + \varepsilon_t \tag{8}$$

[11] Note that, since the variables are of different order of integration, they cannot be cointegrated and therefore there is not an error correction term in equations (5) to (8). Although the Granger representation theorem (Engle and Granger, 1987) states that if two time-series are cointegrated, then there must be Granger causality between them (either one-way or in both directions), the converse is not true and we can still find causality between these variables.

4.3. Empirical results for the whole sample

The resulting FPE statistics for the whole sample (1980-2013) are reported in Tables 3a and 3b.[12]

[Insert Tables 3 here]

Tables 3a and 3b show that in two cases (Finland and France) our results suggest bidirectional Granger-causality; whilst in five cases (Belgium, Germany, Greece, Ireland and Portugal) we find no evidence of Granger-causality relationships in any of the directions. However, for the rest of the countries (Austria, Italy, the Netherlands and Spain) we find evidence of Granger-causality only running from the percentage change in sovereign debt to economic growth. The last column in Table 3 reports the static long-run solution obtained from the estimated dynamic results, which can be taken as an indicator of the impact of one variable over the other (see Hendry, 1997, Chapter 7). As can be seen in Table 3a, it ranges from 0.1079 to 1.2152, suggesting a positive (and in four of the six cases very high) effect of variations in public debt on real growth; whilst Table 3b displays values within the interval (0.7107, 0.8617), also indicating a positive and high impact of economic growth on changes in public debt-to-GDP ratios.

Summing up, when analysing the whole sample period, we find more evidence of causality running from changes in government debt to growth (we find evidence of causality in seven out of the eleven cases studied) rather than from growth to modifications in public debt (for which there is empirical evidence only in the cases of Finland and France). However, it is important to stress that in the cases where we find evidence of Granger-causality between these two variables, the effect is not negative but positive. This finding suggests that in some countries, in line with the first studies by Modigliani (1961), Diamond

[12] These results were confirmed using both Wald statistics to test the joint hypothesis $\hat{\gamma}_1 = \hat{\gamma}_2 = ... = \hat{\gamma}_n = 0$ in equations (6) and (8) and the Williams-Kloot test for forecasting accuracy (Williams, 1959). These additional results are not shown here in order to save space, but they are available from the authors upon request.

(1965) or Saint-Paul (1992) which contended that a public debt increase always contributed to economic growth, there is evidence of a positive effect of changes in government debt on growth, and *vice-versa.*

4.4. Empirical results allowing for structural breaks

In order to gain further insight into the dynamic Granger-causality between the percentage change in the debt-to-GDP ratio and growth, we now examine the existence of breakpoints in causality evolution in order to analyse bi-directional Granger-causality before and after the detected break dates. To do so, we do not set a specific breakpoint based on *a priori* knowledge about the potential break date; we first apply the Quandt-Andrews breakpoint test and let the data select when regime shifts occur in each potential causal relationship, and later we confirm the breakpoint identified by using the tests developed by Bai and Perron (1998, 2003) for detecting multiple structural breaks[13].

The results of our tests for detecting breakpoints are presented in Tables 4a and 4b[14]. In particular, these tables show that in 16 out of the 22 relationships between percentage changes in public debt and growth (73% of the total cases), the break date takes place after 2007 (coinciding with the beginning of the global financial crisis) and in more than a third of the cases (36%) it takes place in 2009 (matching the outbreak of the European debt crisis). The exceptions are Finland, with a break date for the causal relationship that runs from growth to alterations in sovereign debt in 1995 (which might be explained by the fact that Finland joined the European Union on 1 January 1995[15]); Germany, with a break date for the causal relationship from variations in public debt to growth in 1992, and one for the relationship from growth to changes in the public debt-to-GDP ratio in 1990 (just after

[13] We compute the breakpoint tests using a statistic which is robust to heteroskedasticity, since we estimate our original equations with Newey and West (1987) standard errors.
[14] In order to save space, the numerical results of Quandt-Andrews and Bai-Perron tests are not reported in Tables 4a and 4b, but are available from the authors upon request.
[15] The accession of Finland to the European Union took place in a difficult economic environment after the early 1990s recession, but reinforced its tradition of fiscal responsibility. See Honkapohja and Koskela (1999) and OECD (2010).

German reunification in 1990, which led to a boom in demand and higher inflation which the Bundesbank countered with restrictive policy measures)[16]; Greece with a break point in the relationship that runs from differences in government debt to economic growth in 1995 (a year identified as a turning point for the Greek economy after the Bank of Greece's adoption of the "hard drachma policy"- see, for example, Bryant *et al.*, 2001 or Trachanas and Katrakilidis, 2013[17]); and Ireland, with a break date for the causal relationship that runs from growth to modifications in sovereign debt in 1988 (when the government debt-to-GDP ratio reached its peak, with levels close to 110%, before the financial crisis, in a complex economic environment with unemployment and interest rates running into double figures and emigration increasing).

Regarding the Granger-causality before and after the endogenous detected breakpoint, Table 4a shows that only in two countries (Ireland and Portugal) there was no evidence of Granger-causality between changes in public debt and growth at any point in the period, whilst in the case of Germany, we find evidence of a positive causal relationship between variations in the debt-to-GDP ratio and growth only after the break point (1992). Nevertheless, in the eight remaining countries causality between alterations in sovereign debt and economic growth decreases after the break point. In three countries (Austria, Finland and France), the causal relationship declines but remains positive after the break date; while in the case of Belgium, Greece, Italy, the Netherlands and Spain the causal relationship between variations in public debt and growth is positive before the break but becomes negative after that date.

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[16] On 18 May 1990, the two German states (the German Democratic Republic and the Federal Republic of Germany) signed a treaty agreeing on monetary, economic and social union. Reunification was accompanied by a massive domestic fiscal expansion as the Bonn government borrowed in order to rebuild East Germany and transfer income to its relatively poor citizens who, for their part, began to consume more and demanded higher wages.

[17] The aim of this policy, which meant that the exchange rate was used as a nominal anchor, was to improve the convergence of the Greek economy toward the remainder of the EU. We should recall that in 1995, the ratio of government debt-to-GDP reached levels close to 100% in Greece (see Figure 1b), far higher than those required by the Maastricht Treaty (60%, a figure that was never fulfilled).

With respect to the causal relationship running from growth to the percentage change in the public debt-to-GDP ratio (Table 4b), we found no evidence for it in any of the periods studied for four out of the eleven countries: Austria, Greece, Italy and Portugal. In the remaining seven cases the evidence is mixed. While in three cases (Finland, Ireland and Spain) the causal relationship decreases after the break – indeed, in Finland and Spain the relationship between growth and variations in government debt is negative after the break date, suggesting that low economic growth might lead to an increase in the level of government debt – in the other four cases (Belgium, France, Germany and the Netherlands) the causal relationship between growth and the percentage change in the public debt-to-GDP ratio registers a positive increase after the break date detected in each individual country.

Since the evidence seems much more conclusive in the relationships that run from changes in public debt to economic growth, we have plotted the evolution of the two variables, jointly with the break date in the relationship for both central and peripheral countries.

[Insert Figures 1a and 1b here]

If we focus our analysis on central countries (Figure 1a), we observe that in Belgium, variations in public debt began to Granger-cause negatively economic growth from 2009 onwards when the debt-to-GDP ratio was around 96.5%. In the Netherlands, changes in government debt began to have a negative effect on growth also from 2009 onwards, when the debt-to-GDP reached a level of 56%.

In the case of peripheral countries (Figure 1b), in Greece, shifts in sovereign debt have a negative effect on growth after 1995 when the debt-to-GDP level was close to 97%. In Italy, we find evidence that changes in public debt Granger-cause economic growth after 2007 when the debt-to-GDP presented a value of 103%. Finally, in Spain, variations in government debt begin to have a negative effect on real growth after 2009 with a debt-to-GDP ratio of 53%.

This evidence suggests that there is a negative relationship between changes in public debt and economic growth in five of the eleven countries studied after the detected break date, and that the debt threshold above which the relationship becomes negative differs according to the country (see Table 4a). According to our evidence, this threshold ranges from 53% in the case of Spain to 103% in the case of Italy. These results are highly significant, since they not only reinforce the idea that causality should be examined from a dynamic point of view, but also that a specific analysis for each country is necessary, since the debt threshold above which an increase in the level of indebtedness might undermine economic performance varies according to the economic situation of each country. Moreover, three of the five cases in which we found that, after an endogenously determined break date, variations in public debt had a negative effect on growth corresponded to peripheral countries: Spain (from 2009 and a debt-threshold of 53%), Greece (from 1995 and a debt-threshold of 97%) and Italy (from 2009 and a debt-threshold of 103%). Therefore, these results may have important implications for policymakers, in a context, in which they must manage the trade-off between policies that might enhance growth and those that favour a deleveraging of the economy.

Finally, it is also noticeable that growth negatively Granger-causes percentage changes in public debt, after an endogenously determined break date, in the case of Finland and Spain (see Table 4b). In particular, in the case of Spain, we might conclude that our empirical evidence suggests the existence of a "diabolical loop" between low economic growth (the threshold rate seems to be 0.9%) and high public debt levels after 2009.

5. Concluding remarks and policy implications

Although the empirical literature has studied the effects of high levels of debt on economic growth, to our knowledge, no strong case has yet been made for a causal bi-directional relationship between (changes in) public debt and economic growth in EMU countries, taking into account not only cross-country heterogeneity in the euro area, but also its time-varying nature. The aim of this paper was to fill this gap in the literature. In order to consider the potential heterogeneity in the bi-directional causality across both time and space (periods and countries), we applied the Granger-causality approach and endogenous breakpoint tests between public debt and economic growth between 1980 and 2013.

The results for the whole sample between 1980 and 2013 are not at odds with those presented by other authors, but they do qualify them. As in our study, Panniza and Presbitero (2014) and Puente-Ajovín and Sanso-Navarro (2015) do not find negative causation (the latter examined Granger-causality) between sovereign debt and economic growth for a sample period that finishes in 2009.

However, when we analyse Granger-causality after endogenously detecting a breakpoint (which, in most cases, takes place between 2007 and 2009), we find evidence of negative Granger-causality between changes in sovereign debt and growth in some of the countries studied between the break date and the end of the sample period in 2013.

As in every empirical analysis, the results must be taken with caution since they are based on a set of countries over a certain period and on a given econometric methodology. However, our findings suggest that there is evidence of a "diabolic loop" between low economic growth and high public debt levels in Spain since 2009. Moreover, in the case of Belgium, Greece, Italy and the Netherlands, variations in public debt have a negative effect over growth after an endogenously determined breakpoint and above a debt threshold that ranges from 56% to 103%, depending on the country[18]. In our opinion, these results are particularly relevant in an environment where, in a context of low economic growth, deleveraging policies need to be implemented in certain euro area countries and their beneficial effects are still being debated.

Taken together, our findings argue in favour of confronting the excessive fiscal imbalances in three southern EMU countries that have been hit by the recent sovereign debt crisis: Italy, Greece, and Spain, although they still register low rates of economic growth. The empirical evidence suggests that an increase in the level of public indebtedness, which might be accompanied by a relaxation of austerity programs, may not boost economic growth, but increase its decline. Given the importance of these policy implications, further research is required to better understand and quantify the effect of public debt on economic growth.

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[18] Although we use a completely different methodology, our results are in concordance with those presented by the authors who found that above a certain debt-to-GDP threshold an increase in public indebtedness had a negative effect on economic growth [see Reinhart and Rogoff (2010), Cecchetti *et al.* (2011), Minea and Paren (2012), Presbitero (2012), Baum *et al.* (2012), Checherita-Westphal and Rother (2012), and Égert (2013), to name a few]

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Table 1. Augmented Dickey-Fuller tests for unit roots

Panel A: I (2) versus I (1) (Variables in first differences)			
	τ_τ	τ_μ	τ
ΔATDEBT	-5.5165*	-5.9523*	-3.5004*
ΔATGROWTH	-4.8275*	-5.5223*	-5.6278*
ΔBEDEBT	-4.7354*	-4.7354*	-3.0932*
ΔBEGROWTH	-6.2709*	-6.2781*	-6.3878*
ΔFIDEBT	-4.3973*	-3.7552*	-3.6457*
ΔFIGROWTH	-5.8645*	-5.9652*	-6.0526*
ΔFRDEBT	-3.7576*	-3.8345*	-2.6467*
ΔFRGROWTH	-5.8645*	-5.9652*	-6.0526*
ΔGEDEBT	-4.4793*	-3.7383*	-3.3813*
ΔGEGROWTH	-4.4319*	-4.4899*	-4.5908*
ΔGRDEBT	-4.9532*	-3.7961*	-2.7664*
ΔGRGROWTH	-4.5429*	-3.6719*	-2.8259*
ΔIEDEBT	-4.2429**	-3.7123*	-2.9123*
ΔIEGROWTH	-4.8589*	-6.0644*	-6.1605*
ΔITDEBT	-3.7833**	-3.7963*	-3.1035*
ΔITGROWTH	-4.5173*	-4.1937*	-4.8478*
ΔNLDEBT	-3.5620**	-3.5620*	-3.4908*
ΔNLGROWTH	-3.9448**	-3.7223*	-4.5883*
ΔPTDEBT	-4.0966*	-3.7279*	-2.8741*
ΔPTGROWTH	-4.4933*	-6.6656*	-6.6656*
ΔSPDEBT	-4.2195**	-3.7041*	-2.9487*
ΔSPGROWTH	-3.6423**	-3.0983**	-2.2242*
Panel B: I (1) versus I (0) (Variables in levels)			
	τ_τ	τ_μ	τ
ATDEBT	-0.0148	-1.7545	-0.7251
ATGROWTH	-4.6447*	-4.4038*	-2.7344*
BEDEBT	-2.3981	-1.5461	-0.4637
BEGROWTH	-5.1290*	-5.1264*	-3.0833*
FIDEBT	-2.4141	-1.5271	0.5305
FIGROWTH	-3.6080**	-3.6903*	-3.0197*
FRDEBT	-2.5228	0.1262	0.1556
FRGROWTH	-3.6080**	-3.6903*	-3.0197*
GEDEBT	-2.1846	-1.1390	0.7117
GEGROWTH	-4.6796*	-4.3525*	-3.0034*
GRDEBT	-1.5890	-2.0141	-1.2151
GRGROWTH	-3.9533**	-3.1896**	-2.7664*
IEDEBT	-2.2343	-2.5767	-0.4025
IEGROWTH	-4.2246**	-3.9278*	-2.6827*
ITDEBT	-2.4291	-1.9253	0.1092
ITGROWTH	-4.3374*	-3.7448*	-2.8070*
NLDEBT	-2.3467	-2.4335	1.1712
NLGROWTH	-3.5489**	-3.2692**	-2.8502*
PTDEBT	-1.6685	-2.0674	0.8641
PTGROWTH	-3.6980**	-3.6563*	-2.7231*
SPDEBT	-2.3666	-1.5221	0.5903
SPGROWTH	-3.9545**	-3.2041**	-2.8797*

Notes:
The ADF statistic is a test for the null hypothesis of a unit root.
τ_τ, τ_μ and τ denote the ADF statistics with drift and trend, with drift, and without drift respectively.
* denotes significance at the 1% level. Critical values based on MacKinnon (1996)
Δ denotes first difference. AT, BE, FI, FR, GE, GR, IE, IT, NL, PT and SP stand for Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Portugal and Spain respectively.

Table 2. KPSS tests for stationarity

Panel A: I (2) versus I (1) (Variables in first differences)		
	τ_τ	τ_μ
ΔATDEBT	0,1032	0.3314
ΔATGROWTH	0.0986	0.2828
ΔBEDEBT	0.1042	0.3168
ΔBEGROWTH	0.1012	0,2495
ΔFIDEBT	0.0751	0.0826
ΔFIGROWTH	0.1037	0.3077
ΔFRDEBT	0.1118	0.1452
ΔFRGROWTH	0.0539	0.1583
ΔGEDEBT	0.0547	0,0524
ΔGEGROWTH	0.0767	0.3134
ΔGRDEBT	0.1111	0.2116
ΔGRGROWTH	0.0474	0.1592
ΔIEDEBT	0.0971	0.2116
ΔIEGROWTH	0.0574	0.1592
ΔITDEBT	0.1018	0.2036
ΔITGROWTH	0.0913	0.3312
ΔNLDEBT	0.1078	0.2484
ΔNLGROWTH	0.0798	0.3141
ΔPTDEBT	0.1066	0.3395
ΔPTGROWTH	0.0395	0.0382
ΔSPDEBT	0.1067	0.1297
ΔSPGROWTH	0.0893	0.1967
Panel B: I (1) versus I (0) (Variables in levels)		
	τ_τ	τ_μ
ATDEBT	0.2291	0.4955**
ATGROWTH	0.1055	0.1925
BEDEBT	0.6051*	0.5863**
BEGROWTH	0.1027	0.2536
FIDEBT	0.2295*	0.5320**
FIGROWTH	0.0904	0.1580
FRDEBT	0.2660*	0.6803**
FRGROWTH	0.0943	0.2691
GEDEBT	0.1653**	0.7547**
GEGROWTH	0.1012	0.1754
GRDEBT	0.2917*	0.7463*
GRGROWTH	0.1153	0.1724
IEDEBT	0.2483*	0.6771**
IEGROWTH	0.1045	0.1766
ITDEBT	0.1568**	0.5927
ITGROWTH	0.1023	0.1725
NLDEBT	0.2282*	0.5082**
NLGROWTH	0.1012	0.2674
PTDEBT	0.1529**	0.5661**
PTGROWTH	0.0993	0.2502
SPDEBT	0.1968**	0.4650**
SPGROWTH	0.1054	0.1967

Notes:
The KPSS statistic is a test for the null hypothesis of stationarity.
τ_τ and τ_μ denote the KPSS statistics with drift and trend, and with drift respectively * denotes significance at the 1% level. Critical values based on MacKinnon (1996).
Δ denotes first difference. AT, BE, FI, FR, GE, GR, IE, IT, NL, PT and SP stand for Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Portugal and Spain respectively.

Table 3a. FPE statistics, all sample (ΔDebt → Growth)

	FPE(m.0)	FPE(m.n)	Causality	Impact
ΔATDEBT → ATGROWTH	**3.6192 (1.0)**	**3.4518 (1.2)**	**Yes**	**1.2152**
ΔBEDEBT → BEGROWTH	3.4085 (2.0)	3.4521 (2.1)	No	
ΔFIDEBT → FIGROWTH	**10.5553 (1.0)**	**9.1998 (1.1)**	**Yes**	**0.1079**
ΔFRDEBT → FRGROWT	**2.6149 (1.0)**	**2.2787 (1.2)**	**Yes**	**0.7259**
ΔGEDEBT → GEGROWTH	5.5113 (1.0)	5.6604 (1.1)	No	
ΔGRDEBT → GRGROWTH	5.8721 (1.0)	6.2241 (1.1)	No	
ΔIEDEBT → IEGROWTH	9.3414 (1.0)	9.7624 (1.1)	No	
ΔITDEBT → ITGROWTH	**3.8557 (1.0)**	**3.3013 (1.1)**	**Yes**	**0.4256**
ΔNLDEBT → NLGROWTH	**3.4411 (1.0)**	**2.8179 (1.2)**	**Yes**	**0.7021**
ΔPTDEBT → PTGROWTH	4.9873 (1.0)	5.3018 (1.1)	No	
ΔSPDEBT → SPGROWTH	**3.0333 (1.0)**	**3.0098 (1.1)**	**Yes**	**0.7866**

Table 3b. FPE statistics, all sample (Growth → ΔDebt)

	FPE(m.0)	FPE(m.n)	Causality	Impact
ATGROWTH → ΔATDEBT	8.9099 (1.0)	9.5225 (1.1)	No	
BEGROWTH → ΔBEDEBT	10.7905(1.0)	11.3556 (1.1)	No	
FIGROWTH → ΔFIDEBT	**14.2982 (2.0)**	**13.9998 (2.1)**	**Yes**	**0.7107**
FRGROWTH → ΔFRDEBT	**7.2823 (1.0)**	**7.0749 (1.1)**	**Yes**	**0.8617**
GEGROWTH → ΔGEDEBT	4.0397 (1.0)	4.9406 (1.1)	No	
GRGROWTH → ΔGRDEBT	8.7370 (1.0)	9.2978 (1.1)	No	
IEGROWTH → ΔIEDEBT	4.3576 (1.1)	4.6364 (1.1)	No	
ITGROWTH → ΔITDEBT	3.7730 (1.0)	4.2267 (1.1)	No	
NLGROWTH → ΔNLDEBT	4.9931 (1.0)	5.9658 (1.1)	No	
PTGROWTH → ΔPTDEBT	3.3260 (1.0)	3.7053 (1.1)	No	
SPGROWTH → ΔSPDEBT	6.8209 (1.0)	7.8151 (1.1)	No	

Notes:
The figures in brackets are the optimum order of lags in each pair of countries
Impact shows the static long-run solution
Δ denotes first difference. AT, BE, FI, FR, GE, GR, IE, IT, NL, PT and SP stand for Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Portugal and Spain respectively.
Bold values indicate the presence of Granger-causality.

Table 4a Causality analysis by sub-periods (ΔDebt → Growth)

	Before the break	After the break	Break date	Debt threshold
ΔATDEBT → ATGROWTH	**Yes (0.7086)**	**Yes (0.3054)**	**2009**	
ΔBEDEBT → BEGROWTH	No	***Yes (-0.3288)***	2009	**96.5%**
ΔFIDEBT → FIGROWTH	**Yes (0.6651)**	**Yes (0.3065)**	**2008**	
ΔFRDEBT → FRGROWTH	**Yes (0.9652)**	**Yes (0.1273)**	**2009**	
ΔGEDEBT → GEGROWTH	No	***Yes (0.3439)***	1992	
ΔGRDEBT → GRGROWTH	**Yes (0.1900)**	**Yes (-0.7222)**	1995	**97%**
ΔIEDEBT → IEGROWTH	No	No	2008	
ΔITDEBT → ITGROWTH	**Yes (0.5270)**	**Yes (-0.1508)**	**2007**	**103%**
ΔNLDEBT → NLGROWTH	**Yes (2.3476)**	**Yes (-0.1063)**	**2009**	**56%**
ΔPTDEBT → PTGROWTH	No	No	2008	
ΔSPDEBT → SPGROWTH	**Yes (2.8895)**	**Yes (-0.1074)**	**2009**	**53%**

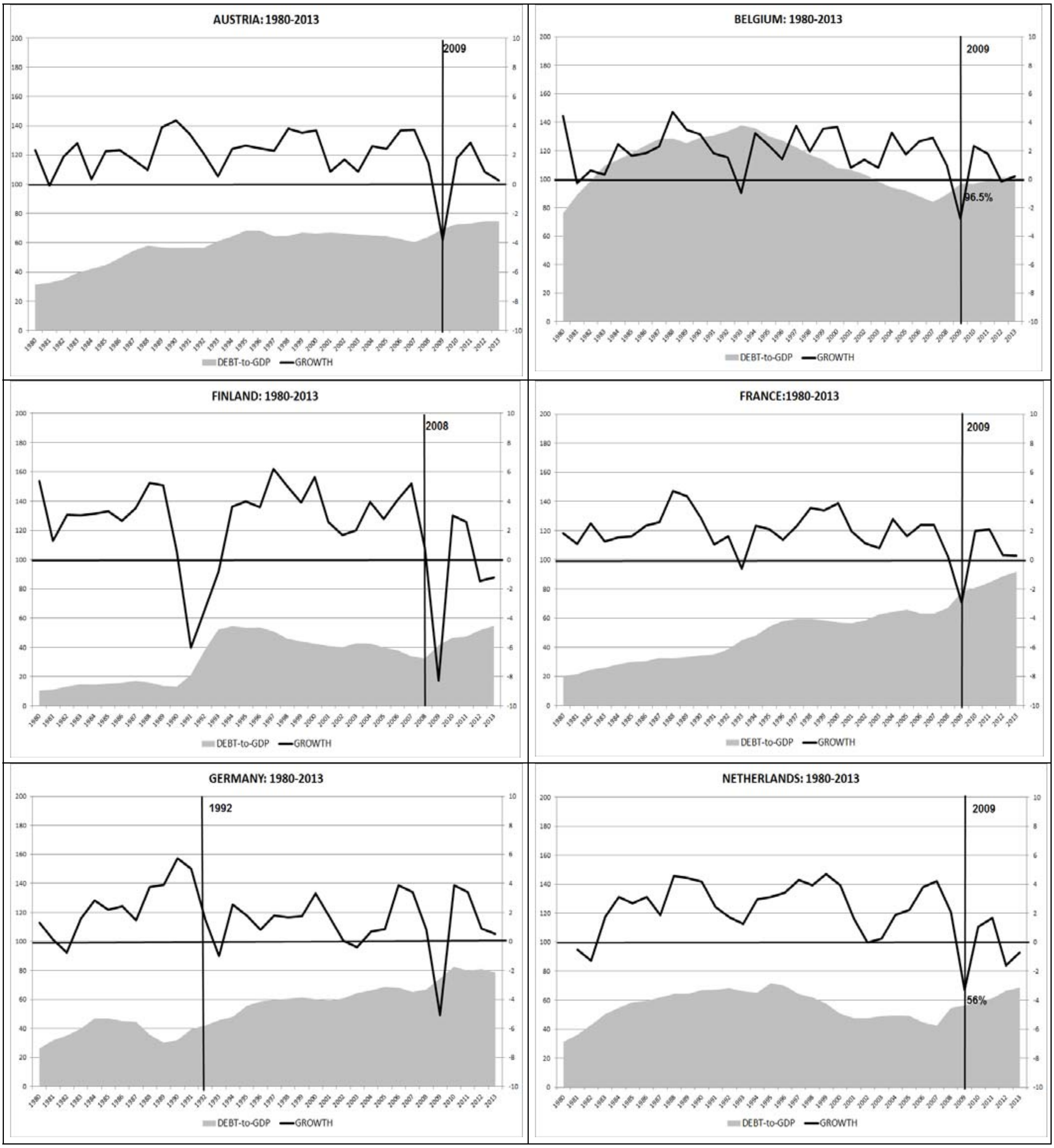
Table 4b Causality analysis by sub-periods (Growth → ΔDebt)

	Before the break	After the break	Break date	Growth threshold
ATGROWTH → ΔATDEBT	No	No	2009	
BEGROWTH → ΔBEDEBT	No	***Yes (4.2152)***	2008	
FIGROWTH → ΔFIDEBT	**Yes (0.4863)**	**Yes (-0.2060)**	**1995**	4%
FRGROWTH → ΔFRDEBT	**Yes (0.4127)**	**Yes (6.3966)**	**2007**	
GEGROWTH → ΔGEDEBT	No	***Yes (0.7742)***	1990	
GRGROWTH → ΔGRDEBT	No	No	2009	
IEGROWTH → ΔIEDEBT	***Yes (3.3019)***	No	1988	
ITGROWTH → ΔITDEBT	No	No	1995	
NLGROWTH → ΔNLDEBT	No	***Yes (1.4729)***	2008	
PTGROWTH → ΔPTDEBT	No	No	2009	
SPGROWTH → ΔSPDEBT	No	***Yes (-12.3380)***	2008	**0.9%**

Notes:
The figures in brackets are the estimated long-run solutions
Δ denotes first difference. AT, BE, FI, FR, GE, GR, IE, IT, NL, PT and SP stand for Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Portugal and Spain respectively.
Bold values indicate Granger-causality.
Bold italic values indicate new Granger-causality not detected in the whole sample.

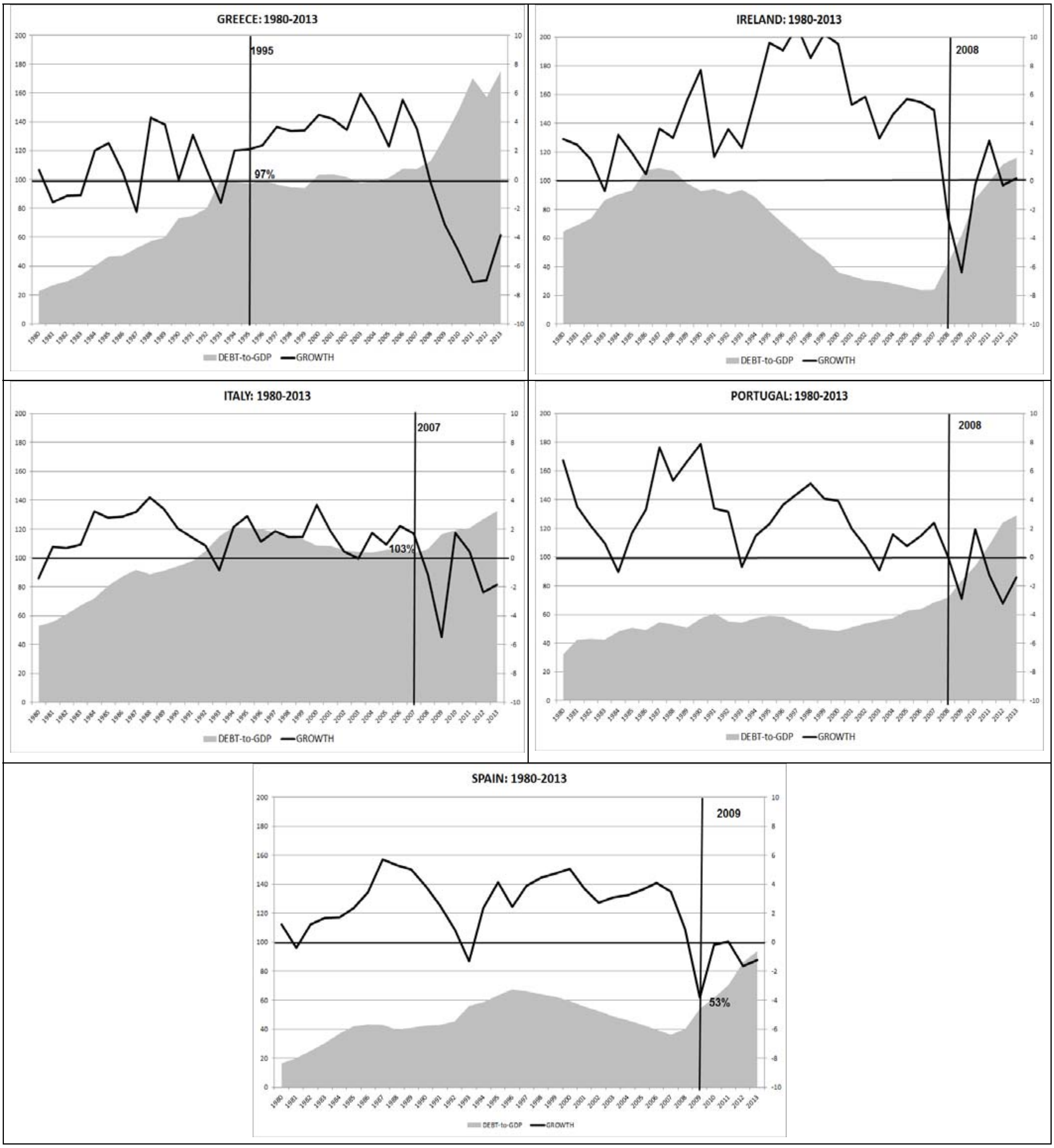
Figure 1. Debt-to-GDP, Growth and breakpoint for the relationship running from Δdebt to growth.

Figure 1a. EMU Central Countries: 1980-2013



Notes: Debt-to-GDP ratios (left scale) and real GDP growth (right scale).

Figure 1b. EMU Peripheral Countries: 1980-2013



Notes: Debt-to-GDP ratios (left scale) and real GDP growth (right scale).