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9 April 2011

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Abstract

We assess the issue of fiscal sustainability in the selected EU countries. Our sample includes those showing the highest government debts, which are nowadays known under the somewhat degrading acronym – PIIGGS (Portugal, Ireland, Italy, Greece, Great Britain and Spain). Assuming the so-called present value borrowing constraint, stationarity of debts presents a sufficient condition for fiscal sustainability. Utilizing various standard panel unit root tests and the test by Im et al. (2010), we examine this condition on quarterly debt-to-GDP ratios over the period 2000 to 2010. Results provide evidence, that when trend breaks in the series are incorporated, not all of these countries exhibit non-stationarity behavior of their debt-to-GDP ratios.

Key words: Fiscal sustainability, Government debt, Panel unit-root tests
JEL classification: E62, C23, H62, H63

Acknowledgement

This paper was supported by grant from Slovak Grant Agency VEGA No. 1/0826/11 and VEGA No. 1/0641/10.

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1. Introduction

Fiscal sustainability is a key issue for policy makers within the European Monetary Union (EMU henceforth) framework. This topic has been studied quite extensively and is interesting for several reasons: (1) individual fiscal discipline of each EMU member state is relevant to establish the common monetary policy; (2) accumulation of debt in conjunction with subsequent budgetary deficit may invoke an increase of long-term interest rates, which is unfavorable in integrated financial markets where eventually the sovereign debt will be placed scarcely; (3) in compliance with the European Union Treaties, member states adopting the euro have to fulfill the Maastricht convergence criteria (namely in the fiscal area – keeping the level of government debt under the 60 % of GDP of reference value and limiting the deficit at most of 3 % of GDP) and then the Stability and Growth Pact assures, that the fiscal discipline will be monitored henceforward.

We have mentioned three major fields of interest for which the government debt sustainability is relevant. The fourth one stemmed during the recent financial and economic crisis which has resulted in the European sovereign debt crisis, the so-called "2010 Euro Crisis". Countries showing the highest deficits and debts are nowadays known under the degrading acronym PIIGGS, which stands for Portugal, Ireland, Italy, Greece, Great Britain and Spain. Among other implications, it is unprecedented that the average rate of return from junk bonds is lower (one year yield is 6.372 % as of 11.3.2011 measured by Merrill Lynch High Yield 100 index), than the yield from the government bonds of some EU countries (latest 5 year Portugal emission yields 7.126 %).

The aim of this paper is to assess the government debt sustainability of PIIGGS countries under the present value borrowing constraint. The recent drop in the output of economies, with the long-term increase of debts makes this topic of high interest to policy makers and investors alike. Analysis is conducted by applying standard unit root tests and the test by Im et al. (2010) which allows for a cross-sectional dependence of time series within the panel and break occurrence in both level and trend.

The paper is organized as follows. Section 2 describes some theoretical background of the present value borrowing constraint and includes a brief survey of the empirical literature. Section 3 presents the dataset and methodology of panel unit root tests. Section 4 provides empirical findings and Section 5 concludes.

2. Theoretical background and empirical literature overview

A sustainability of public finance is usually presented in the form of present value borrowing constraint (PVBC henceforth). In nominal terms, government budget constraint for one country at time *t* can be written as[1]:

$$G_t + (1 + r_t)D_{t-1} = R_t + D_t \qquad (1)$$

where *G* is the government expenditure, *R* is the government revenue, *D* is the government debt and *r* is the interest rate payable on *D*. In the absence of money finance, the eventual budget deficit $G_t - R_t + r_t D_{t-1}$ must be financed by an increase of debt $D_t - D_{t-1}$. Equation *(1)* can be recursively solved for the subsequent periods, whereby inter-temporal budget constraint is formed as:

$$D_t = \sum_{s=1}^{\infty} \frac{R_{t+s} - G_{t+s}}{\prod_{j=1}^{s} (1 + r_{t+j})} + \lim_{s \to \infty} \prod_{j=1}^{s} \frac{D_{t+s}}{(1 + r_{t+j})} \qquad (2)$$

We can consider a fiscal policy as sustainable, when the second term from the right-hand side of Equation *(2)* goes to zero in infinity. The motivation behind stationarity testing lies in the fact, that a stationary D_{t+s} (around a constant, or a deterministic trend) implies slower then exponential growth (in the denominator of the right fraction), which would be needed for the debt to be unsustainable.

The PVBC in Equation *(1)* can also be rewritten using all variables as a percentage of GDP:

$$\frac{G_t}{GDP_t} + \frac{(1 + r_t)D_{t-1}}{(1 + g_t)GDP_{t-1}} = \frac{R_t}{GDP_t} + \frac{D_t}{GDP_t} \qquad (3)$$

where the growth rate of GDP is denoted as *g*. If r_t is assumed to be stationary (with mean *r*) and *g* is constant, the PVBC is given by:

$$\delta_{t-1} = \sum_{s=0}^{\infty} \left(\frac{1+g}{1+r} \right)^{s+1} (\rho_{t+s} - \varepsilon_{t+s}) + \lim_{s \to \infty} \delta_{t+s} \left(\frac{1+g}{1+r} \right)^{s+1} \qquad (4)$$

where $\delta_t = D_t / GDP_t$; $\varepsilon_t = [G_t + (r_t - r)D_{t-1}] / GDP_t$ and $\rho_t = R_t / GDP_t$. If the last term in Equation *(4)* becomes zero ($r > g$), the fiscal policy will be sustainable and growth of public debt will not become an explosive process. This yields the familiar result that fiscal policy will be sustainable if the present value of the future stream of primary surpluses, as a

[1] Following Wilcox (1989), Llorca – Redzepagic (2008) and Afonso – Rault (2010).

percentage of GDP, matches the ‘‘inherited’’ stock of government debt (Afonso – Rault, 2010)[2].

To analyze the fiscal sustainability, two general approaches are applied. The first is to test the government debt for a presence of a unit root[3] and the second is to conduct a cointegration analysis between government revenues (*R*) and expenditures (*G*)[4]. When the analyses are conducted on the individual samples of each country, some methodological issues related to the short length of the time series could arise. Such studies therefore often provide mixed results (see, Wilcox, 1989; Uctum – Wickens, 2000 or Bergman, 2001)[5]. Due to insufficient length of macroeconomic data, several recent empirical papers applied more powerful techniques in a panel framework.

Holmes et al. (2010) analyzed annual budget deficits as a percent of GDP over the sample period 1971 – 2006 for Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Spain, Sweden and the United Kingdom. Using Hadri – Rao (2008) test which allows for cross-sectional dependence and for endogenously detected structural breaks, they conclude that EU countries exhibit fiscal stationarity over the full period, even in the subsamples 1971 – 1990 and 1991 – 2006 (pre- and post- Maastricht Treaty). Evidence against the non-stationarity is considered here as support for the strong form of fiscal sustainability insofar as satisfying the PVBC.

Similar conclusions are provided by Afonso – Rault (2010) for the EU-15 countries during the period 1970 – 2006. They found the first difference of stock of government debt to be stationary, using individual LM unit root tests[6] (Schmidt – Phillips, 1992; Lee – Strazicich, 2003), panel LM unit root test (Im – Lee, 2001), panel unit root tests with cross-sectional independence (Levin et al., 2002; Im et al., 2003), and panel unit root tests allowing for cross-sectional dependence (Choi, 2006; Moon – Perron, 2004).

In the case of the cointegration tests it is assumed, that *R* and *G* are both non-stationary while their first differences are stationary. Nevertheless, if one variable is I(0) in levels and the second one is I(1), sustainability is still possible but not observed by

[2] For more technical details of PVBC see, e.g., Greiner et al. (2007).

[3] For the sake of brevity, we will not distinguish between unit root and stationarity tests (null hypothesis of applied tests will be clearly stated in the results section).

[4] According to our knowledge, both procedures to empirical testing of fiscal sustainability were primary applied by Hamilton – Flawin (1986). Some critics of both approaches regarding to PVBC can be found in Bohn (2007).

[5] Even though some frequently cited papers (e.g., Ahmed – Rogers, 1995) provide clear results from univariate unit root tests to support the fiscal sustainability.

[6] Some results were mixed in this case. Note that Lee – Strazicich (2003) test allows for two structural breaks. Afonso – Rault (2010) used this univariate test to incoroporate series specific breaks into the LM panel test.

cointegration analysis (for details see Afonso – Rault, 2008). Even when *G* and *R* are integrated in different orders, it cannot be clearly stated that there is a sustainability problem (e.g. revenues are systematically above expenditures and budgetary surplus is executed). For this group of empirical studies see, for example, Hamilton – Flawin (1986), Ahmed – Rogers (1995), Prohl – Schneider (2006) or Westerlund – Prohl (2010).

As Afonso – Rault (2010) pointed out; stationarity is a sufficient but not necessary condition for fiscal sustainability. A necessary condition is the existence of a long-run relationship (cointegration) between *R* and *G*. Nevertheless, in this paper we will focus on the first approach of testing the fiscal sustainability, i.e. stationarity testing of government debt while several panel unit root test will be applied. It is worth to mention that our sample of data is much more recent and includes the EU debt crisis.

3. Data description and methodology

To verify the fiscal sustainability of selected EU countries (Portugal, Ireland, Italy, Greece, Great Britain and Spain) we perform various panel unit root tests using debt-to-GDP ratios over the period 2000Q4 – 2010Q3. Some tests require balanced panels, and that is why the time span of our dataset is limited by availability of data which are obtained from the public source – Eurostat. Descriptive statistics and normality tests are presented in the following table.

Table 1: Descriptive statistics and normality test (debt-to-GDP ratios)

	Portugal	Ireland	Italy	Greece	Great Britain	Spain
Obs.	40	40	40	40	40	40
Mean	60.848	37.988	109.350	107.960	45.430	47.348
Median	61.700	32.350	108.700	104.500	41.900	47.550
Max.	84.200	90.500	119.600	140.100	75.100	59.300
Min.	48.200	24.600	103.600	97.200	36.700	35.300
Std. Dev.	8.802	16.258	4.016	10.875	10.723	7.229
Skewness	0.797	1.891	1.086	1.611	1.675	-0.058
Kurtosis	3.266	5.576	3.603	4.620	4.546	1.769
Anderson-Darling	0.843	4.721	1.456	3.416	4.450	0.523
p-value	0.027	<0.005	<0.005	<0.005	<0.005	0.172

Source: Eurostat

Note: For a visualization of our dataset see Figure 1. Although one of the most used normality tests in economics seems to be the Jarque – Berra tests, we have used rather Anderson-Darling test. For example Yazici – Yolacan (2007) had conducted an extensive study of normality tests, where Anderson-Darling on n = 40 performed rather well. The empirical alfa was 0.051 which is close enough to the nominal value. The power against Beta(2;2), Gamma(2;1), Log-normal, Weibull(2,1) and t(4) were 1.000, 0.959, 0.253, 0.999, 0.971.

It is obvious, that maximum values of debt-to-GDP ratios are affected by recent financial crisis, since the debts tend to increase and moreover, GDPs were decreasing. It can be seen from basic statistics, that Italy and Greece did not kept the level of government debts under the 60 % of GDP within the whole analyzed period. According to above mentioned empirical literature, the stock of debt itself does not provide information about the fiscal sustainability. We can make several conclusions only from descriptive statistics, but our point of interest is whether the unit root tests are able to provide clear results on the matter of debt sustainability of PIIGSS countries.

For this purpose some empirical works utilized univariate unit root tests, which have notoriously low power. This is mostly the case for short span data series and data series with sum of the true autoregressive parameters near, but less than one. If one of the series in the panel framework is stationary, panel tests have higher probability of rejecting the joint null hypothesis of non-stationarity (see, Taylor – Sarno, 1998). Therefore whenever possible, it is currently a standard approach to complement stationarity analysis with panel unit root tests. However, in some tests the joint null hypothesis of a (heterogeneous or homogenous) unit root is not always meaningful. According to alternative hypotheses which claim that some time series are stationary, this only tells the researcher that at least one panel member is stationary, with no information about how many series, or which ones, are stationary (see, Breuer et al., 2002, p. 527).

In the last two decades the research on panel tests has grown rapidly. The distinction between panel tests are made on whether they: assume cross-sectional independence/dependence of the series, assume a common or time series specific data generating process, or allow for structural breaks in the series of the panel. As our samples have only small number of observations, we have not analyzed and assessed univariate stationarity tests. In this paper, we have selected standard, in recent years the probably most widely used tests (Levin et al. (2002) – LLC test; Im et al. (2003) – ISP test; Breitung – Das (2005), Maddala and Wu (1999) – Fisher type χ^2 test[7], Hadri (2000) – LM test).

For the purposes of our analysis, we also employ a test proposed by Im et al. (2005, 2010), which allows for cross-sectional dependence and structural breaks in the level and trend of the series. As in the previous cases, we again deal with a panel unit root test, which has a potentially greater power when compared to the basic univariate tests. As the choice of

[7] Choi (2001) proposed three other Fisher type tests (Z, L^* and P_m) in which the power of all the tests increases as N grows, but the size and the power of all three tests decrease when a linear time trend term is included in the model. For these reasons we decided to use the Maddala – Wu Fisher type test.

available methodologies is large, our choice was motivated by several attractive properties of this particular test.

First, it allows for the presence of breaks both in level and trend. These breaks are not determined endogenously, but have to be identified prior to the unit root testing. This necessity to separate the testing of the unit root hypothesis and break identification may be beneficial, as discussed by Kim – Perron (2009). Particularly, it is possible to avoid the problems with the formulation of the null and alternative hypothesis, such as in the case of the Zivot – Adrews test (1992).

Second, this particular test is based on a statistic with an asymptotic distribution that is free of nuisance parameters, related to the position and the size of the breaks. This property has the advantage that the critical values do not have to be recalculated for the specific breaks found in the data.

In the description of the test used on our sample, we follow the notation and description given by Im et al. (2010; ILT henceforth). We consider a balanced panel dataset with N cross-sectional units and T observations per unit. The vector:

$$Z_{it} = (1, t, DT_{it}) \tag{5}$$

describes the deterministic components in the modeled series, where i denotes the cross-sectional unit ($i=1,2,...,N$) and t is a time variable ($t=1,2,...,T$). The variable DT_{it} describing the break in trend is defined as:

$$DT_{it} = \begin{cases} 0; & t < t_{b_i} \\ t - t_{b_i}; & t \geq t_{b_i} \end{cases} \tag{6}$$

where t_{b_i} is the time index for the occurrence of a break in trend within the series for cross-sectional unit i. We identify one break for all series in our analysis (see Appendix for detail results). It was possible to allow for breaks in the level of the series, however judging from the data, we considered trend shifts as adequate.

We then follow ILT, by calculating a detrended series

$$\tilde{y}_{it} = y_{it} - \tilde{\psi} - Z_{it}\tilde{\delta}' \tag{7}$$

where the parameters $\tilde{\delta}'$ are obtained from an auxilary regression:

$$\Delta y_{it} = \Delta Z_{it}\delta' + u_{it} \tag{8}$$

It can be seen, that such a procedure doesn't allow for the estimation for the constant in Z_{it}. However, the parameter $\tilde{\psi}$ is chosen so that $\tilde{y}_{i1} = 0$ holds, and thus accounts for the effect of the constant. ILT suggest that the detrended values should not be used directly, as the distribution of the test statistic would still not be nuissance-free. Instead, they use:

$$\tilde{y}_{it}^{*} = \begin{cases} \dfrac{T}{t_{b_i}} \tilde{y}_{it}; & t \le t_{b_i} \\ \dfrac{T}{T - t_{b_i}} \tilde{y}_{it}; & t > t_{b_i} \end{cases} \tag{9}$$

Using this series, ILT formulate a test equation augmented by cross-sectional averages of the lagged levels and first differences ($\bar{y}_{t-1}^{*}$ and $\Delta\bar{y}_{t}^{*}$) to account for cross-correlation:

$$\Delta y_{it} = \Delta Z_{it}\delta' + \phi_i \tilde{y}_{i,t-1}^{*} + g\bar{y}_{t-1}^{*} + h\Delta\bar{y}_{t}^{*} + \sum_{j=1}^{p} g_{ij}\bar{y}_{t-j}^{*} + \sum_{j=1}^{p} d_{ij}\Delta\tilde{y}_{i,t-j} + u_{it} \tag{10}$$

The choice of lag length p was conducted by examining the Schwartz information criterion (BIC) for each series. As the critical values reported by ILT for the final LM statistic assume a common lag choice, we have used lag order of one, the optimal order for the majority of the series used. The *t*-statistic (called $\tilde{\tau}_i^{*}$) for the null hypothesis of $\phi_i = 0$ in each equation can be used to calculate the *t*-bar statistic:

$$\bar{t} = \sum_{i=1}^{N} \tilde{\tau}_i^{*} \tag{11}$$

which in turn can be used to establish the statistic of the LM test having standard normal distribution:

$$LM(\tilde{\tau}^{*}) = \frac{\sqrt{N}\left(\bar{t} - E(\bar{t})\right)}{\sqrt{\operatorname{var}(\bar{t})}} \tag{12}$$

The expected value and variance of the *t*-bar are tabulated by ILT in the Table 3 placed in the appendix of their paper.

4. Empirical findings

We start our analysis with basic, well-known panel unit root tests. Debt-to-GDP ratios of the selected countries are in their levels non-stationary according to results from LLC, IPS, Fisher type ADF and Breitung – Das tests. The null of unit root in all series is rejected in the case of Fisher type PP test, which means that some of the series in our panel may be stationary (but due to inexact formulation of alternative hypothesis it cannot be resolved which one). In the same matter, Hadri LM test with the null of stationarity claims, that some time series contain a unit root. Unambiguous results are obtained when the panel analysis was performed on the first differences of the debt-to-GDP ratios. Applying such transformation makes most economic time series stationary (i.e. integrated of order one), which is as well true in our case. The more detailed results are provided in the following table.

Table 2: Results from panel unit root/stationarity tests

	Level		First differences	
	statistics	*p*-value	statistics	*p*-value
LLC	3.912	1.000	-11.514	0.000
IPS	4.235	1.000	-12.920	0.000
Fisher type (ADF)	18.182	0.110	104.162	0.000
Fisher type (PP)	25.380	0.013	262.754	0.000
Breitung – Das*	4.081	1.000	-4.966	0.000
Hadri LM	23.310	0.000	-0.072	0.529

*Note: trend is included in all tests; * allows for cross-sectional dependence across panel;*
LLC test H_0: all time series have a unit root; H_1: all time series are stationary.
IPS test H_0: all time series have a unit root; H_1: some time series are stationary
Maddala – Wu Fisher type tests H_0: all time series have a unit root; H_1: some time series are stationary
Breitung – Das test H_0: all time series have a unit root; H_1: all time series are stationary
Hadri LM test H_0: all time series are stationary; H_1: some time series have a unit root

Since we are dealing with the so-called 2010 Euro Crisis, it is reasonable to assume an occurrence of trend breaks in analyzed series. Over the last two years, due to recent financial and economic crisis, GDPs were decreasing and government debts tend to increase. Both these tendencies have potentially resulted in much higher debt-to-GDP ratios. In the light of these propositions, we continue our analysis by identifying possible breaks in the series.

Rather than testing for the true number of trend breaks, the small sample size of our data, visual inspection and the commonly known facts about the economic crisis in recent years have led us to assume only one break, $m = 1$. For the date break estimation technique we have followed the commonly used approach of multiple linear regression model, where we have searched for data partitions where the residual sum of squares (RSS) was at a global minimum. For further details of applications see Bai – Perron (1998, 2003) or Zeileis – Kleiber (2005). We have searched for the break date T_j in each of the series by minimizing the residual sum of squares in the following model: $y_t = \delta_{0,j} + t\delta_{1,j} + u_t$, where j = 1, 2 and $t = T_{j-1}+1, \ldots, T_j$. The trimming parameter was set only to $h = 4$ observations from the beginning and end of the series. The one break estimations pointed to the beginning of the recent crisis, more precisely at the end of the year 2007 (Portugal, Ireland) and during the year 2008 (Italy, Greece, Great Britain, Spain). We also computed 95 % confidence intervals for these breaks, which are quite narrow and confirm their location. Break dates with corresponding regime specific estimates of the coefficients are reported in the Appendix. It is also interesting to see, how the size of trend coefficients increased after the identified breaks. The most notable is the case of Ireland, where the coefficient has changed from negative value before the break to highest positive value within the whole sample. These breaks are apparent from graphical visualization of all debt-to-GDP ratios which is presented in the Figure 1.

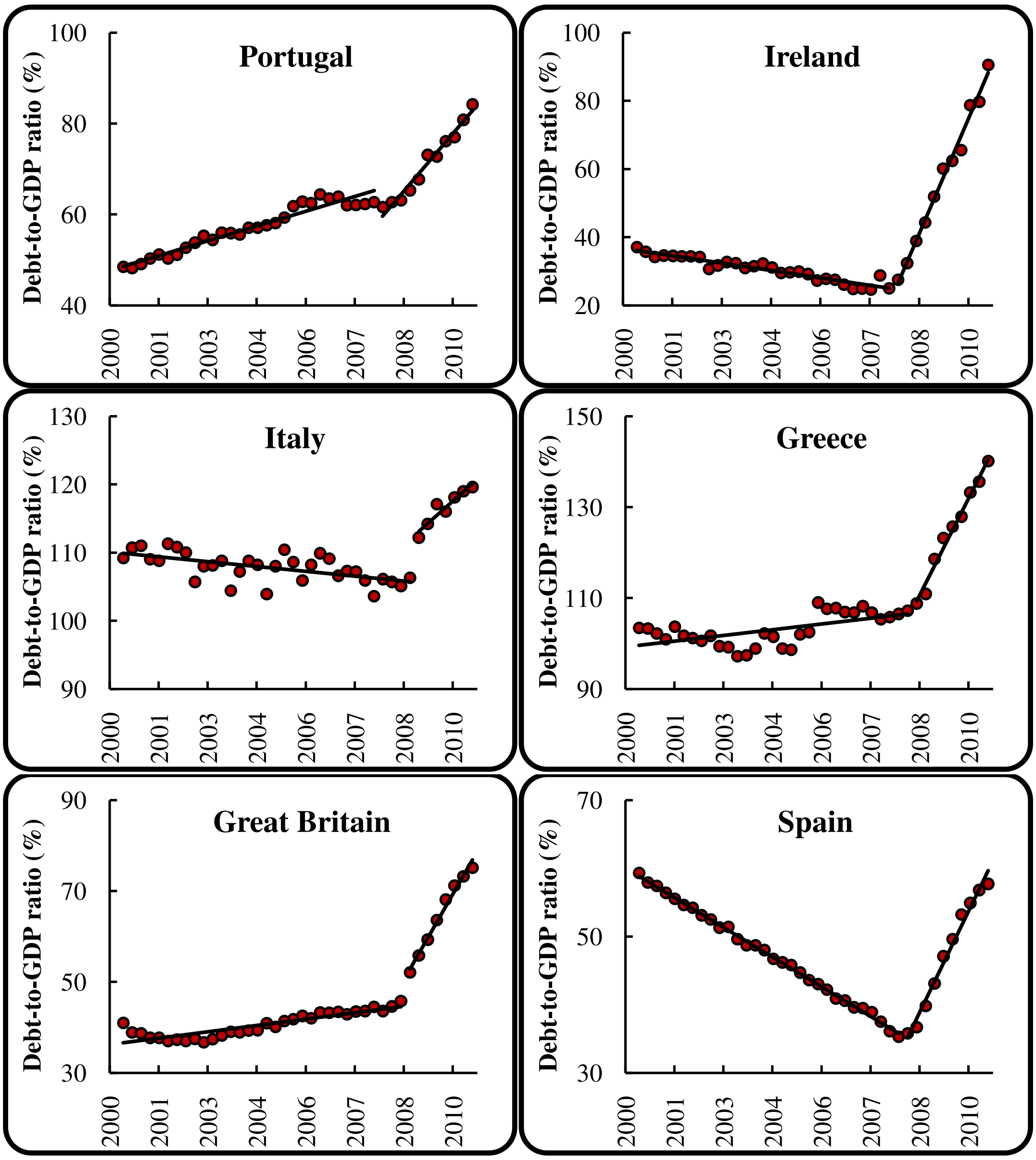

Figure 1: Debt-to-GDP ratios and detected breaks in trend

Note: The breaks in trends correspond to those obtained by the minimization of RSS, see Appendix for further details.

As the occurrence of the trend breaks in the data was obvious, we have decided to employ one of the more recent panel unit root tests of Im et al. (2010) which takes into account cross-sectional dependence within the panel and breaks in level and trend as well. The results of the testing procedure on our dataset, together with the parameters tabulated by ILT are shown in the following table.

Table 3: Results from the ILT panel unit root test

N	T	N_{ILT}	T_{ILT}	$\bar{t}$	$LM(\tilde{\tau}^{*})$	p-value
6	40	10	32	-1.616	7.962	0.000
6	40	10	50	-1.616	8.891	0.000

The null hypothesis of the ILT test is that all series in the panel contain unit roots, with the alternative that some of the series are stationary. Our results indicate the rejection of the null hypothesis, that is, we are able to reject the non-stationarity assumption for at least some PIIGGS countries. Thus the sustainability of government debts within the PVBC framework remains partly unresolved for the PIIGGS countries despite of applying one of the latest panel unit root test.

5. Conclusion

In the title of this paper we put a simple question, regarding to ability of standard econometric techniques to capture the recent European sovereign debt crisis. Under the so-called present value borrowing constraint, stationarity of debts is a sufficient condition for fiscal sustainability. Consequently, to resolve this question some standard panel unit root tests has been applied along with one of the latest test proposed by Im et al. (2010).

Standard tests provided evidence of non-stationarity of debt-to-GDP ratios (in their levels) with an exception of Maddala – Wu (1999) Fisher type PP test, which rejected the null in the favor of stationarity of some series. After one trend break has been identified in each series and included to the ILT test, the results suggested that not all of the PIIGGS countries have non-stationary debt-to-GDP ratios (i.e. providing evidence of debt sustainability). Nevertheless, it is still questionable whether the trend stationarity is still a sufficient condition for fiscal sustainability. Perhaps a sequential employment of the LM test could shed light on the issue, which countries contribute to this uncertainty of the results. Another approach might lie in supplementing of our analysis with a stationarity panel test which allows for structural breaks, like Hadri – Rao (2008) or with the SUR-ADF approach, advocated by Breuer et al. (2002). At the end we would like to emphasize, that it was our intention to avoid any personal opinions on the subject of recent debt crisis and our focus was strictly limited to the potential of the quantitative approach.

Appendix

Identified break dates with 95 % confidence intervals

	Break date P(2007:Q3 ≤ 2007:Q4 ≤ 2008:Q1) = 95 %			
Portugal	coefficient	variance	coefficient	variance
Intercept	47.98	8.34	-9.42	18.57
Time	0.59	0.10	2.31	0.01
	Break date P(2007:Q3 ≤ 2007:Q4 ≤ 2008:Q1) = 95 %			
Ireland	coefficient	variance	coefficient	variance
Intercept	36.58	0.16	-157.80	2.49
Time	-0.40	0.00	6.15	0.00
	Break date P(2008:Q3 ≤ 2008:Q4 ≤ 2009:Q1) = 95 %			
Italy	coefficient	variance	coefficient	variance
Intercept	110.02	0.35	73.26	11.91
Time	-0.13	0.00	1.17	0.01
	Break date P(2008:Q1 ≤ 2008:Q2 ≤ 2008:Q3) = 95 %			
Greece	coefficient	variance	coefficient	variance
Intercept	99.35	2.11	-15.03	41.87
Time	0.23	0.01	3.89	0.03
	Break date P(2008:Q2 ≤ 2008:Q3 ≤ 2008:Q4) = 95 %			
Great Britain	coefficient	variance	coefficient	variance
Intercept	36.35	0.28	-60.43	99.79
Time	0.26	0.00	3.43	0.07
	Break date P(2007:Q4 ≤ 2008:Q1 ≤ 2008:Q2) = 95 %			
Spain	coefficient	variance	coefficient	variance
Intercept	59.63	0.02	-48.51	148.50
Time	-0.79	0.00	2.70	0.14

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