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Are EU budget deficits sustainable?

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Abstract

In this paper, we test for the stationarity and sustainability of European Union budget deficits over the period 1971 to 2006, using a panel of thirteen member countries. Our testing strategy addresses two key concerns with regard to unit root panel data testing, namely (i) the identification of which members-states are stationary, and (ii) the presence of cross-sectional dependence. We employ a moving block bootstrap approach to the Hadri (2000) procedure that tests the null of joint stationarity. In contrast to the existing literature, we find that the EU countries considered are characterised by fiscal sustainability over the full sample period. This conclusion also holds when analysing sub-periods based on before and after the Maastricht treaty.

JEL Classification: C33; F32; F41.

Keywords: Heterogeneous dynamic panels, fiscal sustainability, mean reversion, panel stationarity test.

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1 Introduction

For the European Union (EU), the size of the government budget deficit has assumed a particular importance in recent years with the Maastricht Treaty and Stability and Growth Pact making fiscal sustainability an explicit criterion for membership of European Monetary Union (EMU). The original Maastricht requirement, set in 1992, that governments run a budget deficit of no more than 3% of GDP as a precondition to enter EMU has, for many EU countries, implied a shift towards a more restrictive fiscal stance and with this, the possibility of adverse consequences with respect to output and unemployment. However, if the budget is out of control, economic policies at both the macro- and microeconomic levels will quickly become unsupportable, and require changes to be made. Sustainability is the criterion which is frequently used to evaluate whether fiscal policy is under control. In this context, *sustainable* fiscal policies have been judged in terms of whether the government intertemporal budget constraint (IBC) holds in present value terms (see, for example, Hamilton and Flavin 1986). The IBC is based on the equality of current debt with the sum of expected future discounted primary surpluses. *Unsustainable* policies, on the other hand, are characterized by violation of the IBC meaning that, at some time in the future, such policies will have to be changed, otherwise they will lead to the government becoming insolvent or to a collapse of the policy regime. This in turn would have serious implications for the credibility and functioning of EMU.

It is important, therefore, to view fiscal sustainability as a long-run concept. The literature on budget deficit sustainability is primarily concerned with whether government's intertemporal solvency constraint is violated. Empirical examinations of this issue have fallen into one of the following two categories. The first group of studies has examined the possibility of non-stationarity in budget deficits by conducting tests of unit roots. Evidence against the existence of unit roots has been considered as support for the strong form of budget deficit sustainability. The results of this line of research have been mixed. Studies such as Caporale (1995) and Vanhorebeek and Rompuy (1995) paint a varied picture for the EU. Regarding

the US, while Hamilton and Flavin (1986) find that the budget deficit follows a stationary stochastic process and thus is sustainable, Wilcox (1989), Trehan and Walsh [(1988), (1991)], and Kremers (1989) find that the budget deficit is non-stationary implying an unsustainable budgetary process.

The second group of studies has addressed the issue of sustainability of budget deficits by examining the long-run relationship between government revenues and expenditures using cointegration methodology. The existence of cointegration between revenues and expenditures has been considered as evidence consistent with the intertemporal budget constraint and can be regarded as the weak form of budget deficit sustainability. The results of this line of research have also been mixed. With regard to EU countries, studies such as Bravo and Silvestre (2002) and Afonso (2005) find limited evidence in favour of cointegration. For example, Bravo and Silvestre (2002) find evidence consistent with sustainable budgetary paths in the cases of Austria, France, Germany, Netherlands and the UK, but not in the cases of Belgium, Denmark, Ireland, Portugal, Italy and Finland. As to the US, Haug (1991) finds support for the existence of cointegration and the intertemporal budget constraint, whereas Hakkio and Rush (1991) question the existence of cointegration when the sample period is extended towards the end of the 1980s, arguing that deficit sustainability may not hold in the later part of their sample period. The lack of consensus on both these approaches has motivated a further line of research that finds stronger evidence in favour of stationarity, cointegration and sustainability when allowance is made for the existence of structural breaks (see, *inter alia*, Tanner and Liu 1994, Quintos 1995, Martin 2000, Cunado *et al.* 2004).

In this paper, we test for stationarity and long-run sustainability of the EU budget deficits using data for a panel comprising thirteen EU members. Since unit root tests applied to single series suffer from low power, panel unit root techniques offer a way forward in terms of enhanced test power. In recent years a number of alternative procedures have been proposed to test for the presence of unit roots in panels that combine information from the time-series dimension with that from the

cross-section dimension, such that fewer time observations are required for these tests to have power. The most commonly used unit root test applied to panels include Maddala and Wu (MW) (1999) and Im, Pesaran and Shin (IPS) (2003), which test the joint null hypothesis of a unit root against the alternative of at least one stationary series, by using the augmented Dickey–Fuller (ADF) (1979) statistic across the cross-sectional units of the panel. A recent study of EU budget sustainability by Prohl and Schneider (2006) utilises a range of panel unit root tests and finds evidence in favour of sustainability.[1] It should, however, be noted that IPS (2003, p.73) warn that due to the heterogeneous nature of the alternative hypothesis in their test, one needs to be careful when interpreting the results, because the null hypothesis that there is a unit root in each cross section may be rejected when only a fraction of the series in the panel are stationary. A further issue here is that the presence of cross-sectional dependencies can undermine the asymptotic normality of the IPS test and lead to over-rejection of the null hypothesis of joint non-stationarity.

In contrast to the existent literature on fiscal sustainability, this study examines the long-run sustainability of EU budget deficits using the Hadri (2000) test of the null hypothesis that all of the individual series are stationary (either around a mean or around a trend), against the alternative of at least a single unit root in the panel. The Hadri tests thus offer the advantage that if the null hypothesis is not rejected, there is evidence that all of the current account deficits in the panel are stationary. In addition to this, an important novel feature of our analysis is that we allow for the presence of potential cross-sectional dependencies, since failing to account for this leads to over-rejection of Hadri test statistics. More specifically, we consider a procedure based on a moving block bootstrap of the Hadri tests.

The outline of the paper is as follows. Section 2 discusses the government IBC framework that is developed and used to define fiscal sustainability. We also briefly review the Hadri approach for testing for stationarity of the budget deficit in het-

[1] The range of panel tests are augmented by a procedure advocated by Banerjee and Carrion-i-Silvestre (2006) that tests the null of joint non-stationarity with an allowance for endogenously-determined structural breaks.

erogeneous panels of data allowing for the likely case in which there is cross section dependence. Section 3 describes the data and presents the results of the empirical analysis. We find that in contrast to earlier estimates, our bootstrap approach confirms sustainability across all members of the EU sample. Section 4 concludes.

2 Testing sustainability of the budget deficit in heterogeneous panel data

The link between current fiscal policy and the outstanding government debt, which constrains future policies, is summarized, in an accounting sense, in the government's IBC. In the absence of money finance, the government budget deficit must be funded through new debt creation. The IBC may be written in nominal terms as follows

$$-D_t + r_t B_{t-1} = B_t - B_{t-1}, \tag{1}$$

where D is the budget deficit defined as government (tax) revenue minus the value of government expenditure (purchases of goods and services and transfer payments), B is the value of government debt and r is the interest rate payable on B. It is common in the literature to express the IBC in terms of ratios with respect to nominal GDP. This enables us to write

$$b_t = (1 + r_t)(1 + \pi_t + \eta_t)^{-1} b_{t-1} - d_t, \tag{2}$$

where the lower case b and d refer to debt and the budget deficit expressed as a proportion of nominal GDP, π_t is the rate of price inflation and η_t is the rate of growth of real GDP.[2] Equation (2) can be re-expressed as

$$b_t = (1 + \theta_t) b_{t-1} - d_t, \tag{3}$$

where $\theta_t = (r_t - \pi_t - \eta_t)$ is the ex post real interest rate adjusted for real output growth. Looking forward one period, this provides in *ex ante* terms

[2] Equation (2) is derived from dividing equation (1) by nominal GDP ($P_t Y_t$) and using $P_t Y_t = (1 + \pi_t)(1 + \eta_t) P_{t-1} Y_{t-1} \cong (1 + \pi_t + \eta_t) P_{t-1} Y_{t-1}$.

$$b_t = E_t \left[(1 + \theta_{t+1})^{-1} (b_{t+1} + d_{t+1}) \right]. \tag{4}$$

More generally, solving forwards yields the following representation of the IBC

$$b_t = E_t \sum\nolimits_{k=0}^{\infty} \Pi_{i=1}^{k} (1 + \theta_{t+k})^{-1} d_{t+k} + E_t \sum\nolimits_{k=0}^{\infty} \Pi_{i=1}^{k} (1 + \theta_{t+i})^{-1} b_{t+k}, \tag{5}$$

where $\Pi_{i=1}^{k} (1 + \theta_{t+k})^{-1}$ is a time-varying real discount factor, adjusted for the real GDP growth rate. Equation (5) states that outstanding debt must equal the present value of the stream of funds needed to fund the interest and principal on that debt. To rule out Ponzi-type financing schemes where the government indefinitely pays debt interest by issuing more debt, the transversality condition is imposed

$$\lim_{k \to \infty} E_t \sum\nolimits_{k=0}^{\infty} \Pi_{i=1}^{k} (1 + \theta_{t+i})^{-1} b_{t+k} = 0. \tag{6}$$

It now follows that the current debt is offset by the sum of current and expected future discounted surpluses, implying that the IBC holds in present value terms with

$$b_t = \lim_{k \to \infty} E_t \sum\nolimits_{k=0}^{\infty} \Pi_{i=1}^{k} (1 + \theta_{t+i})^{-1} d_{t+k} \tag{7}$$

The IBC equation (7) implies that almost any short-run path for the primary budget surplus is consistent with budget balance in present value terms. Moreover, there will invariably exist some future set of policies, which, if implemented, can ensure that current policies do actually satisfy the IBC. However, such policies may be a long way from being optimal, or even politically feasible. Therefore, to give the concept of sustainability some value as a practical tool for policy evaluation, we must give more structure to the framework provided by equation (7). For this purpose, we follow the approach originated by Hamilton and Flavin (1986), by assuming that current processes behind fiscal policy will remain unchanged indefinitely and evaluating whether such policy is consistent with the IBC. This approach relies on the underlying stability of past data processes. In this paper, we focus on the time-series properties of the government budget deficit.

In the econometric side of the literature, Hadri (2000) proposes an LM procedure to test the null hypothesis that all of the individual series are stationary (either around a mean or around a trend) against the alternative of at least a single unit root in the panel. The two LM tests proposed by Hadri (2000) are panel versions of the test developed by Kwiatkowski, Phillips, Schmidt and Shin (KPSS) (1992). Following Hadri (2000), consider the models:

$$y_{it} = f_{it} + \varepsilon_{it}, \tag{8}$$

and

$$y_{it} = f_{it} + \gamma_i t + \varepsilon_{it}, \tag{9}$$

where f_{it} is a random walk,

$$f_{it} = f_{it-1} + u_{it},$$

and ε_{it} and u_{it} are mutually independent normal distributions. Also, ε_{it} and u_{it} are $i.i.d$ across i and over t, with $E[\varepsilon_{it}] = 0$, $E[\varepsilon_{it}^2] = \sigma_\varepsilon^2 > 0$, $E[u_{it}] = 0$, $E[u_{it}^2] = \sigma_u^2 \geq 0$, $t = 1, ..., T$ and $i = 1, ..., N$.

Let $\hat{\varepsilon}_{it}^{\mu}$ ($\hat{\varepsilon}_{it}^{\tau}$) be the residuals from the regression of y_{it} on an intercept, for model (8) (on an intercept and a linear trend term, for model (9)). Let $\hat{\sigma}_{\varepsilon^\mu}^2$ $\left(\hat{\sigma}_{\varepsilon^\tau}^2\right)$ be a consistent estimator of the error variance from the appropriate regression, which is given by:

$$\hat{\sigma}_{\varepsilon_i^l}^2 = \frac{1}{NT}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{\varepsilon}_{it}^{l2}, \qquad l = \mu, \tau.$$

Also, let S_{it}^l be the partial sum process of the residuals,

$$S_{it}^l = \sum_{j=1}^{t}\hat{\varepsilon}_{ij}^l, \qquad l = \mu, \tau.$$

Then, the LM statistic is:

$$LM_l = \frac{1}{N}\left(\sum_{i=1}^{N}\left(\frac{1}{T^2}\sum_{t=1}^{T}\frac{S_{it}^{l2}}{\hat{\sigma}^2_{\varepsilon_i^l}}\right)\right), \qquad l = \mu, \tau.$$

It should be noted that the LM statistic is based on averaging the individual KPSS test statistics. In order to obtain a consistent estimator of $\hat{\sigma}^2_{\varepsilon_i^l}$ which is efficient in the presence of residual serial dependence, we follow Hobijn *et al.* (2004) who suggest applying the Newey and West (1994) automatic bandwidth selection procedure for the Quadratic Spectral kernel.

Finally, Hadri (2000) considers the standardised statistics:

$$Z_\mu = \frac{\sqrt{N}\left(LM_\mu - \xi_\mu\right)}{\zeta_\mu} \Rightarrow N\left(0,1\right),$$

and

$$Z_\tau = \frac{\sqrt{N}\left(LM_\tau - \xi_\tau\right)}{\zeta_\tau} \Rightarrow N\left(0,1\right).$$

The asymptotic mean and the variance of Z_μ are $\xi_\mu = \frac{1}{6}$ and $\zeta_\mu^2 = \frac{1}{45}$, respectively, while the asymptotic mean and the variance of Z_τ are $\xi_\tau = \frac{1}{15}$ and $\zeta_\tau^2 = \frac{11}{6300}$, respectively. In a subsequent paper, Hadri and Larsson (2005) find the exact formulae for the two finite-sample moments of the KPSS statistic.

The Monte Carlo experiments of Hadri (2000) demonstrate that these tests have good size properties for T and N sufficiently large. However, as noted by Giulietti *et al.* (2006), even for relatively large N and T the Hadri tests suffer from severe size distortions in the presence of cross-sectional dependence, the magnitude of which increases as the strength of the cross-sectional dependence increases. This finding is in line with the results obtained by Strauss and Yigit (2003) and Pesaran (2007) for the IPS and MW panel unit root tests. To correct the size distortion caused by cross-sectional dependence, Giulietti *et al.* (2006) apply the bootstrap method and find that the bootstrap Hadri tests are approximately correctly sized.

3 Data and empirical analysis

We examine the sustainability of the budget deficit for a panel of thirteen EU countries over the study period 1971-2006. The following thirteen countries are included in the sample: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Portugal, Spain, Sweden and the United Kingdom.[3] Annual data for each of these countries are taken from the European Commission AMECO (Annual Macroeconomic Data) database.[4] In all cases, fiscal deficits are expressed as a proportion of GDP. The data exclude seigniorage whereby countries use money finance to fund a budget deficit. This exclusion is justified on the grounds that eleven countries from the sample have proceeded towards adopting the Euro as their currency and no longer have the ability to pursue an independent monetary policy.[5] Since money finance is no longer an option for these countries, it therefore seems appropriate to judge sustainability using measures of the budget deficit that exclude money financing. A second issue with the data set concerns German unification in July 1990. For data availability reasons Germany is actually measured as West Germany for 1971-1990 and Germany for 1991-2006. A final issue concerns the Maastricht treaty which was singed on February 1992 but the negotiations were completed in 1991. To consider the possibility of differences in pre- and post-Maastricht behaviour as well as the change in the measurement of the German series, we consider two sub-samples: i) from 1971 to 1990 and ii) from 1991 to 2006 for our analysis of fiscal sustainability.

Table 1 presents the results of applying the KPSS stationarity test to the budget deficits of the countries listed above (based on the model with intercept only). As indicated earlier, the tests statistics are calculated applying the Newey and West (1994) automatic bandwidth selection procedure for the Quadratic Spectral kernel.

[3]Note that Belgium France, Germany, Italy and Luxemburg are the founding states (with Holland), Ireland and the UK joined in 1973, Greece in 1981, Spain and Portugal in 1986 and Austria, Finland and Sweden in 1995.

[4]This range of countries is dictated by the availability of consistent data with respect to the study period. For this reason, Denmark and Luxembourg are excluded from the sample.

[5]Sweden and the UK are not members of the single currency.

Focusing first on the pre-Maastricht period, the null hypothesis of stationarity is rejected for six out of the thirteen countries under consideration; for five countries rejection is at the 10% significance level, and for one more country rejection is at the 5% level. Turning to post-Maastricht period, the null hypothesis of stationarity is rejected for four countries at a level of significance of 10%. Lastly, when one considers the full sample period, stationarity is rejected for two countries at the 10% level of significance. The evidence here is mixed and does not provide a clear indication of sustainability.

Next, we apply the Hadri test to the current account deficits of the countries under consideration. The main motivation for testing stationarity in a panel of data instead of individual time series is that it has been noted that the power of the tests increases with the number of cross-sections in the panel. The results of the Hadri test are reported in the last line of Table 1. As can be seen, the results indicate that the null hypothesis that all of the series in the panel are stationary is clearly rejected when considering the pre- and post-Maastricht period, as well as when analysing the full sample period (in this latter case rejection is at the 5% significance level). However, as indicated above, an important assumption underlying the Hadri test is that of cross section independence among the individual time series in the panel. The Hadri test suffers from severe size distortions in the presence of cross section dependence. Thus, to allow for potential cross section dependence, we apply the bootstrap method to the Hadri tests by resampling the residuals from either a regression of y_i on a constant for the Z_μ test, or on a constant and a trend for the Z_τ test. As suggested by Maddala and Wu (1999, p.646), we resample the residuals with the cross-section index fixed, so that we preserve the cross-correlation structure of the error term.

With dependent data, a further refinement in the bootstrap described above can be obtained by applying the idea of bootstrapping overlapping blocks of residuals rather than the individual residuals, also known as the moving block bootstrap approach.[6] This approach requires the researcher to choose the block size, i.e. the

[6]For a discussion of the moving block bootstrap see Künsch (1989), Maddala and Kim (1998)

number of contiguous residuals to be resampled with replacement. The choice of the block size is based on the values suggested by the inspection of the correlogram of the series, which involves identifying the smallest integer after which the correlogram becomes negligible, as suggested by Künsch (1989; p.1226). In particular, the results shown in Table 2 are based on 2,000 bootstrap replications used to derive the empirical distribution of the Z_μ statistic (since we focus on the model with intercept only), for alternative block sizes of 1, 2, 3 and 4 residuals. Although for some countries the smallest integer we identified is around 2, we also allowed for larger blocks in order to ensure the robustness of the results for longer block sizes.

The results of the Hadri test using the moving block bootstrap approach are reported in Table 2. As can be seen from the table, we are unable to reject the null hypothesis of panel stationarity for all three sample periods, and independently of the block size considered. These findings provide support to the view that the budget deficits of the EU countries are sustainable in the long run. Indeed, we find that sustainability is not just restricted to the post-Maastricht era, but rather has been present across the EU during the past thirty years. Having said this, Table 2 indicates an increase in the p-values associated with accepting the null during the second sub-period. This is consistent with Maastricht era providing a stronger impetus towards sustainability than was the case previously.

4 Concluding remarks

This paper applies the Hadri (2000) tests for panel stationarity to examine evidence on budget deficits stationarity and sustainability for thirteen EU countries. In contrast to standard panel unit root tests, the Hadri tests employ the null hypothesis of joint stationarity. The standard tests are of a joint non-stationary null, the rejection of which may be attributable to the stationary behaviour of as little as one panel member. This study also addresses problems associated with cross-sectional

and Berkowitz and Kilian (2000). Details on the implementation of the moving block bootstrap can be found in these references, and so are not presented here in the interests of brevity.

dependence among panel members through pursuing a bootstrap approach to the Hadri tests.

The use of individual KPSS tests for stationarity does not provide a clear indication that budget deficits are sustainable in the long run. However, within a panel context, and after allowing for the potential effect of cross sectional dependencies, we find support of the view that the budget deficits of the EU countries are sustainable in the long run. This finding also holds after considering the possibility of differences in pre- and post-Maastricht behaviour as well as the change involved in the measurement of the German data series after unification. On that basis, member countries are likely to offer long-run compliance with the Maastricht requirements concerning fiscal discipline.

Table 1. KPSS test for mean stationarity

Countries	1971 – 1990	1991 – 2006	1971 – 2006
Austria	0.410 *	0.319	0.206
Belgium	0.254	0.399 *	0.350
Denmark	0.162	0.449 *	0.135
Finland	0.162	0.318	0.225
France	0.445 *	0.225	0.479 **
Germany	0.262	0.127	0.097
Greece	0.515 **	0.362 *	0.244
Ireland	0.165	0.297	0.495 **
Italy	0.478 *	0.320	0.291
Netherlands	0.417 *	0.292	0.247
Spain	0.360 *	0.396 *	0.168
Sweden	0.173	0.332	0.120
United Kingdom	0.199	0.197	0.112
Hadri test	3.428 [0.000]	3.496 [0.000]	1.815 [0.035]

For the individual KPSS tests the finite sample critical values are based on the response surfaces in Sephton (1995). * and ** indicate 10 and 5 per cent levels of significance, respectively. The tests statistics are calculated applying the Newey and West (1994) automatic bandwidth selection procedure for the Quadratic Spectral kernel. The p-value of the Hadri test appears in [ ].

Table 2. The bootstrap Hadri test for the budget deficit

Sample period	Z_μ	Block size	p-value
1971 – 1990	3.428	1	0.580
		2	0.465
		3	0.393
		4	0.335
1991 – 2006	3.496	1	0.649
		2	0.602
		3	0.483
		4	0.492
1971 – 2006	1.815	1	0.373
		2	0.337
		3	0.312
		4	0.314

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