Fiscal deficit sustainability of the Spanish regions[*]

Autores y e-mail de la persona de contacto:

Josep Lluís Carrion-i-Silvestre[†], carrion@ub.edu

Departamento:

Econometria, Estadística i Economia Espanyola

Universidad:

Universitat de Barcelona

Área Temática: *(indicar el área temática en la que se inscribe el contenido de la comunicación)*

Finanzas públicas

Resumen: *(máximo 300 palabras)*

The fiscal deficit of the Spanish Autonomous Communities (AC) is investigated using non-stationary panel data analysis. We consider the two main approaches in the literature, first, assessing whether there is a long-run relationship between revenues and expenditures of the ACs and, second, focusing on the use of fiscal rules. The paper shows that it is possible to relate these approaches in a unified framework.

Palabras Clave: *(máximo 6 palabras)*

Fiscal deficit, fiscal rules, panel data, cross-section dependence, common factors

Clasificación JEL:
E62, H62, C12, C22

---

[*] The author acknowledges the financial support from the IEA2011 grant from the Institut d'Estudis Autonòmics, Department of Institutional Relations and Participation of the Government of Catalonia, and from the Spanish Ministerio de Ciencia e Innovación under grant ECO2011-30260-C03-03.

[†] AQR-IREA research group. Department of Econometrics, Statistics and Spanish Economy, University of Barcelona, Av. Diagonal, 690. 08034 Barcelona. Tel: 934024598 Fax: 934021821, Email: carrion@ub.edu

1. Introduction

The sustainability of government fiscal policy is a major issue especially in the current context where the developed economies are facing the effects of the global crisis. Efforts to contain public spending and streamlining the provision of public goods is an objective of the present governments trying to reactivate the economies in an environment where there is a difficulty in finding funding and liquidity. Borrowers monitor governments accounts when deciding where to locate their investment and loans. In this scenario, Spain is a case of relevant interest, given the adjustment procedures that have been implemented in order to reduce the level of debt and the pressure of the fiscal deficit of the Spanish economy.

The interest is also given by the fact that since the beginning of the democratic period in 1978 Spain started a process of competences transference towards the Spanish regions (Autonomous Communities, AC), which involves transference of some taxes and the provision of public services – security, health and education, essentially. This fiscal decentralized situation has lead the central government to monitor the ACs when trying to reduce the excessive deficit and debt levels of the Spanish economy.

The aim of this paper is to analyze the sustainability of the fiscal policy of the ACs as a whole using the two approaches that have been mainly adopted in the literature. The concept of sustainability of fiscal policy implies the fulfillment of the so-called intertemporal budget constraint, which succinctly states that the current level of debt in an economy should equal the present value of its future fiscal surpluses. If this condition is met, economies cannot be indefinitely issuing debt to cover fiscal deficits as the markets would observe a risk of bankruptcy. In order to test whether this condition is satisfied, we will analyze two different relationships. First, we use panel data cointegration techniques to assess whether exists a long-run relationship between revenues and expenditures. Second, we investigate if the fiscal rule that relates the fiscal primary surplus and debt levels holds for the Spanish ACs.

The information available to conduct the study covers the period 1984-2009, thus defining series on fiscal variables of revenues, expenditures and debt relatively in a relative short time period.[3] This suggests that the analysis of the fiscal deficit sustainability should be based on the use of panel data techniques to combine the information of both temporal and cross-section dimensions when conducting statistical inference. The analysis should also take into account that the fiscal variables that we are using show a high degree of persistence, i.e., they can be I(1) non-stationary variables. Complied with this, we should apply econometric techniques that consider this feature, if meaningful conclusions are to be obtained. To the best of our knowledge, this approach has not been used to study the case of the Spanish fiscal deficit sustainability at the AC level.

Throughout the paper we discuss the different alternatives that exist in the literature when specifying models that will allow assessing the sustainability of fiscal deficit, and the (necessary and sufficient) conditions that must be checked. As discussed below, there are methodological positions that may seem contradictory, although the present paper shows the connecting links among them.

The paper is structured as follows. Section 2 provides a selected review of the literature on fiscal sustainability, indicating which are the seminal works of the different approaches. It also discusses empirical research that exists in a sub-central government level, noting that, in general, the evidence is scarce. Section 3 develops the arithmetic of debt and its relation to the fiscal deficit, while explaining the conditions of sustainability. Section 4 presents the so-called fiscal rules as an alternative way to assess fiscal sustainability. Section 5 details the database used in the paper. Section 6 details the econometric methodology and the results of its application. Finally, Section 7 concludes with some general comments.

2. Review of the literature

The literature on fiscal deficit sustainability has experienced a major breakthrough since the early nineties, attributable, to some extent, to the emergence of techniques for non-stationary time series analysis. The seminal papers of Hamilton and Flavin (1986), Hakkio and Rush (1991) and Trehan and Walsh (1988, 1991) showed that time series

[3]It should be bear in mind that the Spanish ACs territorial organization was implemented in 1984 so there is no previous information concerning this level of government.

cointegration analysis could configure a valid strategy to assess the sustainability of governments fiscal policy. The approach is based on testing the order of integration of the variables that reflect the fiscal practice of the governments – fiscal deficit and debt – as a way to establish whether the intertemporal budget constraint is satisfied ensuring a sustainable fiscal policy in the long-run.

The contributions in the literature can be broadly classified in two different approaches. First, we have the analyses based on an univariate approach, which study the order of integration of the deficit (including interest payments on debt) – see Hamilton and Flavin (1986) – or the stock of public debt – see Wilcox (1989). Second, we have the studies that base on a multivariate approach, which analyze if there is a long-run relationship between the flows of revenues and expenditures – see Trehan and Walsh (1988), Hakkio and Rush (1991), Haug (1991), Quintos (1995) and Martin (2000), among others. Aiming to reconcile both approaches, Trehan and Walsh (1991) derive sufficient conditions for fiscal sustainability, conditions that require, first, that there exists a cointegration relationship between the primary deficit and debt and, second, that the quasi-difference of the primary deficit is an I(0) stationary process.

These non-stationary techniques are appropriate to assess long-run stochastic properties of the time series, something that is inextricably linked to fiscal sustainability and the intertemporal budget constraint. These early studies mentioned above focused mainly on the U.S. fiscal sustainability. Other subsequent studies have refined the analysis by incorporating the possibility of different economic systems (structural changes) that are associated with different degrees of sustainability – see Quintos (1995), Martin (2000) and Afonso (2005), among others – and have also generalized the definition of sustainability to distinguish between strict and weak sustainability – see the discussion below. Afonso (2005) provides a quite comprehensive summary of empirical studies in the literature indicating the country – mainly U.S., but also European countries – period, variable(s) under analysis and key findings.

Bohn (1998) criticizes these analyses arguing that, in principle, any order of integration of the public debt is consistent with the fulfilment of the intertemporal budget constraint. To overcome this criticism, Bohn (1998, 2007) offers an alternative way to assess the sustainability of the public deficit, a proposal that is based on the specification of a fiscal rule that measures the reaction of the primary surplus with respect to variations in the level of debt. According to this approach, a (statistical significant) positive response of the primary surplus to changes in debt would constitute a sufficient condition for the sustainability of fiscal policy. According to Bohn (2007), the relationship between the primary deficit and debt is of economic interest, interest that goes beyond establishing whether or not there exist a cointegration relationship between the fiscal variables. However and although Bohn (1998, 2007) argument is correct, Quintos (1995) shows that the assessment of the order of integration of the public debt is still a relevant question, provided that it gives information on the degree of sustainability (strong or weak) of fiscal deficit.

The encouraging point of Bohn's (2007) criticism would be trying to find significant relationships between the primary deficit and debt. However, it should be bear in mind that the estimation and statistical analyses that he proposes require assessing the order of integration of the variables involved in the relationship, if misleading conclusions are to be avoided. It is well known that the consistence of the estimated parameters of the models that relate I(1) non-stationary variables depends on whether cointegration takes place. Consequently, as a preliminary step to estimate the models advocated by Bohn (1998, 2007) we should proceed to assess the order of integration of the fiscal variables, provided the risk of facing a spurious relationship if the variables in the model are I(1) non-stationary stochastic processes – a spurious relationship produces inconsistent estimates of the parameters and invalidates the statistical inference.

The two methodological approaches discussed so far have also been applied in a regional environment. On one hand, Mahdavi and Westerlund (2011), and Westerlund, Mahdavi and Firoozi (2011) analyze the relationship between revenues and expenditures – at the state and local governments levels for the U.S. – using panel data cointegration techniques. On the other hand, Esteller and Solé (2004) applied the methodology of Bohn (1998) to analyze the sustainability of the fiscal policy of the Spanish ACs, while Claeys, Suriñach and Ramos (2010) do so for the U.S. states and German landers. As can be seen, the empirical evidence of fiscal sustainability at the regional level is still scarce, and mostly concentrated in the U.S. economy. In the case of Spain, we are only aware of the study in Esteller and Solé (2004), which is based on the estimation of a dynamic model specification that generalizes the proposal of Bohn (1998), but without considering the non-stationarity of the variables.

The analysis that is conducted in this paper is interesting provided that it increases the empirical evidence focusing on the Spanish ACs regions where the decentralization system and fiscal sustainability is one of the hot political discussions at present times. The approach that is adopted in the paper uses procedures designed to work with non-stationary panel data, a strategy that has not been implemented yet in the case of the Spanish regions.

Finally, it should be mentioned that the estimated specifications in the papers mentioned above are heterogeneous in terms of the definition the variables involved in the analysis. There are studies that use fiscal variables in nominal terms, in real terms relative to the GDP or to the population. Bohn (2005, 2007) indicates that this issue is not important as long as the discount factor is measured adequately. In our case, the variables are used in levels and expressed in real terms.

3. The arithmetic of debt and fiscal sustainability

This section derives the algebra for an ad-hoc version of the intertemporal budget constraint (IBC) and the implicit stationarity restrictions. The government budget constraint for each period can be written as:

$$\Delta B_t = G_t - R_t = DEF_t, \qquad (1)$$

where B_t is the market value of government debt in real terms, G_t is the government spending in real terms, including interest payments, R_t represents the revenues in real terms. The deficit (DEF_t) is the difference between government revenues and expenditures of a time period, a variable that, by definition, equals the change in the debt. However, and as mentioned in Bohn (2005), while the condition set by equation (1) is given in nominal terms, changes in the real value of debt differ from the actual deficit by an inflation term. Therefore, it is important in this context to use a definition of the dynamics of the debt that is invariant to changes in scale when distinguishing between the stock of debt and the flows of revenues and expenditures.

If we denote by i_t the real interest rate[4] and assuming that this variable is I(0) stationary around a mean value i – see Hakkio and Rush (1991) – it is possible to define:

$$G_t = GE_t + i_t B_{t-1}, \tag{2}$$

where GE_t is the actual expenditure excluding interest payments, and the second term of the right side of equation (2) represents the payment of interests on the accumulated debt at the end of the previous period. Note that the debt can be expressed as:

$$B_t = (1+i)B_{t-1} + EXP_t - R_t,$$

where $EXP_t = GE_t + (i_t - i)B_{t-1}$ or, alternatively, $B_t = (1/(1+i))(R_{t+1} - EXP_{t+1}) + (1/(1+i))B_{t+1}$. Since the government is subject to the same budget constraint in t+1, t+2,…, we can intertemporally add the budgetary constraints of each period and obtain:

$$B_t = \sum_{j=0}^{\infty}\left(\frac{1}{(1+i)}\right)^{j+1}(R_{t+j+1} - EXP_{t+j+1}) + \lim_{j\to\infty}\left(\frac{1}{(1+i)}\right)^{j+1} B_{t+j+1}. \tag{3}$$

The intertemporal budget balance (or deficit sustainability or IBC) occurs if and only if the present value of government debt equals the present value of future budget surpluses,

$$B_t = \sum_{j=0}^{\infty}\left(\frac{1}{(1+i)}\right)^{j+1}(R_{t+j+1} - EXP_{t+j+1}),$$

that is, if and only if the transversality condition holds:

[4]Note that the variables could be expressed in nominal terms or as a ratio of real GDP. If the variables are in nominal terms, i_t is the nominal interest rate. If the variables are expressed in real terms, i_t is the real interest rate. Finally, if the variables are expressed as a ratio to the GDP, 1+ i_t would be the interest rate adjusted by the growth rate of the economy, which is obtained by dividing the nominal growth rate of the GDP.

$$E_t\left(\lim_{j\to\infty}\left(\frac{1}{(1+i)}\right)^{j+1}B_{t+j+1}\right)=0, \tag{4}$$

where $E_t(\cdot)$ denotes the conditional expectation on the information set available at time t. If the condition given by (4) is satisfied, then the deficit is sustainable, given that the stock of debt that remains in the hands of the economic agents will grow at a slower rate, on average, than the growth rate of the economy (approximated by the real interest rate). Therefore, this implies that the government is not financing its deficit by issuing new debt following a Ponzi scheme game. In order to implement the empirical testing of the fiscal sustainability, we can take the first difference in (3) to obtain:

$$\Delta B_t = G_t - R_t = \sum_{j=0}^{\infty}\left(\frac{1}{(1+i)}\right)^{j+1}\left(\Delta R_{t+j+1}-\Delta EXP_{t+j+1}\right)+\lim_{j\to\infty}\left(\frac{1}{(1+i)}\right)^{j+1}\Delta B_{t+j+1},$$

so that sustainability is associated with the transversality condition:

$$E_t\left(\lim_{j\to\infty}\left(\frac{1}{(1+i)}\right)^{j+1}\Delta B_{t+j+1}\right)=0. \tag{5}$$

If the condition given by (5) is satisfied, then we can conclude that there is an intertemporal budget balance (deficit sustainability) because this would imply that the government would incur in a future surplus equal, in expected value, to the market value of the debt.

As mentioned above, the empirical literature has followed two different approaches when assessing the sustainability of the fiscal deficit. One group of studies – univariate based approach – has concentrated on analyzing the stochastic properties of B_t – see Hamilton and Flavin (1986) and Wilcox (1989). In this case, fiscal deficit sustainability would require ΔB_t to be I(0) stationary, i.e., a condition that is equivalent to check whether the I(1) non-stationary vector of variables (G_t, R_t) are cointegrated with cointegration vector (1,-1)’.

A second group of studies – multivariate based approach – analyzes whether the vector of variables (G_t, R_t) generate a cointegration relationship assuming that the cointegration vector is known and equal to (1,-1)' – in this case, we end up with the approach of the first group of studies – or estimating the cointegration vector – see Trehan and Walsh (1988, 1991), Hakkio and Rush (1991), Haug (1991), Quintos (1995) and Martin (2000).

Trehan and Walsh (1991) can be thought as the first contribution that unifies both approaches. In particular, Trehan and Walsh (1991) are the first to explicitly derive the conditions for fiscal deficit sustainability in terms of a relationship between the primary deficit (the deficit excluding interest payments on debt) and debt. In addition, Hakkio and Rush (1991) implicitly point to a relationship between the deficit and the debt, although they concentrated in the cointegration relationship between the components of the primary deficit. Hakkio and Rush (1991) postulate that if the total revenues and total expenditures are I(1) non-stationary variables that define the cointegration relationship:

$$R_t = \mu + \beta G_t + u_t, \tag{6}$$

with $0 < \beta \leq 1$, then the condition that prevents a Ponzi game situation is satisfied. In this model the value of β in (6) determines the degree of sustainability. Thus, if $0 < \beta < 1$ we have weak sustainability, whereas $\beta = 1$ defines the sustainability in the strict sense (or strong sustainability). In economic terms, sustainability in the weak sense corresponds to a situation where the government reacts to the increase in public debt, but this correction is not equal to the growth of the public expenditure. In this case, an unsteady growing deficit and an increase in public debt can be observed. Consequently, Hakkio and Rush (1991) argue that a cointegration relationship between R_t and G_t would be necessary for a strict interpretation of the sustainability of the deficit. However, Quintos (1995) indicates that $0 < \beta \leq 1$ in (6) would be a necessary and sufficient condition for the fiscal deficit sustainability, and the cointegration relationship between R_t and G_t – regardless of whether or not the cointegration vector is imposed – would only be a sufficient condition for fiscal deficit sustainability. In this regard, the debt could be either I(1) or I(2) and the fiscal deficit sustainability will still hold, although the interpretation would be qualitatively different – in the event that the debt is I(1) we have strict sustainability, while if the debt is I(2) the sustainability would be weak.

Quintos (1995) makes an interesting remark by pointing out that, although $0 < \beta < 1$ constitutes a necessary and sufficient condition for the sustainability of the public deficit, this situation is not consistent with the possibility that the government might market its debt in the long-run. The fact that $0 < \beta < 1$ has important implications in terms of economic policy. If a government spends more than it raises, it will have a high risk of failure and will have to offer a higher interest rate in order to put its debt on the market.

Another interesting aspect highlighted in Quintos (1995) is the different rate at which the fiscal deficit tends towards the sustainability, which is determined by the order of integration of B_t – see Theorem 1.1 in Quintos (1995). Thus, the rate at which (5) tends to zero is higher if $B_t \sim I(1)$ than in the case where $B_t \sim I(2)$. This allows us to distinguish between strong and weak sustainability, respectively, as mentioned above.

The main weakness of the proposal in Quintos (1995) is the way in which the implementation of the testing strategy is carried out. The problem lays in the fact that, in order to apply her strategy, we need, first, a consistent estimation of the parameter β in (6) and, second, we have to test if $0 < \beta \leq 1$. In the case where the revenues and expenditures of the government are I(1) variables, estimating equation (6) can lead either to a spurious relationship – where the estimated parameters are inconsistent – or to a cointegration relationship – where the estimated parameters are (super) consistent. On the one hand, if we end up with a spurious relationship, the value of β cannot be identified by using the estimation techniques that base on the individual analysis (country-by-country or region-by-region analysis) – this is the case of the empirical application in Quintos (1995), where the sustainability of the U.S. fiscal deficit is analyzed. On the other hand, in the case where the model defines a cointegration relationship, it is possible to get a consistent estimate of β, where the statistical inference requires to apply an efficient estimation method – for example, with the procedures of the fully modified OLS (FM-OLS) of Phillips and Hansen (1990), the canonical cointegration regression (CCR) of Park (1992) or the dynamic OLS (DOLS) of Saikonnen (1991) and Stock and Watson (1993). Therefore, the main problem lies in the identification of the β parameter in the possible case of no cointegration. As discussed below, this problem can be solved if, instead of using a strategy that focuses on a single unit (a single country), we use procedures based on panel data analysis, which is the methodology used in this paper.

4. Sustainability of fiscal deficit and fiscal rules

The strategy to test for fiscal deficit sustainability that has been presented in the previous section has received some criticism in the literature giving rise to alternative approaches. In this regard, Bohn (1998) proposed to estimate a fiscal rule in order to assess the sustainability of the fiscal policy of the government. Basically, Bohn (1998) suggests checking whether there exists a corrective response by the government to increases of the public debt. The focus is set on the response from the primary surplus – non-financial revenues less non-financial expenditures (excluding interest payments on debt) – to changes in the level of the public debt. The model suggested in Bohn (1998) for the U.S. economy takes the form:

$$S_t = \rho B_t^* + \alpha Z_t + \varepsilon_t, \tag{7}$$

where the primary surplus is given by $S_t = R_t - G_t^*$, being G_t^* the government spending excluding interest payments on debt, B_t^* the level of debt in the economy at the beginning of period t – it can be approximated by the level of debt in the period t-1 – and, finally, Z_t is a vector of explanatory variables that capture the economic cycle.[5] The sufficient condition for sustainability requires $\rho > 0$ in equation (7) so that the government would be taking corrective actions – reducing the level of expenditures (excluding interest on debt) and/or increasing the tax revenues – in order to offset the changes in the level of debt. Bohn (1998) mentions that it is possible to proceed in two different ways. First, if the primary surplus and debt are I(1) non-stationary variables, one might consider the relationship:

[5]In fact, Bohn (1998) defines the primary surplus and debt at the beginning of the period divided by the GDP of the economy. This transformation has no influence on the interpretation of his model, so in order to be consistent with the definition of the variables used so far, we use variables in levels. As for the other explanatory variables (Z_t), Bohn (1998) uses the variables GVAR and YVAR in Barro (1986), which aim to capture the temporary government spending and cyclical variations of the output of the economy, respectively.

$$S_t = \rho B_t^* + v_t, \qquad (8)$$

and test for the presence of cointegration between S_t and B_t^*. If cointegration holds, that would mean that $v_t = \alpha Z_t + \varepsilon_t$ is an I(0) stationary process, so that, according to Bohn (1998), it would not be necessary to explicitly model the effect of the economic cycle on the primary surplus in order to obtain a consistent estimate of ρ. Second, if the variable primary surplus and debt are I(0) stationary variables, then we should estimate equation (7) with the inclusion of the cyclical determinants of the fiscal surplus to ensure a consistent estimate of ρ.

Beyond the specific case that is analyzed in Bohn (1998), we must think about whether there is a link between the testing strategy that has been described in the previous section and Bohn's proposal. First, one can see that in order to implement the approach in Bohn (1998), we require knowing the order of integration of the variables, something that, in itself, is the first way to check if the fiscal deficit is weak or strong sustainable – see the discussion in the previous section. Therefore, we can establish one link between the two approaches. Notwithstanding, the relationship between the two approaches goes further.

Suppose that the fiscal variables involved in the model are I(1) non-stationary variables, so that equation (8) can be expressed as:

$$R_t - G_t^* = \rho B_t^* + v_t$$

$$R_t = G_t^* + \rho B_t^* + v_t. \qquad (9)$$

If we now compare equation (9) with equation (6) we see that, apart from a constant term – which Bohn (1998) also included when estimating the model – there are, apparently, two differences. First, in (6) the total expenditure:

$$G_t = G_t^* + r_t B_{t-1}, \qquad (10)$$

is used, while in (9) we have a quite similar explanatory variable:

$$\bar{G}_t = G_t^* + \rho B_{t-1}, \tag{11}$$

and, second, in (6) no restriction is imposed on β, while (9) imposes β = 1. As can be seen, the main difference lies in the definition of the interest rate that is used, since while (10) takes into account the interest rate for each period, (11) considers an average interest rate – this is a plausible assumption provided that the model assumes that the real interest rate is an I(0) stationary variable.

The model of Bohn (1998) allows us to relate his sufficient conditions for fiscal deficit sustainability with the ones drawn from the approaches described in the previous section, which rely on cointegration analysis. Thus, Bohn (2007) indicates that, in the case where equation (9) represents a cointegration relationship, three situations may occur:

- That $\rho > r$, r being the average interest rate of the debt, a situation that would imply I(0) stationarity of the deficit and debt
- That $0 < \rho < r$, a situation that would cause a slightly explosive behavior of the deficit and debt, but with a sufficiently slow growth to satisfy the intertemporal budget constraint
- That $\rho = r$, which implies that the debt would be an I(1) non-stationary process and the deficit an I(0) stationary process, fulfilling the intertemporal budget constraint *à la* Quintos

To sum up, the model of Bohn (1998) can be seen as a special case of the approach based on the analysis of cointegration discussed in the previous section, where: (i) it is imposed that the cointegration vector is known and equal to (1, -1)’ and (ii) that the payment of debt interests is calculated using a constant interest rate – which can be defined as the average of the real interest rate. Given these features and as set forth in Quintos (1995), we would be faced with a particular definition of necessary and sufficient condition for sustainability of the public deficit, which will require to impose that $r_t = \rho$.

Finally, it should be noticed that the model given by (8) relates a flow variable (primary surplus) and a stock variable (debt), both possibly being I(1) non-stationary variables. If so, this specification is related to the concept of multicointegration proposed by Granger and Lee (1989) and applied to the analysis of fiscal sustainability in Berenguer-Rico and Carrion-i-Silvestre (2011), Escario, Gadea and Sabaté (2012), and Camarero, Carrion-i-Silvestre and Tamarit (2012), among others. In particular, the analysis of multicointegration allows to estimate in a single stage the cointegration vector that relate the flow variables – in the case of the Bohn's model it is imposed that the cointegrating vector is (1, -1)' – and the one that relates the flow and the stock variables. With these final remarks, we stress that there are close ties linking the proposals testing the fiscal deficit sustainability using cointegration analysis and the ones based on the estimation of fiscal rules, something that is explored in the reminder of this paper.

5. Data and descriptive analysis

The main source of information used in this paper is the Spanish Ministry of Economy and Finance, which gives information on consolidated revenues and expenditures, settled by chapters, for the seventeen Spanish ACs regions for the period 1984-2009. From the breakdown in which the data is available, we can proceed to obtain the non-financial revenues and expenditures of the ACs, which allows us to compute the deficit and primary surplus of the Spanish ACs.[6]

Another variable that will be used in the analysis is the debt of the Spanish ACs, which has been obtained from various issues of the Monthly Bulletin of the Bank of Spain. We also require the GDP deflator of each AC in order to express the variables in real terms. In this case, the sources of information have been the BDMORES database and the Regional Accounting of the Spanish national statistical institute (INE).

The overall debt of the ACs has experienced a sustained growth over the analyzed period. There are, however, exceptions to this behavior in some sub-periods. First, note that some ACs have experienced a reduction of debt in real terms – Andalusia (2000-2008), the Basque Country and Navarre (1996-2008) – others see a real deadlock in

[6]One could think of removing the two Spanish AC foral regions that have a funding system different from the other ACs, which give them greater autonomy in their decision of raising and spending. However, these ACs also face the same conditions as the rest of the ACs when assessing whether the fiscal policy is sustainable or not, and therefore we have decided to keep them in the sample.

debt – the case of La Rioja in the period 1991-2008, Aragon, Canary Islands, Catalonia, Galicia and Murcia in the period 1996-2008, and Madrid in 2003-2008. For the rest of the ACs the debt increase has been sustained throughout the period – we should highlight the Balearic Islands, Castilla La Mancha and Valencia.

6. Panel data integration and cointegration analyses

Macroeconomic variables are usually characterized by high persistence that makes the values that they take on a particular moment of time to depend heavily on previous values. This feature makes that many macroeconomic variables are characterized as non-stationary stochastic processes that are governed by stochastic trends, i.e., integrated processes I(d) with $d > 0$. This is a relevant issue, since the estimation of models involving non-stationary time series can lead to spurious relationships. As is well known, the parameter estimates from a spurious relationship are inconsistent and the test statistics that are usually computed to validate the estimated model can lead to think that we are facing a causal relationship with economic meaning, when in fact the variables are not related.

Previous analyses in the literature have characterized the fiscal variables involved in our model specification as I(1) non-stationary processes, so caution should be taken when estimating the parameters of the models if meaningful conclusions are to be obtained. However, by working with I(1) non-stationary stochastic processes does not necessarily mean that we have to face a spurious relationship when estimating models like the ones given by (6) or (8). Thus, it is possible that relationships among I(1) variables lead to consistent estimates of the parameters if the variables generate a cointegration relationship. Therefore, the analysis should proceed assessing, first, the order of integration of the variables involved in the model and, second, testing for the presence of cointegration if these variables are characterized as I(1) stochastic processes. If evidence of cointegration is found, the estimation of the parameters will be consistent, but the inference that can be made about them is carried out in a different way as commonly done when working in a stationary framework.

In this paper the order of integration and cointegration analyses are performed using panel data techniques. The advantage of taking into account the statistical information coming from both the temporal and the cross-section dimensions is the improvement of the statistical inference, provided that panel data unit root and cointegration test statistics are supposed to be more powerful than the ones based on the individual information. However, non-stationary panel data techniques can lead to misleading conclusions if the presence of cross-section dependence among the units of the panel data sets is not taken into account. The first generation of non-stationary panel data techniques assumed the independence among the units of the panel data sets, an assumption that, if not satisfied, will introduce a bias to conclude in favor of the stationarity of the panel data – see Banerjee, Marcellino and Osbat (2004, 2005). Although it is now a common practice to apply panel data unit root and stationarity test statistics that account for cross-section dependence, few studies test whether such dependence exist. Further, the application of these cross-section dependence test statistics can give some hints on the type of cross-section dependence that is present and, hence, how should we control for it.

6.1. Panel data cross-section dependence

In this section we compute the test statistics in Pesaran (2004, 2013), Ng (2006) and Bailey et al. (2012), which specify the null hypothesis of cross-section independence against the alternative hypothesis of cross-section dependence. Aside from whether there is evidence of dependence among the units of the panel data sets, the application of the test statistic in Ng (2006) is interesting because it provides information about the degree of dependence – in the sense that we can conclude whether dependence is pervasive. The paper in Bailey et al. (2012) also focus on measuring the strength of the cross-section dependence. We briefly discuss these test statistics.

The WCD test statistic in Pesaran (2004, 2013) is given by:

$$WCD = \sqrt{\frac{2T}{n}} \sum_{j=1}^{n} \hat{\rho}_j,$$

where $\hat{\rho}_j$, $j = 1, 2, \ldots, n$, $n = N(N-1)/2$, denotes the pair-wise Pearson's correlation coefficients of the ordinary least squares (OLS) residuals $\hat{e}_{i,t}$ obtained from the estimation of the AR(p_i). Under the null hypothesis of cross-section independence, the

WCD statistic converges to the standard normal distribution. Pesaran (2004, 2013) also mentions that the Breusch-Pagan Lagrange multiplier test statistic (WCD_{LM}) can be applied to test the null hypothesis of cross-section independence. This statistic takes the expression:

$$WCD_{LM} = \sqrt{\frac{1}{2n}\sum_{j=1}^{n}\hat{\rho}_j^2},$$

which under the null hypothesis of cross-section independence converges to the standard normal distribution. Pesaran (2013) indicates that these test statistics can be used to detect weak cross-section dependence – i.e., cross-section dependence that is not pervasive. However, large values of these statistics can be taken as an informal indication that there might be strong dependence among the units of the panel. In this regard, Bailey et al. (2012) propose a statistics to measure the degree of cross-section dependence (δ), for which confidence intervals can be computed. A value of δ around one indicates that strong dependence is present.

As for the previous statistics, Ng (2006) suggests getting rid of the autocorrelation pattern in the individual time series through the estimation of an AR model, which allows us to isolate the cross-section regression from serial correlation. Taking the estimated residuals from the AR regression equations as individual series, we compute the absolute value of Pearson's correlation coefficients $\left(\bar{\rho}_j = \left|\hat{\rho}_j\right|\right)$ for all possible pairs of individuals, $j = 1, 2, \ldots, n$, $n = N\left(N-1\right)/2$, from which the *svr* statistic to test the null of independence is specified – under the null hypothesis of independence the test statistic converges to the standard normal distribution. Ng (2006) proposes also to define a group of small (*S*) correlation coefficients and a group of large (*L*) correlation coefficients, where η denotes the proportion of correlation coefficients in the *S* group. Once the sample of correlation coefficients has been split, we can proceed to test the null hypothesis of non correlation in both sub-samples.

Table 1 presents the results of calculating the *WCD*, WCD_{LM} and *svr* statistics for each panel data. The qualitative conclusion that we can drawn is that the *WCD* test clearly rejects the null hypothesis of no correlation, regardless of the deterministic function that is chosen – this conclusion is supported by the WCD_{LM} test statistic. The large values of these statistics can be taken as an indication that strong cross-section dependence is affecting the units of the panel data. This can be confirmed computing the degree of cross-section dependence in Bailey et al. (2012). As can be seen, the point estimate δ is close to one for the two variables for which it can be computed, although the 90% confidence interval defined by (δ_L, δ_U) define a wide range of values for this parameter.

The *svr* statistic in Ng (2006) shows that the null hypothesis of no correlation cannot be rejected at the 5% significance level for the small sub-sample of correlations for the debt, deficit and primary surplus – see the p-values associated to the *svr*(*S*) statistic – while it is clearly rejected when analyzing the sub-sample of large correlations – see the results for the *svr*(*L*) statistic. It should also be noted that the *L* group is largely more numerous than the *S* one, which indicates that, first, there is evidence of strong cross-section correlation and, second, that the correlation is pervasive – see Ng (2006). When we consider all correlations, the null hypothesis of cross-section independence is clearly rejected for all variables. It should be mentioned that the pervasiveness of the cross-section dependence suggests that panel data unit root and cointegration test statistics can capture the cross-section dependence by defining common factor models, as suggested by Bai and Ng (2004).

To sum up, this section has shown that the time series that define the panel data sets of our model are (pervasively) cross-section correlated. This conclusion is supported in an informal way when analyzing the pictures of the variables, where similarities in their evolution are easily observed. Either formally or informally, this feature indicates that the statistical inference that bases on the use of panel data analysis has to consider the presence of cross-section dependence.

6.2. Panel data order of integration analysis

The non-stationary panel data literature has proposed various ways to incorporate cross-section dependence when assessing the order of integration in panels of data. In the first stage, the dependence was intended to be captured removing the cross-section mean of the series, with the hope that the resulting variables were already independent – see Im, Pesaran and Shin (2003), among others. This strategy is equivalent to introduce panel data temporary effects and implies assuming the existence of a common I(0) stationary factor that has the same effect on all series. This solution, though computationally simple, implies assuming a situation that hardly occurs in practice, i.e., it is difficult to argue that the common factor will have the same effect on all units. A second approach is proposed by Maddala and Wu (1999), who suggest obtaining the empirical distribution of the test statistics using bootstrap techniques. The idea is to resample the whole cross-section dimension to preserve the cross-section dependence. Finally, a third approach models the cross-section dependence by specifying an approximate common factor model, which constitutes a simple device to capture the dependence structure. In this section and provided the conclusions obtained above, we opt for the third option, applying panel data unit root test statistics that incorporate unobservable common factors to capture the cross-section dependence.

Bai and Ng (2004), Moon and Perron (2004) and Pesaran (2007) are three of the proposals available in the literature that include the use of common factors when testing for the order of integration. Because Bai and Ng (2004) is more general than the other two mentioned approaches, we proceed to briefly describe it here. The framework of Bai and Ng (2004) assumes that the observable variable $y_{i,t}$ can be decomposed into a deterministic component $D_{i,t}$, a common component $\lambda_i' F_t$ and an idiosyncratic component $e_{i,t}$:

$$y_{i,t} = D_{i,t} + \lambda_i' F_t + e_{i,t} \tag{12}$$

$$(1-L) F_{j,t} = C_j(L) w_{j,t}; \quad j = 1, \ldots, r \tag{13}$$

$$(1-\rho_i L) e_{i,t} = H_i(L) \varepsilon_{i,t}, \tag{14}$$

where $D_{i,t}$ denotes the deterministic part of the model – either a constant or a linear time trend – F_t is a $(r \times 1)$-vector of unobservable common factors, and $e_{i,t}$ is the idiosyncratic disturbance term. The $(r \times 1)$-vector of loading parameters λ_i measures the effect that the common factors have on the *i*-th time series. The unobserved common factors and idiosyncratic disturbance terms are estimated using principal components on the first difference model – the estimation of the number of common factors is obtained using the panel Bayesian information criterion (BIC) in Bai and Ng (2002).

Once both the idiosyncratic and common components have been estimated, we can proceed to test their order of integration using unit root tests. On the one hand, it is possible to test whether there are I(0) stationary or I(1) non-stationary common factors (F_t) using the ADF (for the one common factor case, r = 1) or the MQ test statistics in Bai and Ng (2004) (for the general case where there are more than one common factor, r > 1) – either in its parametric ($MQ_f^j(m)$) and/or non-parametric ($MQ_c^j(m)$) version, where $j = c$ for the model that includes a constant, $j = \tau$ for the model that includes a linear time trend and m denotes the number of stochastic trends under the null hypothesis. The critical values for the MQ tests for up to six common factors can be found in Table 1 of Bai and Ng (2004), whereas the usual critical values of the Dickey-Fuller test can be used in the case of one common factor.

Using these statistics we will be able to conclude how many (if any) of the r common factors that have been estimated are I(0) stationary common factors (r_0) and how many are I(1) non-stationary common factors (r_1), so that $r = r_0 + r_1$. On the other hand, we can test the panel unit root hypothesis focusing on the idiosyncratic shocks ($e_{i,t}$). In this case, Bai and Ng (2004) propose to compute the usual ADF pseudo t-ratio statistic applied to the idiosyncratic component. If the model contains only an intercept, the pseudo t-ratio statistic is denoted as $ADF_{\hat{e}_i}^{c}$ and its asymptotic distribution coincides with the Dickey-Fuller distribution for the case of no constant for the deterministic specification. If the model has an intercept and a linear trend the statistic is denoted as $ADF_{\hat{e}_i}^{\tau}$, which asymptotic distribution is function of a Brownian bridge. Assuming that $e_{i,t}$ are cross-section independent, a pooled ADF test statistic can be defined to test the null hypothesis of panel unit root.

As can be seen, this technique can determine the source of the non-stationarity that is present on the observable variable. It is possible that the non-stationarity of the observed variables ($y_{i,t}$) is the result of the presence of I(1) common factors – or a combination of I(0) and I(1) common factors – which would imply that the panel data set is non-stationary and that the source of non-stationarity is a common cause for all the units in the panel. In this case, we should conclude that there are global permanent shocks affecting the whole panel. It could also be the possible that the source of non-stationarity of the panel is idiosyncratic – i.e., the idiosyncratic disturbance terms are I(1) non-stationary processes – a fact that implies that shocks that affect only each time series have a permanent character.

The approach of Bai and Ng (2004) nests the ones in Moon and Perron (2004) and Pesaran (2007). As noted by Bai and Ng (2009), the proposals in Moon and Perron (2004) and Pesaran (2007) control the presence of cross-section dependence allowing for common factors, although the common factors and idiosyncratic shocks are restricted to have the same order of integration. Therefore, it is not possible to cover situations in which one component (e.g., the common factors) is I(0) and the other component (for example, the idiosyncratic shocks) is I(1), and vice versa. In practical terms, the test statistics in Moon and Perron (2004) and Pesaran (2007) turn out to be statistical procedures to make inference only on the idiosyncratic shocks, where the dynamics of both the idiosyncratic and the common components are restricted to be the same.

Table 2 provides the results of the two test statistics – denoted as CIPS and CIPS* – proposed in Pesaran (2007) for different values of the order of the autoregressive correction (p) that is used when estimating the augmented Dickey-Fuller (ADF) auxiliary regression equations. As can be seen, the results vary depending on the value of p that we use. In general, we can see that, with marginal exceptions, the results that are obtained lead to the non-rejection of the null hypothesis of panel data unit root at the 5% significance level, regardless of the order of the autoregressive correction. Table tab-pdur also includes the results of the test statistics proposed by Moon and Perron (2004) – denoted by t_a and t_b. As we can see, the null hypothesis of panel data unit root cannot be rejected at a significance level of 5% for any of the variables and neither test statistic, regardless of the number of common factors (r) that is considered. Therefore, the widespread evidence obtained by applying these two testing approaches is that we can consider that all the variables in the study are I(1) non-stationary stochastic processes. However, the evidence obtained with Pesaran and Moon-Perron test statistics may be biased because of the assumption that the dynamics of the common factors is the same as the one driving the idiosyncratic disturbance term. This limitation is overcome by the proposal in Bai and Ng (2004), which analyses the order of integration of the common factors and the idiosyncratic disturbance terms in a separate way.

Table 2 also reports the test statistics in Bai and Ng (2004). The conclusion obtained from these statistics is that all variables present of symptoms of being I(1) non-stationary stochastic processes, as in all cases we detected the presence of I(1) non-stationary common factors. Therefore, regardless of the stochastic properties of the idiosyncratic disturbance terms, the panel data sets used in the study are I(1) non-stationary panel data sets.

In summary, the panel data unit root test statistics that have been applied in this section allow us to conclude that the fiscal panel data are I(1) non-stationary panels. This conclusion is robust to the presence of a strong cross-section dependence structure linking the units of the panels, a feature that, according to the test statistic in Ng (2006), is present in our data sets.

6.3. Panel data cointegration

The panel data unit root test statistics that have been applied in the previous section indicate that the variables involved in our model are I(1) non-stationary variables. The use of these variables in levels may lead to obtain wrong conclusions as we might be facing a spurious relationship. In this regard, it is necessary to test whether the relationship posed by the model that analyzes the fiscal deficit sustainability is a long-run relationship (an equilibrium relationship with economic meaning) or not (a spurious relationship). In order to decide which is the actual situation, we proceed to apply panel data cointegration test statistics taking into account the presence of cross-section dependence. The econometric literature in this area is limited and quite recent, although it is possible to find some proposals that fit our requirements. In this regard, we propose to apply the panel cointegration test statistics in Banerjee and Carrion-i-Silvestre (2011, 2014), Westerlund (2008) and Bai and Carrion-i-Silvestre (2013). These test statistics meet the needs that our analysis requires, since they account for the presence of cross-section dependence among the units of the panel through the specification of an approximate common factor model.

It is worth mentioning that there are some important features that share and distinguish these proposals. First, they assume that the disturbance term of the model – for instance, let us focus on the model given in (6) – can be decomposed as the sum of a common and an idiosyncratic components:

$$u_{i,t} = \lambda_i' F_t + e_{i,t},$$

where F_t is a $(r \times 1)$-vector of unobservable common factors, λ_i is a $(r \times 1)$-vector of loading parameters and $e_{i,t}$ is the idiosyncratic disturbance term. One important difference concerns to the order of integration of F_t, since Westerlund (2008) considers that all common factors are I(0) stationary common factors, whereas the other approaches assume that there might be a combination of I(0) and I(1) common factors as in Bai and Ng (2004). Second, Bai and Carrion-i-Silvestre (2013) considers the most general case where the common factors might both affect the dependent variable and the stochastic regressors, whereas the other proposals assume that the common factors and the stochastic regressors are orthogonal. Third, in Banerjee and Carrion-i-Silvestre (2011) the effect of the unobserved common factors is taken into account as in Pesaran (2006), who uses cross-section averages to proxy the common factors. The other proposals estimate the common factors using principal components as in Bai and Ng (2004). Finally, the deterministic specification in Westerlund (2008) is given by a constant term, so that his test statistics cannot be computed if a linear time trend is specified – this does not apply for the other proposals.

We have discussed above two alternative approaches to analyze the sustainability of fiscal policy. First, we estimate the model given by equation (6) relating the revenues and expenditures:

$$R_t = \mu + \beta G_t + u_t,$$

and test the presence of cointegration in panel data. The second alternative is postulated in Bohn (1998) and is based on the estimation of the fiscal rule given by equation (8):

$$S_t = \mu + dt + \rho B_t^* + v_t, \tag{15}$$

with the inclusion of a linear trend as a deterministic component to capture the growth of debt. In order to estimate (15) Bohn (1998) considers that the level of debt at the beginning of period *t* is proxied by the level of debt in the period *t*-1. We address each of these two approaches in a separate way.

6.3.1. Relationship between revenues and expenditures

Table 3 shows that the procedure of Banerjee and Carrion-i-Silvestre (2006) detects the presence of a I(1) non-stationary common factor – the MQ test statistics in Bai and Ng (2004) indicate that there is an I(1) non-stationary common stochastic trend – that drives the cross-section dependence of the panel data model. As for the panel cointegration, we see that the ADF test applied to the idiosyncratic disturbance terms lead to reject the null hypothesis of no panel cointegration at the 5% significance level. Consequently, there is evidence for a long-term relationship (cointegration) between the revenues and expenditures once the cross-section dependence has been taken into account.

This conclusion is also achieved with the application of the $CADF_p$ test statistic regardless of the order of the autoregressive correction (*p*) that is applied. With the application of this procedure we obtain a parameter estimate of $\hat{\beta} = 0.948$ in (6), estimation that, regardless of the presence of panel cointegration, is a consistent estimate of the relationship between the revenues and expenditures.

As for the DH test statistics of Westerlund (2008), the conclusions depends on the degree of homogeneity that is assumed. While the test statistic that allows for heterogeneity in the autoregressive coefficient (DH_g) leads to reject the null hypothesis of no panel cointegration, the one that imposes homogeneity (DH_p) does not – the exception is found when three common factors are specified. This contradiction between these two test statistics and the results obtained with the other two procedures may be indicating that the assumption of homogeneity when testing for the presence of cointegration does not seem advisable.

The application of the test statistic in Bai and Carrion-i-Silvestre (2013) provides more evidence for the presence of a cointegration relationship between revenues and expenditures, regardless of the number of common factors that are considered in the analysis – see Table 4. As can be seen, the estimated β parameter in (6) is around one, with values ranging from 0.942 to 1.01, depending on the number of factors that is included.

6.3.2. Relationship between primary surplus and debt

In this case, the panel data cointegration test statistics in Banerjee and Carrion-i-Silvestre (2011, 2014) indicate that there is no cointegration between the primary surplus and the debt. This conclusion is reversed when computing the test statistics in Bai and Carrion-i-Silvestre (2013) as the null hypothesis of no panel data cointegration is clearly rejected by the P_m and P statistics regardless of the number of common factors that is specified, and for the MSB_{ξ} statistic if four common factors are considered. The fact that the latter proposal is more general than the previous ones – as mentioned above, it considers that the common factors that affect the dependent variable may also be affecting the explanatory variables – leads us to attribute more weight to the findings that indicate that there is a long-run relationship between the primary surplus and the debt.

It should be noted that regardless of whether there is a cointegration relationship, using the common correlated effects (CCE) estimator in Banerjee and Carrion-i-Silvestre (2011) allows us to obtain a consistent estimate of the ρ parameter that measures the relationship between the primary surplus and the debt. Table 3 indicates that the value of this parameter is 0.064, whereas the estimates in Table 4 ranges between 0.047 and 0.105, values that, in principle, would satisfy the sustainability condition of Bohn (1998) – at this stage it is worth noticing that we still need to assess the statistical significance of such parameter estimates.

6.4. Estimation of the cointegration relationships: Fiscal deficit sustainability analysis

In this section we present the estimated cointegrating relationships of the two approaches that are used to determine the sustainability of the fiscal policy of the Spanish ACs. Due to the presence of cross-section dependence, the procedures for estimating the cointegrating relationships that are used are the ones proposed in Bai, Kao and Ng (2009) and Kapetanios, Pesaran and Yamagata (2011).

The approach of Bai, Kao and Ng (2009) estimates the cointegration vector using procedures that render consistent and efficient estimates of the parameters – Continuous Updated Fully-Modified (CUP-FM) and Continuous Updated Bias Corrected (CUP-BC) estimators – considering the presence of I(0) and/or I(1) common factors. Given the efficiency property of these estimators, we can perform inference on the estimated parameters – in the limit the estimated parameters are distributed according to a normal distribution. The strategy in Kapetanios, Pesaran and Yamagata (2011) bases on the CCE approach in Pesaran (2006). These authors show that under panel cointegration, the pooled CCE estimator is a consistent estimator of the cointegration vector, which is asymptotically distributed as a normal distribution. There is an important feature that distinguish both proposals. Thus, whereas Kapetanios, Pesaran and Yamagata (2011) assume that the stochastic regressors are weakly exogenous, Bai, Kao and Ng (2009) specify a more general framework where the stochastic regressors might be endogenous.

Table 5 shows the results of the estimates for the two models and three estimators. Regarding the cointegration relationship between the revenues and expenditures, we see that the parameter estimates are clearly statistical significant with values that are placed around one. In this regard, the null hypothesis that the parameter is equal to one is not rejected at the 5% significance level values for the CCE and CUP-FM estimates, where it is rejected when using the CUP-BC estimate – note that the parameter in this case is larger than one, which leads to fiscal surpluses. Consequently, these results indicate, in the first place, that the fiscal deficit of the Spanish ACs is clearly sustainable and, secondly, the cointegration vector is (1, -1). As argued in Quintos (1995), the combination of these two conditions defines a necessary and sufficient condition for the sustainability of public deficit in the sense that the fiscal policy of the ACs is consistent with the intertemporal budget constraint. Therefore, the fiscal deficit is sustainable in the strict sense (strong sustainability).

The second part of Table 5 shows the results of estimating Bohn’s (1998) model. The estimated parameters are greater than zero, although the estimates that are statistical significant are those that rely on the use of the efficient procedure in Bai, Kao and Ng (2009). Thus, from the three estimates that have been obtained, two of them indicate that the sufficient condition of sustainability advocated by Bohn (1998) is met. The estimates indicate that for every 100 euros that increases the debt of the Spanish ACs, the primary surplus increases by around 8 euros (7.5 according to the CUP-FM and 8.2 according to the CUP-BC estimates), which is consistent with a reaction that satisfies the intertemporal budget constraint.

It is interesting to note that the average interest rate of the debt for the analyzed period and for all the Spanish ACs is $r = 0.075$, a value that is in accordance with the values estimated for the ρ parameter (in fact, equals the estimated CUP-FM parameter). Therefore, if one defines the hypothesis that $\rho = r = 0.075$, one can see that this null hypothesis cannot be rejected when using either the CUP-FM estimates – being the value of the estimated parameter equal to the value under the null hypothesis – of the CUP-BC estimates – in this case, the value of the test statistic is (0082-0075) / 0006 = 1.17, a value that is smaller than the critical value at the 5% level of significance of a standard normal distribution. Following Bohn (2007), these results would seem to indicate that the debt is an I(1) non-stationary variable – which is the result that has been obtained when we have conducted the order of integration of the variable above – and that the fiscal deficit is an I(0) stationary process – a result that has also been obtained previously when it was concluded that the revenues and expenditures generate a cointegration relationship with a cointegration vector (1, -1).

In summary, the different approaches that have been followed in this paper have led to conclude that the fiscal policy of the Spanish ACs is sustainable in the sense that in the long-run the intertemporal budget constraint is met and, in addition, that the primary surplus reacts positively and significantly to changes in the level of debt of the ACs.

7. Conclusions

The paper analyzes the sustainability of the deficit of the Spanish ACs in the period 1984-2009 using consolidated data on revenues, expenditures and debt, all expressed in real terms. The literature on the sustainability of public deficit is divided into two major approaches. First, we have the analyses that are based on the cointegration analysis of the fiscal variables. Second, we have the studies based on fiscal rules that relate the primary surplus to the level of debt of the economies. Both approaches have been presented in the paper, discussing the similarities and differences that characterize them in order to draw a conclusion as robust as possible on the question of the sustainability of fiscal policy.

A first set of results in the paper clearly shows that the variables involved in both approaches share the characteristics of being I(1) non-stationary variables and being affected by the presence strong cross-section dependence. The cross-section dependence has been captured through the use of parsimonious models by specifying common factor models, which are able to account for global stochastic trends. This just makes it clear that the system of financing of the ACs and the competences that they have undertaken are driven by the same legal framework that brings up this strong (penetrating) cross-section dependence.

A second important result has been to evidence how the deficit of the Spanish ACs is sustainable in the long-run, regardless of the analytical approach that is chosen. It has been shown that the revenues and expenditures of the ACs generate a long-run equilibrium relationship, with a (1, -1) cointegrating vector. This result defines a sufficient condition for the sustainability of the public deficit.

The paper has also shown that the fiscal rule that relates the primary surplus and debt levels at the beginning of the period has also proved to define a cointegration relationship, with a parameter of interest that is similar to the average interest rate of the debt of the analyzed period. This implies that both approaches have come to the same conclusion, namely, the fiscal deficit of the ACs in Spain is sustainable in the strong sense for the analyzed period.

References

Afonso, A., 2005. Fiscal sustainability: the unpleasant European case. FinanzArchiv, 61, 19-44.

Bai, J., Carrion-i-Silvestre, J.L., 2013. Testing panel cointegration with unobservable dynamic common factors that are correlated with the regressors. Econometrics Journal, forthcoming.

Bai, J., Kao, C., Ng, S., 2009. Panel cointegration with global stochastic trends. Journal of Econometrics 149, 1, 82-99.

Bai, J., Ng, S., 2004. A PANIC attack on unit roots and cointegration. Econometrica 72, 4, 1127-1177.

Banerjee, A., Carrion-i-Silvestre, J. L., 2011. Testing for panel cointegration using common correlated effects estimators. Document de treball. Universitat de Barcelona.

Banerjee, A., Carrion-i-Silvestre, J. L., 2014. Cointegration in panel data with breaks and cross-Section dependence. Journal of Applied Econometrics, forthcoming.

Banerjee, A., Marcellino, M., Osbat, C., 2004. Some cautions on the use of panel methods for integrated series of macro-economic data. Econometrics Journal, 7, 322-340.

Banerjee, A., Marcellino, M., Osbat, C., 2005. Testing for PPP: should we use panel methods?. Empirical Economics, 30, 77-91.

Barro, R., 1986. U.S. deficits since World War I. Scandinavian Journal of Economics, 88, 195-222.

Berenguer-Rico, V., Carrion-i-Silvestre, J. L., 2011. Regime shifts in stock-flow I(2)-I(1) systems: the case of US fiscal sustainability. Journal of Applied Econometrics, 26, 298-321.

Bohn, H., 1998. The behavior of U.S. public debt and déficits. The Quarterly Journal of Economics, 13, 949-963.

Bohn, H., 2005. The sustainability of fiscal policy in the United States. Tech. Rep. 1446, CESifo, Munchen.

Bohn, H., 2007. Are stationarity and cointegration restrictions really necessary for the intertemporal budget constraint?. Journal of Monetary Economics, 54, 1837-1847.

Camarero, M., Carrion-i-Silvestre, J. L., Tamarit, C., 2012. The relationship between debt level and fiscal sustainability in OECD countries: a multicointegration approach. Working paper, Universitat de Barcelona.

Claeys, P., Ramos, R., Suriñach, J., 2010. Fiscal sustainability across government tiers. International Economics and Economic Policy, 5, 139-163.

Escario, R., Gadea, M., Sabaté, M., 2012. Multicointegration, seignorage and fiscal sustainability. Spain 1857-2000. Journal of Policy Modeling, 34, 270-283.

Esteller, A., Solé, A., 2004. Estabilidad presupuestaria y financiación autonómica. Hacienda Pública Española, monografia 2004, 173-201.

Granger, C. W. J., Lee, T., 1989. Multicointegration. Advances in Econometrics, 8, 71-84.

Hakkio, C., Rush, M., 1991. Is the budget deficit too large? Economic Inquiry, 24, 429-445.

Hamilton, J., Flavin, M., 1986. On the limitations of government borrowing: A framework for empirical testing. American Economic Review, 76, 808-816.

Haug, A., 1991. Cointegration and government borrowing constraints: Evidence for the U.S. Journal of Business & Economic Statistics 9, 97-101.

Holly, S., Pesaran, M. H., Yamagata, T., 2010. A spatio-temporal model of house prices in the US. Journal of Econometrics 158, 160-173.

Im, K., Pesaran, M. H., Shin, Y., 2003. Testing for unit roots in heterogeneous panels. Journal of Econometrics, 115, 53-74.

Kapetanios, G., Pesaran, M. H., Yamagata, T., 2011. Panels with nonstationary multifactor error structures. Journal of Econometrics 160, 326-348.

Maddala, G. S., Wu, S., 1999. A comparative study of unit root tests with panel data and a new simple test. Oxford Bulletin of Economics and Statistics Special Issue, 61, 631-652.

Martin, G., 2000. U.S. deficit sustainability: A new approach based on multiple endogeneous breaks. Journal of Applied Econometrics, 15, 83-105.

Moon, H. R., Perron, B., 2004. Testing for unit root in panels with dynamic factors. Journal of Econometrics, 122, 81-126.

Ng, S., 2006. Testing cross-section correlation in panel data using spacings. Journal of Business and Economic Statistics, 24, 12-23.

Pesaran, M. H., 2004. General diagnostic tests for cross section dependence in panels. CESifo Working Papers. No. 1233.

Pesaran, M. H., 2006. Estimation and inference in large heterogeneous panels with a multifactor error structure. Econometrica 74, 967-1012.

Pesaran, M. H., 2007. A simple panel unit root test in the presence of cross section dependence. Journal of Applied Econometrics, 22, 265-312.

Pesaran M. H., 2013. Testing weak cross-sectional dependence in large panels. Econometric Reviews, forthcoming.

Phillips, P. C. B., Hansen, B. E., 1990. Statistical inference in instrumental variables

regression with I(1) processes. Review of Economic Studies, 57, 99-125.
Phillips, P. C. B., Moon, H. R., 1999. Linear regression limit theory for nonstationary panel data. Econometrica, 67, 1057-1111.

Quintos, C. E., 1995. Sustainability of the deficit process with structural shifts. Journal of Business & Economic Statistics, 13, 409-418.

Saikonnen, P., 1991. Asymptotically efficient estimation of cointegrating regressions. Econometric Theory, 7, 1-21.

Stock, J. H., Watson, M. W., 1993. A simple estimator of cointegrating vectors in higher order integrated systems. Econometrica, 61, 783-820.

Trehan, B., Walsh, C., 1988. Common trends, the government budget constraint, and revenue smoothing. Journal of Economic Dynamics and Control, 12, 425-444.

Trehan, B., Walsh, C., 1991. Testing intertemporal budget constraints: Theory and applications to U.S. federal budget and current account deficits. Journal of Money, Credit and Banking, 23, 206-223.

Westerlund, J., 2008. Panel cointegration tests of the Fisher effect. Journal of Applied Econometrics, 23, 193-233.

Westerlund, J., Mahdavi, S., Firoozi, F., 2011. The tax-spending nexus: evidence from a panel of US State-local governments. Economic Modelling, 28, 885-890.

Wilcox, D., 1989. The sustainability of government deficits: Implications of the present value borrowing constraint. Journal of Money, Credit and Banking, 54, 1837-1847.

Table 1. Cross-section dependence

	Pesaran (2004, 2013)		Ng (2006)							Bailey et al. (2012)		
			Whole sample SVR		Small sample SVR(S)		Large sample SVR(L)					
	WCD	WCD_{lm}	Test	p-valor	Test	p-val	Test	p-val	η	δ	δ_L	δ_U
Revenues	28.65	15.80	3.74	0.00	2.65	0.00	3.17	0.00	0.22	-		
Expenditures	23.48	12.78	3.19	0.00	2.14	0.02	4.39	0.00	0.29	-		
Debt	27.80	19.95	6.64	0.00	-0.88	0.81	4.96	0.00	0.10	-		
Deficit	33.15	23.17	4.13	0.00	-0.50	0.69	3.80	0.00	0.10	0.9297	-24042	24044
Primary surplus	37.43	23.81	5.07	0.00	1.73	0.04	4.99	0.00	0.10	0.9273	-25739	25741

Table 2. Panel data unit root tests

	Revenues		Expenditures		Debt		Deficit		Primary surplus	
Pesaran (2007)										
p	CIPS	CIPS*	CIPS	CIPS*	CIPS	CIPS*	CIPS	CIPS*	CIPS	CIPS*
0	-2.28	-2.28	-1.65	-1.65	-2.18	-2.18	-2.85**	-2.80**	-2.67**	-2.56**
1	-2.47	-2.47	-1.79	-1.79	-2.32	-2.32	-2.02	-2.02	-1.88	-1.88
2	-2.48	-2.48	-1.65	-1.65	-2.37	-2.37	-1.86	-1.86	-1.79	-1.79
3	-2.75**	-2.75**	-1.45	-1.47	-2.53	-2.47	-1.31	-1.32	-1.33	-1.33
4	-1.97	-1.97	-1.05	-1.06	-1.97	-2.14	-1.24	-1.33	-1.26	-1.26
5	-2.25	-2.15	-1.48	-1.45	-1.63	-1.67	-0.97	-0.97	-0.87	-0.87
Moon and Perron (2004)										
	Revenues		Expenditures		Debt		Deficit		Primary surplus	
r	t_a	t_b	t_a	t_b	t_a	t_b	t_a	t_b	t_a	t_b
1	-1.05	-0.98	-1.38	-1.51	0.32	0.30	-7.90	-3.70	-5.39	-2.60
2	-0.24	-0.23	-1.18	-1.19	0.67	0.70	-9.22	-4.63	-7.62	-3.63
3	-1.74	-1.82	-0.84	-0.87	-1.25	-1.24	-8.92	-4.74	-10.88	-5.19
4	-1.23	-1.08	-1.66	-1.73	-3.08	-3.03	-9.09	-4.55	-12.43	-5.60
5	-0.81	-0.72	-1.50	-1.83	-1.48	-1.25	-11.67	-5.38	-12.24	-5.63
6	-0.24	-0.20	-1.34	-1.32	-1.77	-1.53	-12.44	-5.55	-14.07	-5.78
Bai and Ng (2004)										
	Revenues		Expenditures		Debt		Deficit		Primary surplus	
	Test	p-val	Test	p- val	Test	p- val	Test	p- val	Test	p- val
ADF idiosyncratic	-1.75	0.04	-0.54	0.30	-1.39	0.08	-3.24	0.00	-2.61	0.01
	Revenues		Expenditures		Debt		Deficit		Primary surplus	
	Test	$\hat{r}$	Test	$\hat{r}$	Test	$\hat{r}$	Test	$\hat{r}$	Test	$\hat{r}$
MQ (no par.)	-24.49	6	-16.77	6	-23.53	6	-18.94	6	-22.89	6
MQ (par.)	-24.25	6	-23.67	6	-21.72	6	-22.23	6	-21.08	6

Note: ** indicates rejection of the null hypothesis of unit root at the 5% level of significance. Moon and Perron (2004) statitistics distribute as a standard normal distribution under the null hypothesis of unit root

Table 3. Banerjee i Carrion-i-Silvestre (2011, 2014) and Westerlund (2008) panel cointegration test statistics

Banerjee and Carrion-i-Silvestre (2014)

	Revenues and expenditures			Primary surplus and debt		
	Test	p-val		Test	p-val	
ADF-idio	-1.71	0.044		-0.02	0.492	
	Test	$\hat{r}_1$	$\hat{r}$	Test	$\hat{r}_1$	$\hat{r}$
MQ (no par.)	-3.20	1	1	-7.36	1	1
MQ (par.)	-2.65	1	1	-2.73	1	1

Banerjee and Carrion-i-Silvestre (2011)

	Revenues and expenditures		Primary surplus and debt	
p	$CADF_p$	$\hat{\rho}$	$CADF_p$	$\hat{\rho}$
0	-3.32**	0.948	-3.89**	0.064
1	-2.68**		-2.56	
2	-2.64**		-2.20	
3	-2.33**		-1.84	
4	-2.50**		-2.04	

Westerlund (2008)

	Revenues and expenditures			
r	DH_g	p-val	DH_p	p-val
1	26419.74	0.000	0.47	0.318
2	14390.17	0.000	-0.23	0.590
3	14.18	0.000	1.69	0.046
4	4.83	0.000	0.33	0.371
5	6.87	0.000	-0.72	0.763
6	5.84	0.000	-0.84	0.800

Table 4. Bai and Carrion-i-Silvestre (2013) panel cointegration test statistics

Revenues and expenditures

r	$MSB_{\tilde{\xi}}$	p-val	P_m	p-val	P	p-val	$\hat{\beta}$
1	-2.67	0.00	7.17	0.00	93.11	0.00	0.977
2	-2.45	0.01	4.08	0.00	67.63	0.00	1.010
3	-2.71	0.00	5.44	0.00	78.83	0.00	0.952
4	-2.64	0.00	5.25	0.00	77.27	0.00	0.974
5	-2.66	0.00	4.60	0.00	71.94	0.00	0.963
6	-2.59	0.01	3.50	0.00	62.90	0.00	0.942

Primary surplus and debt

r	$MSB_{\tilde{\xi}}$	p-val	P_m	p-val	P	p-va	ρ
1	-0.86	0.20	3.46	0.00	62.54	0.00	0.080
2	-0.13	0.45	2.20	0.01	52.16	0.02	0.105
3	-0.78	0.22	1.38	0.08	45.39	0.09	0.082
4	-1.10	0.14	2.34	0.01	53.27	0.02	0.075
5	-1.40	0.08	1.98	0.02	50.30	0.04	0.085
6	-1.97	0.03	2.26	0.01	52.65	0.02	0.047

Table 5. Panel data cointegrating vector estimates)

Revenues and expenditures

	CCE	CUP-FM	CUP-BC
β	0.948	0.998	1.024
s.e.	0.048	0.005	0.004
t-ratio $\beta = 0$	19.577	219.612	228.390
t-ratio $\beta = 1$	-1.079	-0.440	5.353

Primary surplus and debt

	CCE	CUP-FM	CUP-BC
ρ	0.064	0.075	0.082
s.e.	0.072	0.006	0.006
t-ratio $\rho = 0$	0.883	12.423	13.466