Macroeconomic stability and fiscal decentralization in the Spanish autonomous communities

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Abstract

The paper analyzes the relationship among fiscal decentralization and the inflation using regional data for the Spanish economy. The use of non-stationary panel data techniques indicates that there is a long-run relationship among these variables, although the effect of fiscal decentralization on the inflation rate depends on the degree of competences that each autonomous community has assumed.

Keywords: Fiscal decentralization, inflation, cointegration

JEL codes: E62, H77

1 Introduction

The economic literature has analyzed the stability of the economies by focusing on key economic variables such as inflation and unemployment. The evolution of these variables is clearly influenced by the fundamentals of the economy that eventually determines the level of production of the economies. These principles encompass both elements involved directly in the production phases – the level of physical capital, human capital of workers and economies – and aspects such as the institutional and political framework that guarantees a stable relationship among the agents involved in the market. Economic agents sometimes demanded the participation of government institutions in order to undertake legislative and regulatory initiatives that promote the improvement of the relational framework of the economy so that they can attract new investment and increase productivity. In this regard, government institutions must perform actions for ensuring that the conditions for carrying out economic activities are the more appropriate ones.

These actions give rise to tax policies that, by the raising and spending capacities, make governments to act as an agent in the economy. However, fiscal policy actions are not without critics, depending on the prevailing economic paradigm, which can range from Keynesian visions to more liberals positions. On one hand, a too interventionist government may introduce

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distortions in the markets that make them not to act efficiently. On the other hand, the performance of the government can correct market imbalances leading to more equitable situations from a social point of view.

Developed economies have experienced a gradual process of fiscal decentralization that has pursued closer administrative practice in the territories where eventually the governments have an impact. However, this decentralization process might cause possible dangers in macroeconomic terms, because as the level of the final decision comes to citizens, there is a tendency to increase public spending. In this sense, some authors advocate a more centralized government activity in one agent, while others advocate decentralization – with the objective of providing more properly what is needed in each region.

This paper focuses on the area of fiscal decentralization and its relationship with macroeconomic stability, analyzing whether the process of fiscal decentralization, which is based in part on an institutional decentralization process, has affected the evolution of variables such as the inflation rate of the economies. Our study increases the empirical evidence available in the literature analyzing the case of the Spanish economy, working at a level of territorial breakdown of the seventeen autonomous communities (ACs) in the period 1979-2000. The topic of the paper is interesting because the Spanish case is an economy that has experienced a fiscal decentralization process since the beginning of the democracy era that, surprisingly, has not been addressed so far. The panel data analysis that has been conducted in this paper leads us to conclude that there is a long-run relationship between inflation and fiscal decentralization, where the effects that (regional and local) fiscal decentralization have on the inflation depend on the degree of competences that the ACs have assumed.

The paper is structured as follows. Section 2 describes briefly what was the process of fiscal decentralization that has been followed in Spain since the beginning of the constitutional period to the present. Section 3 provides an overview of the main theories and empirical studies that examined the relationship between inflation and fiscal decentralization. Section 4 presents the empirical model and the database that is used. Section 5 is devoted to the econometric methodology that is applied. Section 6 discusses the results of the estimation, relating them to the theories presented above. Finally, Section 7 concludes.

2 Fiscal decentralization in Spain

The Spanish constitution in 1978 created the division of Spain into seventeen autonomous communities (regions), fifty provinces and, approximately, eight thousand municipalities. Associated with this territorial division there are three levels of government: central government, regional governments and local governments.

The distribution of competences established by the Spanish constitution for each of these levels of government is ruled by the Spanish constitution, by the regional own laws (regional constitutions) and the Act of Autonomy of Local Governments. The central government reserve competences affecting the entire territory, in matters such as defense, foreign affairs, economic stability and the part of the social security system that has to do with pensions and unemployment benefits. In addition, with the exception of Catalonia and the Basque Country, it also have competences in public safety.

The Spanish constitution set up two types of ACs in accordance with the competences that have been granted. The ACs regulated by Article 143 of the Spanish constitution are known as the ACs of common regime – these ACs have the competences on education and health temporarily excluded. The ACs regulated by Article 151 of the constitution were assigned common competences, education and health since 1978. From this point of view, we conclude that the ACs governed by Article 143 of the constitution have a low level of responsibility when compared with regions of Article 151 of the constitution. It is worth noticing, however, that since the beginning of democracy period there has been a process of transferring competences from the central government to the Article 143 ACs in matters of education and health. Competences in education have been shifting gradually between 1995 and 1999 in the ACs ruled by Article 143, while competences in health were transferred to all ACs in 2002.

The regional constitutions of each AC have deployed the competences assigned in a different way so that, even though in theory all regions have attributed common competences, in practice not all regions have developed them. Therefore, not all regions have de facto the same responsibilities when providing public goods and services. At this point we want to make evident the existence of a clear division among the Spanish ACs according on the assumed and deployed competences.

It is also possible to establish another division of the ACs depending on their financing system. The financing of the ACs is set by the Finance Act of the Autonomous Communities (LOFCA), which establishes two funding regimes: (i) the foral regime, which is only applicable for the Basque Country and Navarre, and (ii) the common regime, which is applicable to the remaining ACs. The foral regime establishes an economic agreement where the Basque Country and Navarre manage all taxes and make a transfer of revenue to the central government for the common services provided in their territory. In the case of the ACs under the common regime the relationship is the inverse, that is, the central government collects the taxes and makes transfers to the governments of the ACs depending on the competences that they have attributed – see Carrion-i-Silvestre, Espasa and Mora (2008) for details on the resources allocated to the different regions.

The different competences that have assumed the ACs and the different autonomous financing systems that are applied establish a clear division among the Spanish ACs. On one side, the ACs ruled by Article 143 of the Spanish constitution, the ones governed by Article 151 of the Spanish constitution and, finally, the foral ones (Basque Country and Navarre). This division involves different levels of fiscal decentralization, which can cause different impacts on the macroeconomic stability.

The lower level of government is defined by the local governments, which are regulated by the Act of Local Governments, which sets the minimum level of services required to be provided to the population. However, it should be noticed that many of the responsibilities that go beyond the municipal level are shared by the central and regional governments – for example, roads and local transport, railways, housing, social services and development policies.

Figures 1 and 2 show the evolution of public expenditures and revenues, respectively, in the period 1970-2001 for the whole of Spain. As can be seen, the process of fiscal decentralization has been gradual over the period, being more intense in the area of public expenditures than in revenues. The change in the degree of decentralization in the period 1980-2001 shows a pattern that reflects the growing process of institutional change designed in the Spanish constitution. Consequently, the case of Spain is interesting to analyze from the perspective of fiscal decentralization and its relationship with macroeconomic stability.

3 Macroeconomic stability and fiscal decentralization: an overview

3.1 Theories that explain the relationship between macroeconomic stability and fiscal decentralization

The economic literature has not established a clear relationship between fiscal decentralization and macroeconomic stability. If we focus on inflation as a proxy variable of macroeconomic stability, we can see that this situation of uncertainty about whether exists a relationship between fiscal decentralization and inflation depends on, first, the action of the central government and sub-central (regional and local) governments and, second, of their political and institutional relationship.

Treisman (2000) distinguishes three types of theoretical arguments that, depending on the role of the different levels of government, may give rise to a potential relationship between inflation and fiscal decentralization: (i) a negative relationship (commitment hypothesis), (ii) a positive relationship (collective action problem), or (iii) absence of relationship (continuity hypothesis).

The commitment hypothesis argues that the dominant strategy of the central government is to broke the promise of monetary stability since the non-anticipated inflation has no real effects on the economy – for example, the central governments are more likely to increase public spending in periods prior to the electoral process. In this regard, central government fiscal policy actions can be controlled or limited by sub-central governments if fiscal policy is decentralized. Consequently, fiscal decentralization is expected to help meeting the objectives of monetary stability – keeping inflation controlled – which implies a negative relationship between inflation and the level of fiscal decentralization – higher decentralization degrees involve lower inflation rates.

Inflation can also be explained from the point of view of collective action, where as more agents are decision makers on monetary and fiscal policy, the greater the incentive to violate the purpose of monetary stability by sub-central governments. Thus, sub-central governments would have incentives to moderate their spending because, on the one hand, the effects (benefits) occur in their region and, on the other hand, the inflation that is generated (costs) due to the increase of the money supply is diluted in the overall state economy. Therefore, the central government is the only one interested in price stability, while the sub-central governments have less incentive to restrictive fiscal policies. If this is the dominant strategy in a given economy, we should expect that higher levels of fiscal decentralization would lead to increases in inflation.

Finally, it is possible to consider the relationship between inflation and fiscal decentralization from the perspective of the continuity hypothesis. This hypothesis argues that the level of decentralization does not directly affect inflation, but that would shield inflation levels (either high or low) of the economy. According to Tsebelia (1995), as the number of agents with veto power increases, more anchored would be the economic policy to the current situation, regardless of whether it is either an inflationary or austerity policy. The fact that the agents with veto have to reach an agreement to change economic policy, makes that the persistence or inertia of economic policies is high on over time. For example, if the level of inflation in the economy is high as a result of an expansionary fiscal policy of the sub-central government, they have no incentive to change to a more restrictive one, since the current situation benefits their territory, even at the expense of maintaining high levels of inflation. If, however, the level of inflation in the economy is low, the central government will have more incentive to keep prices stable, trying to prevent public spending increases. Consequently, initial situations concerning the inflation rate (either low or high) would tend to perpetuate themselves, so that the level of decentralization would not have any effect on the inflation.

So far, we have discussed the different approaches that can help to explain the possible relationship between fiscal decentralization and inflation. However, another relevant agent that adds to the economy ability to control inflation is the central bank. We often find that in developed economies, the central bank is in charge of controlling inflation, regardless of the fiscal policy that may undertake the different levels of government. Thus, governments are no longer able to easily monetize their debt, having to resort to domestic and international financial markets to raise funds. In this case, the central government is more interested in maintaining price stability as the risk premium that would pay to put their debt in the market will be lower. Therefore, any analysis that aims to study the possible relationship between inflation and fiscal decentralization should take into account the degree of independence that the central bank may have in terms of monetary policy.

3.2 Empirical evidence

The theoretical contributions that have been summarized in the previous section offer different cases that can be found in practice when focusing on a particular economy. This explains the diversity of results that are found when reviewing the available empirical evidence.

First, we can highlight a group of papers that find a negative relationship between fiscal decentralization and inflation. Gramlich (1995), Shah (1999) and Rodden and Wibbel (2002) recognize that giving some decision-making power to sub-central governments can promote macroeconomic stability. King and Ma (2001) and Neyapti (2004) analyzed two samples of OECD countries and found a significant negative relationship between fiscal decentralization and inflation. Martínez-Vázquez and McNab (2006) also concluded on the same line when analyzing a sample that consider a group of developed, developing and transition economies, but also in the sub-sample of developed countries.

Among the group of papers that suggest a positive relationship between fiscal decentralization and inflation we have those of Rodden (2002) and Rodden, Eskeland and Litvak (2003), which indicate that fiscal decentralization per se exacerbates macroeconomic instability due to the lack of sub-central governments budget control. Similarly, Martínez-Vázquez and McNab (2006) when analyzing a sub-sample of developing and transition countries found a statistically significant relationship between fiscal decentralization and inflation, which suggests a reduction in the level of fiscal decentralization if price stability is to be maintained.

Thornton (2007) criticizes the approach of King and Ma (2001) and Neyapti (2004) for having used measures of fiscal decentralization that do not take into account the autonomy level of government to take decisions on the spending or revenues.[1] In an attempt to take this into account, Thornton (2007) uses information published by the OECD on the real autonomy of each level of government to collect taxes, but only for 1995. He concludes that the level of fiscal decentralization is not a relevant variable in determining inflation, an opposite conclusion to the one reached in King and Ma (2001) and Neyapti (2004).[2]

Similar to Thornton (2007), Triesman (2000) and Rodden and Wibbels (2002) argue that there is clear evidence of the existence of a relationship between inflation and fiscal decentralization. Triesman (2000) finds that fiscal decentralization tends to shield the levels of inflation regardless of whether they are low or high. Triesman (2000) mentions that in some developed countries the federal structure has contributed to ensure the central bank independence, helping the inflation levels to maintain at low levels. However, in some developing countries, the federal structure has not helped to ensure either the central bank independence or the reduction of central government debt.

The papers mentioned above analyze the possible existence of a relationship between inflation and the level of fiscal decentralization by estimating regression models that incorporate other relevant explanatory variables. Among others, the models tend to include economic variables – such as the level of income per capita, the growth rate or the degree of openness of the economy – social, institutional and political variables – indicators of armed conflict, defense spending over income, political instability or independence of the central bank. Although these variables are relevant from a theoretical point of view, they might not be available when the analysis focus on regional data.

Interestingly, a variable that has not been considered as an explanatory variable in the papers mentioned above is the unemployment rate, although there is an extensive literature that relates inflation and unemployment rates – see, for instance, the analysis of the so-called Phillips curve. In the case of the Spanish economy, the trade-off between inflation and unemployment rates has been robustly established in papers like Dolado, López-Salido and Vega (2000), so we think that any model specification that attempts to explain the evolution of inflation should contain the unemployment rate among the set of explanatory variables.

[1] King and Ma (2001) and Neyapti (2004) and other papers mentioned above, used as measures of fiscal decentralitzation the weight that represents the volume of public spending (or taxes) for each level of government relative to the total public spending (or taxes) for all levels of government. As noted by Oates (1972), this is an imperfect measure of the level of fiscal decentralization, although in most cases these are the only ones that can be calculated.

[2] It should be noted that the analysis in Thornton (2007) is not strictly comparable with those of King and Ma (2001) and Neyapti (2004) because it does not consider the same set of countries. In fact, Thornton (2007) produced a series based on the application of the weights obtained in 1995 by the OECD. Such transformation does not solve the criticism to the work of King and Ma (2001) and Neyapti (2004), as by applying equal weights for each year what you are doing is changing the scale of the time series. If he would have estimated the model considering the same countries that King and Ma (2001) and Neyapti (2004) took into account, he would have obtained their exact same results. Consequently, the conflicting results of Thornton (2007) and King and Ma (2001) and Neyapti (2004) are simply due to the fact that they used a different set of countries and estimation techniques.

Finally and to the best of our knowledge, we have not found any study that analyzes the Spain case for a regional perspective, although the process of fiscal, institutional and political decentralization that has experienced Spain since the beginning of the democratic period has been very important. In this sense, the analysis that is carried out in this paper is a novel contribution insofar as we increase the set of empirical evidence available in the literature.

4 The empirical model

The specification of the empirical model that we use to analyze the relationship between fiscal decentralization and inflation bases on the proposal in Triesman (2000) and Martínez-Vázquez and McNab (2006). The model can be written in a general way as:

$$\pi_{i,t} = \gamma_i + \beta_1 AC_{i,t} + \beta_2 Local_{i,t} + x'_{i,t}\beta_3 + u_{i,t}, \tag{1}$$

where $\pi_{i,t}$ denotes the inflation rate, $AC_{i,t}$ reflects the level of fiscal decentralization at ACs government level, and $Local_{i,t}$ reflects the level of fiscal decentralization at local government level, $i = 1, \dots, 17$, $t = 1979, \dots, 2000$. Finally, the $(k \times 1)$-vector $x_{i,t}$ contains other determinants of inflation proposed in the literature. By working at the regional level means that we either do not have statistical information of some of the determinants that have been commonly used in the literature or do not make sense to use them. For example, since no information is available on the external sector of each autonomous community in Spain, we cannot estimate the degree of openness of the economies at a regional level – the absence of trade flows among the different regions prevents to obtain variables of the external sector for each region. Consequently, and given the restrictions imposed by the availability of statistical information concerning the ACs, $x_{i,t}$ collects the Gross Value Added per capita (GVApc) and the unemployment rate (U) of the Spanish ACs.

The time period that is analyzed in this paper is restricted by the availability of statistical information, and covers the years between 1979 and 2000. The Spanish regional autonomic system was created in 1978, so that the National Statistical Institute of Spain (INE) does not provide data on price levels at ACs level before this year. It is noteworthy that despite having information of price levels for the capital of the Spanish provinces from the second half of the thirties, it is not possible to build a series of price level at ACs level from the price series of the capital of the provinces, as it would no longer consider the influence of rural areas. Consequently, it is from 1978 on when data on ACs price levels is available, so that the series of inflation for the ACs begin in 1979. Figure 3 shows the evolution of the series in that period.

The second set of variables relevant to our study refers to the levels of fiscal decentralization. The fiscal decentralization indicators that have been traditionally used are defined the ratio of expenses (or revenues) of each level of government over the total expenditure (or revenue) made by all levels of government. Thus, the studies that have carried out international comparative use the database of the International Monetary Fund – Government Finance Statistics Annual Yearbook – to develop indicators of fiscal decentralization since the beginning of the seventies to the present.

Unfortunately, the Spanish statistical information on expenditures and tax revenues territorialized by regions does not cover a large time period, which prevents us to carry out the analysis of the effect of fiscal decentralization on stability macroeconomics. An alternative to the classical measure of fiscal decentralization is used in Carrion-i-Silvestre, Espasa and Mora (2008) when analyzing the impact of fiscal decentralization on economic growth of the Spanish ACs. These authors define indicators of fiscal decentralization based on the investment ratios of the different levels of government in each region. Thus, we define the fiscal decentralization variables as:

$$AC_{i,t} = \frac{I_{i,t}^{AC}}{I_{i,t}^{T}}; \qquad Local_{i,t} = \frac{I_{i,t}^{L}}{I_{i,t}^{T}},$$

where $I_{i,t}^{AC}$ denotes the investment made by the i-th autonomous community, $I_{i,t}^{L}$ the one made by the local governments of the i-th AC and $I_{i,t}^{T}$ the total investment that make the central, AC and local governments on the i-th AC. As usual, the investments include investments in roads, railways, airports, ports, hydraulic structures and education. We have also taken into account the investment in health when such competence lies in the government of the autonomous region. However, it has been considered that competences in health care were transferred to some regional governments in different times, which required the implementation of some adjustments to obtain ratios of investment.[3] It should be noted that the use of investment ratios as an indicator of the level of fiscal decentralization makes it possible to capture only a portion of government spending, but it is the most productive one. The computation of these fiscal decentralization indicators are based on the territorialized public investment series by concepts developed by the Valencian Institute of Economic Research (IVIE). Unfortunately, IVIE has disrupted the production of these series so that our analysis period inevitably ends in 2000, the year for which the latest data on territorialized public investment by concepts is available. Figures 4 and 5 show the evolution of the indicators of fiscal decentralization in the period 1964-2000 at the AC and local levels, respectively.

GVApc series in real terms have been obtained from the work of Doménech, Escribá and Murgui (1999) and the BDMores database of the Spanish Ministry of Economy and Finance, which allow to cover the period 1964-2000. Figure 6 shows the temporal evolution. Finally, the unemployment rate is obtained from IVIE and covers the period 1964-2001 – see Figure 7. See the appendix for more details on the definition of variables and statistical sources that have been used.

As can be seen, the graphs show a clear trend behavior for inflation, unemployment, the level of fiscal decentralization to the ACs and the GVA per capita. In the case of fiscal decentralization for the local level the trending behavior is less clear. These characteristics are important for the analysis that is carried out in the following sections, where we require to define the deterministic component that is used when computing the unit root tests statistics that are applied below. For the variables of inflation, GVA per capita, unemployment and fiscal decentralization in terms of ACs the preferred deterministic specification is the linear trend. For the variable of

[3] Between parenthesis, we list the years in which competence in health at regional level were effective for each AC: Andalusia (1985), Canary Islands (1995), Catalonia (1982), Galicia (1991), Valencia (1988), Basque Country (1988) and Navarra (1991). From this time forward investment in health has been added to the investment made by the corresponding autonomous community instead of being regarded as an investment of the central government.

fiscal decentralization at the local level the trending patters is not so clear, so that we consider both the constant and the linear time trend deterministic specifications.

5 Panel data integration and cointegration analyses

Economic variables are 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 processes that are governed by stochastic trends, i.e., integrated processes $I(d)$ with $d > 0$. This is a relevant issue, since the validity of the estimates 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.

In the literature there are many papers that have characterized the series of inflation, unemployment and output variables as I(1) stochastic processes, i.e., non-stationary variables. These are just some of the variables involved in the model specified in the previous section, so some caution should be taken if the dangers of a spurious relationship are to be avoided. Surprisingly, the papers cited in Section 3 assume that the variables are I(0) stationary processes, which does not seem to be in agreement with other studies that examine the order of integration of these variables.

By working with I(1) non-stationary stochastic processes does not necessarily mean that we have to face a spurious relationship when estimating the model given by (1). Thus, it is possible that relationships between I(1) variables lead to consistent estimates of the parameters if the variables generate a cointegration relationship. Therefore, the analysis should proceed determining, first, the order of integration of the variables involved in the model and, second, proceed to test 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 dealing with stationary panel data models.

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.

5.1 Panel data cross-section dependence

In this section we compute the test statistics in Pesaran (2004, 2012), 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 if 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.

Let $y_{i,t}$ denote a generic macroeconomic variable – for instance, the inflation rate or the GVA per capita – for which the following autoregressive ($\mathrm{AR}(p_i)$) model is specified:

$$y_{i,t} = \mu_i + \beta_i t + \phi_{1,i} y_{i,t-1} + \cdots + \phi_{p_i,i} y_{i,t-p_i} + e_{i,t}, \tag{2}$$

$i = 1, \ldots, N$, $t = 1, \ldots, T$, where the inclusion of a linear time trend depends on whether the time series shows a trending behaviour – see the comments above. The test statistic in Pesaran (2004, 2012) is based on the average of pair-wise Pearson's correlation coefficients $\hat{p}_j$, $j = 1, 2, \ldots, n$, $n = N(N-1)/2$, of the ordinary least squares (OLS) residuals $\hat{e}_{i,t}$ obtained from the estimation of the $\mathrm{AR}(p_i)$ model given in (2) – the estimation of an $\mathrm{AR}(p_i)$ model aims at wiping out the effect that the time dependence might have on the cross-section dependence among the unit of the panel.[4] The test statistic in Pesaran (2004, 2012) is given by:

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

where under the null hypothesis of cross-section independence the WCD statistic converges to the standard normal distribution. Pesaran (2004, 2012) 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:

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

which under the null hypothesis of cross-section independence converges to the standard normal distribution. Pesaran (2012) 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 – i.e., strong rejection of the null hypothesis of cross-section independence – 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

[4] The order of the autoregressive model (p_i) is selected using the t-sig criteria in Ng and Perron (1995) allowing for a maximum of $p^{\max} = 10$ lags.

around one indicates that strong dependence is present – the computation of this statistic can only be done for non-trending regressors.

The procedure proposed by Ng (2006) works as follows. As for the previous statistics, we get 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 ($\bar{p}_j = |\hat{p}_j|$) for all possible pairs of individuals, $j = 1, 2, \ldots, n$, where $n = N(N-1)/2$, and sort them in ascending order. As a result, we obtain the sequence of ordered statistics given by $\left\{\bar{p}_{[1:n]}, \bar{p}_{[2:n]}, \ldots, \bar{p}_{[n:n]}\right\}$. Under the null hypothesis that $p_j = 0$ and assuming that individual time series are Normally distributed, $\bar{p}_j$ is half-normally distributed. Furthermore, let us define $\bar{\phi}_j$ as $\Phi\left(\sqrt{T}\bar{p}_{[j:n]}\right)$, where Φ denotes the cdf of the standard Normal distribution, so that $\bar{\phi} = \left(\bar{\phi}_1, \ldots, \bar{\phi}_n\right)$. Finally, let us define the spacings as $\Delta\bar{\phi}_j = \bar{\phi}_j - \bar{\phi}_{j-1}$, $j = 1, \ldots, n$.

Ng (2006) proposes splitting the sample of (ordered) spacings at arbitrary $\vartheta \in (0,1)$, so that we can define the group of small (S) correlation coefficients and the group of large (L) correlation coefficients. The definition of the partition is carried out by minimizing the following sum of squared residuals:

$$Q_n(\vartheta) = \sum_{j=1}^{[\vartheta n]} \left(\Delta\bar{\phi}_j - \bar{\Delta}_S(\vartheta)\right)^2 + \sum_{j=[\vartheta n]+1}^{n} \left(\Delta\bar{\phi}_j - \bar{\Delta}_L(\vartheta)\right)^2,$$

where $\bar{\Delta}_S(\vartheta)$ and $\bar{\Delta}_L(\vartheta)$ denotes the mean of the spacings for each group, respectively. A consistent estimate of the break point is obtained as $\hat{\vartheta} = \arg\min_{\vartheta\in(0,1)} Q_n(\vartheta)$, where some trimming is required – following Ng (2006), the trimming is set at 0.10. Once the sample has been split, we can proceed to test the null hypothesis of non correlation in both sub-samples. Obviously, the rejection of the null hypothesis for the small correlations sample will imply also rejection for the large correlations sample as the statistics are sorted in ascending order. Therefore, the null hypothesis can be tested for the small, large and the whole sample using the Spacing Variance Ratio (SVR) in Ng (2006):

$$SVR(\eta) = \hat{\sigma}_q^2/\hat{\sigma}_1^2 - 1,$$

where $\hat{\sigma}_q^2 = (m_q q)^{-1}\sum_{k=q+1}^{\eta}\left(\bar{\phi}_k^n - \bar{\phi}_{k-q}^n - \hat{\mu}_q\right)^2$, $\hat{\sigma}_1^2 = (\eta)^{-1}\sum_{k=1}^{\eta}\left(\bar{\phi}_k^n - \bar{\phi}_{k-1}^n - \hat{\mu}_1\right)^2$, with $m_q = \eta - q$ and where $\hat{\eta} = [\hat{\vartheta} n]$ denotes the number of small correlations. Ng (2006) shows that under the null hypothesis that all correlations are zero, $\sqrt{\eta}SVR(\eta) \sim N\left(0, \omega_q^2\right)$, $\omega_q^2 = 2(2q-1)(q-1)/(3q)$, as $T \to \infty$. Using these elements, we can define the standardized test statistic $svr(\eta) = \sqrt{\eta}SVR(\eta)/\omega_q$, which converges to the standard normal distribution.

Table 1 presents the results of calculating the WCD and SVR statistics for each panel data, considering both deterministic functions of interest. 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 $\mathring{\delta}$ is close to one in all cases except for the AC variable. The 90% confidence interval defined by $(\mathring{\delta}_L, \mathring{\delta}_U)$ indicates that, except for the AC, the cross-section dependence is strong.

The SVR statistic in Ng (2006) shows that the null hypothesis of no correlation cannot be rejected at a significance level of 5% when working with small sub-sample correlations – 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 $svr(L)$ statistic. It should also be noted that the large sub-sample of correlations is largely more numerous than the small correlations one, which indicates that, first, there is evidence of strong cross-section correlation and, second, that the correlation is pervasive – see Ng (2006). 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) correlated. This conclusion is supported in an informal way when analyzing the pictures of the variables, where similarities in the evolution of the series of inflation, unemployment, fiscal decentralization and GVA per capita are easily observed. Either formally or informally, this feature indicates that the statistical inference has to consider the presence of cross-section dependence if meaningful conclusions are to be obtained.

5.2 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 – this was a recommendation made in 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. The 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 (the same value for the factor loadings) on all units. A second approach is proposed by Maddala and Wu (1999), who suggest to obtain 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, whatever it is. The drawback of this approach is that we cannot get an explanation of the source or form of the cross-section dependence. Finally, a third approach models the cross-section dependence by specifying an approximate common factor model. The common factors are a simple device to capture the dependency structure, allowing the possibility of an ulterior identification of the sources that drive such common factors. In this section and provided the conclusions obtained above, we opt for the third option, applying panel unit root test statistics that incorporate unobservable common factors as a device 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 approaches, we proceed to briefly describe here this procedure. 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 $e_{i,t}$:

$$\begin{aligned} y_{i,t} &= D_{i,t} + \lambda_i' F_t + e_{i,t} && (3) \\ (1-L)\, F_{j,t} &= C_j\,(L)\, w_{j,t}; \quad j = 1\ldots, r && (4) \\ (1-\rho_i L)\, e_{i,t} &= H_i\,(L)\, \varepsilon_{i,t}, && (5) \end{aligned}$$

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 (importance) 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 stationary or non-stationary common factors (F_t) using the ADF (for the one common factor case) or the MQ test statistics in Bai and Ng (2004) (for the general case where there are more than one common factor) – 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. Therefore, 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. If the model has an intercept and a linear trend that 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, if this is the case. 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 case 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), 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.

As can be seen, this technique allows us to determine the possible source of the stochastic trends. It is possible that non-stationarity is the result of the presence of I(1) common factors – or a combination of I(0) and I(1) common factors. This will imply that panel data is non-stationary because there are global stochastic trends affecting the observable variables. In this case, we can conclude that there is permanent global shocks that affect all series of the panel data. It could also be the case that the idiosyncratic disturbance terms were I(1) non-stationary processes, which would imply that the shocks that affect each time series (not global shocks) have a permanent effect. This defines idiosyncratic stochastic trends.

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 for values of $p > 2$ the result that is obtained leads to the non-rejection of the null hypothesis of unit root at the 5% level of significance, regardless of the deterministic specification that is chosen.

Table 3 reports the results of the test statistics proposed by Moon and Perron (2004) – denoted by t_a and t_b. We can see that the null hypothesis of unit root is clearly rejected for the inflation, GVApc, AC (using the deterministic trend), Local, and not rejected for the unemployment rate at the 5% level of significance.

These results are not conclusive, since in some cases we found that the variables are I(0) and in other cases they are characterized as I(1) stochastic processes. However, evidence from the statistics in Moon and Perron (2004) and Pesaran (2007) can be biased because of assuming that the dynamics of the common factors and the idiosyncratic disturbance term is the same. The procedure in Bai and Ng (2004) overcomes this limitation provided that it conducts a separate analysis of the common factor and idiosyncratic components.

Table 4 presents the test statistics of Bai and Ng (2004). The conclusion that is obtained from these statistics is that all observable variables are I(1) non-stationary variables since in all cases we detect the presence of I(1) common factors. Therefore and regardless of the stochastic properties presented by the idiosyncratic disturbance terms, the panel data sets that appear in (1) are I(1). If we focus on the idiosyncratic component, the ADF panel data statistic in Bai and Ng (2004) indicates that the null hypothesis of panel unit root cannot be rejected at the

5% level of significance for the inflation rate, the unemployment rate, the GVApc and the fiscal decentralization at the AC level, while it is rejected for fiscal decentralization at the local level. This result indicates that idiosyncratic shocks have a permanent effect on the inflation rate, the unemployment rate, the GVApc and the fiscal decentralization at the AC level, while the character is temporary for the fiscal decentralization at the local level.

5.3 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 relates inflation and fiscal decentralization is a long-run relationship (an equilibrium relationship with economic sense) 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 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.

The proposal in Bai and Carrion-i-Silvestre (2013) can also envisage that common factors that affect the dependent variable (inflation in our case) may also be affecting some of the explanatory variables of the model (especially unemployment). The model is specified as:

$$\pi_{i,t} = \gamma_i + \beta_{1,i} AC_{i,t} + \beta_{2,i} Local_{i,t} + x_{i,t}' \beta_{3,i} + F_t' \lambda_i + e_{i,t}, \tag{6}$$

where F_t denotes a $(r \times 1)$-vector of unobservable common factors, λ_i is a $(r \times 1)$-vector of factor loadings, $e_{i,t}$ denotes the idiosyncratic disturbance term, and $x_{i,t} = GVApc_{i,t}$ or $x_{i,t} = (GVApc_{i,t}, U_{i,t})'$, depending on whether we include the unemployment rate among the explanatory variables. The common factors can be I(0) stationary or I(1) non-stationary stochastic processes, although in order for the cointegration relationship to exist, the idiosyncratic disturbance terms must be I(0) stationary stochastic processes.

This framework allows us to give a richer interpretation to the usual definition of cointegration. Thus, the common factors can be used to capture misspecification errors in the models due, for instance, to the omission of relevant variables. The papers that have analyzed the relationship between fiscal decentralization and macroeconomic stability have focused on the international scope and have been able to include explanatory variables that are not available when working at a regional level – for example, variables such as the independence of the central bank, the interest rate of the economy, or whether the currency exchange rate is fixed or variable, cannot be included in our model as they are the same for all regions of the economy.

The consistent estimation of the common factors is carried out working with the model given in (6) in first difference, as suggested in Bai and Ng (2004). However, the fact that the common factors that affect the dependent variable may also be affecting (some of) the explanatory variables implies that all elements of the model – i.e., the common factors, the factor loadings and the parameters of the cointegrating vector – have to be estimated using the iterative estimation method proposed in Bai and Carrion-i-Silvestre (2013). Further, this approach enables us to specify heterogeneous or homogeneous parameters, which defines an environment flexible enough to collect all the casuistry of the Spanish autonomous communities.[5] This iterative estimation method gives consistent estimates of the parameters, common factors and idiosyncratic disturbance terms of the model. The inference about whether the relationship is a spurious or a cointegration relationship needs to focus on the idiosyncratic disturbance terms, as if these processes are characterized as I(0) stationary processes, then we will conclude that we are facing a cointegration relationship, and otherwise it will be a spurious relationship.

Table 5 presents the results of computing the panel data cointegration test statistics in Bai and Carrion-i-Silvestre (2013), all based on the modified Sargan-Barghava (MSB) test statistic – the three test statistics are denoted as Z, P_m and P and test the null hypothesis of no panel cointegration against the alternative hypothesis of panel cointegration. Based on the results from simulation experiments, Bai and Carrion-i-Silvestre (2013) conclude that P_m and P are the preferred test statistics provided that they show a better finite sample performance.

We have computed the test statistics for two possible estimates of the relationship between inflation and decentralization. The existing papers in the area of fiscal decentralization and macroeconomic stability ignore the unemployment rate as a possible explanatory variable, although,as pointed out above, there are other fields of the economic analysis where a clear relationships between inflation and unemployment is argued. Therefore and in order to conduct a robustness check of our analysis, we have essayed the sensitivity of the conclusions depending on whether the unemployment rate is included in the model as an additional explanatory variable.

The results presented in Table 5 are provided for different values of the number of common factors that are specified – when the panel BIC information criterion proposed by Bai and Ng (2002) is used the estimated the number of common factors coincides with the maximum value of common factors that have been allowed (six in our case). As can be seen, when the model does not include the unemployment rate among the explanatory variables, the P_m and P test statistics reject the null hypothesis of no cointegration at the 5% level of significance for $r = 2$, 3 and 5 factors – it can be seen that when specifying six common factors the level of significance which would lead to reject the null hypothesis of spurious relationship is 12%, a value not too far from the 10%. In any case the values of the Z statistic do not allow to reject the null hypothesis of spurious regression, although, as mentioned above, P_m and P are preferable to the Z statistic.

The inclusion of the unemployment rate among the set of explanatory variables increases the evidence in favor of panel cointegration, since the P_m and P test statistics reject the null hypothesis of no cointegration at the 5% level of significance for $r = 2$, 4, 5 and 6 factors – with one common factor the p-value is 12%, not too far from the 10% level of significance. As before,

[5] As can be seen, the parameters in (6) include a subscript for the i-th individual, which lets you specify a large heterogeneity degree when estimating the model. However, the estimation of the parameters can also be obtained imposing the homogeneity constraint on the $\beta_{1,i}$, $\beta_{2,i}$ and $\beta_{3,i}$ parameters.

Z statistic does not allow to reject the null hypothesis of spurious regression in any case.

When the parameter homogeneity restriction is imposed, the evidence in favor of panel data cointegration is reduced. Table 6 shows that for the model specification that does not include the unemployment rate among the explanatory variables, the P_m and P test statistics lead to reject the null hypothesis of no panel cointegration when $r = 2$ and 4 at the 5% level of significance, and for $r = 3$ at the 10% level of significance. When the unemployment rate is included in the model, the null hypothesis of no panel cointegration is clearly rejected by the P_m and P test statistics at the 5% level of significance for $r = 2$ and 3.

To sum up, the analysis that has been conducted in this section indicates that there is evidence of panel cointegration, although this evidence is stronger when the unemployment rate is included in the model specification.

6 Macroeconomic stability and fiscal decentralization

Once evidence of panel cointegration has been found, we move on the estimation of the cointegrating vector. In this section we consider different estimation procedures. First, we use the panel common correlated effects (CCE) pooled estimator proposed in Kapetanios et al. (2011). Since the cross-section dependence is sometimes caused by unobservable common factors, Pesaran (2006) uses cross-section averages of the dependent and the explanatory variables as proxies for common factors. This approach suggest the estimation of the following model:

$$\pi_{i,t} = \gamma_i + \beta_1 AC_{i,t} + \beta_2 Local_{i,t} + x_{i,t}'\beta_3 + \bar{z}_t'\eta_i + e_{i,t}, \tag{7}$$

where $\bar{z}_t = (\bar{\pi}_t, \overline{AC}_t, \overline{Local}_t, \bar{x}_t')'$ is the vector of cross-section means of the dependent and explanatory variables, and $x_{i,t} = GVApc_{i,t}$ or $x_{i,t} = (GVApc_{i,t}, U_{i,t})'$, depending on whether we include the unemployment rate among the explanatory variables. Following Pesaran (2006), Holly et al. (2010) and Kapetanios et al. (2011), the estimation of the $\beta = (\beta_1, \beta_1, \beta_3')'$ parameters in (7) bases on the use the following pooled estimator:

$$\hat{\beta}_{CCEP} = \left(\sum_{i=1}^{N} z_i'\bar{M}z_i\right)^{-1}\left(\sum_{i=1}^{N} z_i'\bar{M}\pi_i\right),$$

where $z_i = (z_{i,1}, z_{i,2}, ..., z_{i,T})'$, $z_{i,t} = (AC_{i,t}, Local_{i,t}, x_{i,t}')'$, $\pi_i = (\pi_{i,1}, \pi_{i,2}, ..., \pi_{i,T})'$ and the matrix $\bar{M}$ is defined in Holly et al. (2010). The pooled CCE estimator assumes that the stochastic regressors are weakly exogenous, in which case it is shown that the pooled CCE estimator provides consistent estimates of the parameters and statistical inference can be carried out using the standard test statistics – for instance, in the limit, the t-ratio test statistics that are used for testing the individual significance of the parameters converge to the standard normal distribution.

We have also estimated the parameters of the model following the strategy in Bai, Kao and Ng (2009) and Bai and Carrion-i-Silvestre (2013), since this technique allows us to consistently estimate the panel data cointegrating vector when there are common factors that capture the cross-section dependence. In this case, the model that has been specified takes the form:

$$\pi_{i,t} = \gamma_i + \beta_1 AC_{i,t} + \beta_2 Local_{i,t} + x_i'\beta_3 + F_t'\lambda_i + e_{i,t}, \tag{8}$$

with $x_{i,t}$ defined above. This procedure estimates the cointegrating vector, the common factors and the factor loadings in an iterative fashion, so it is possible to take into account the fact that the common factors that affect the dependent variable can also be affecting the stochastic regressors of the model – the maximum number of common factors that have been allowed in the analysis is six, and the panel BIC information criterion in Bai and Ng (2002) has selected the maximum in all cases. As can be seen, one important difference between (7) and (8) is the way in which the cross-section dependence is accounted for. However, there is also a crucial feature that distinguish both approaches, since the proposals in Bai, Kao and Ng (2009) and Bai and Carrion-i-Silvestre (2013) allow the stochastic regressors to be endogeneous.

In order to conduct statistical inference about the estimated parameters, we have applied the dynamic ordinary least squares (DOLS) estimation method as suggested in Bai and Carrion-i-Silvestre (2013), which consists on adding leads and lags of the first difference of the stochastic regressors as additional explanatory variables – the selection of the number of leads and lags is based on the BIC information criterion considering all possible combinations of model specifications that can be formed with a maximum of two leads and lags. The advantage of using DOLS as estimation method is that it provides optimal estimates of the parameters, as it takes into account the possible endogeneity that may exist in any of the stochastic regressors. The characteristic of optimality of the estimates allows us to apply standard inference on the estimated parameters, since they are asymptotically distributed as a mixture of normal distributions. Therefore, the use of the standard normal distribution sets an upper bound when analyzing the statistical significance of the individual parameters. Finally, we have applied DOLS instead of the fully-modified OLS (FM-OLS) estimation procedure used in, for instance, Bai, Kao and Ng (2009) because the performance in finite samples of the former is superior to that of FM-OLS.

Table 7 presents the results of estimating models (7) and (8) for various alternative specifications. Let us first focus on the CCE estimates. As can be seen from columns (a) and (b) of Table 7, none of the estimated parameters are statistical significant, a surprising result provided that panel cointegration has been found. However, we should bear in mind that the pooled CCE estimator assumes that the stochastic regressors are weakly exogenous, an assumption that might not hold in this case. In order to carry out an analysis that is robust from the possible endogeneity of the stochastic regressors, we focus on the parameters that have been estimated using DOLS.

Column (1) of Table 7 presents the DOLS estimates of the model specification that incorporates GVApc, AC and Local as explanatory variables. As you can see, all parameters are statistically significant at the 5% level of significance. We can see that the sign of the parameter corresponding to GVApc is positive and the parameters associated with the fiscal decentralization variables are negative. Therefore, this initial evidence suggests that fiscal decentralization can reduce the level of inflation, helping to macroeconomic stability. Column (2) of Table 7 collects the estimates for the model that includes the unemployment rate as an additional explanatory variable. The conclusions obtained with the estimate from the previous model are not modified in qualitative terms, although we see that the corresponding parameter for the unemployment rate is negative and highly statistical significant. This indicates that many of the models that have been used so far to study the possible relationship between inflation and fiscal decentralization may have been facing a misspecification error. Moreover, the negative sign of the parameter is coherent with the evidence found in other empirical studies for the Spanish economy where a negative relationship between inflation and unemployment has been obtained – see Dolado, López-Salido and Vega (2000).

From these two initial estimates we can see that fiscal decentralization to local governments has had a major effect in ensuring price stability. However, the heterogeneity of ACs in terms of the competences acquired over time, on the one hand, and the pace to put them into practice, on the other hand, can be taken into account when estimating the models. Thus, we can incorporate dummy variables that interact with the indicators of fiscal decentralization as a way to disentangle the effect that the different degree of competence acquisition has on the macroeconomic stability. To address this feature, we define the following dummy variables: D1 indicates those ACs ruled by Article 143 of the Spanish Constitution, D2 designates those ACs with competences defined in Article 151 of the Spanish Constitution and, D3 is a dummy variable that identifies the foral autonomous communities of the Basque Country and Navarre. Finally, we have also taken into account the existence of ACs with a single province through the definition of the dummy variable D4.

Columns (3) to (5) of Table 7 provide the parameter estimates of the specifications that include interactions with the dummy variables. As can be seen, the sign and statistical significance of the parameters associated with the GVApc and the unemployment rate do not vary significantly compared with the previous estimates. The most notable difference refers to the qualitative effects of fiscal decentralization, as the magnitude of the coefficients varies significantly depending on the type of region.

Consider first the estimates reported in column (3) of Table 7. Now the fiscal decentralization at the AC level has a positive and significant effect on inflation if the autonomous community belongs to the group of the Article 143 ACs. However, the ACs with higher level of competences – those of the article 151 of the Spanish constitution and the foral ones – see how fiscal decentralization at the AC level helps ensure price stability, since the associated parameters are negative and statistical significant. On the other hand, fiscal decentralization at the local level has a negative and statistical significant impact on inflation, although its magnitude depends on the type of AC – the ACs with higher competences are the ones where the fiscal decentralization at the local level has more impact on inflation.

Columns (4) and (5) of Table 7 report the estimation of the model with the D4 dummy variable. The estimates in column (4) reveal that two parameters are not statistical significant, so we proceed to re-estimate the model excluding such regressors, providing the results that appear in column (5). The inclusion of the dummy variable D4 has two consequences. First, the fiscal decentralization at the AC level turns not statistical significant for the Article 143 ACs. Second, the fiscal decentralization at the AC level has a positive and significant effect for the ACs with a single province.

Regarding the effects of fiscal decentralization at the local level, we see that the estimates do not vary significantly in relation to the aforementioned. Therefore, we conclude that fiscal decentralization at the local level helps to control inflation, with a greater effect on inflation for the Article 151 ACs, followed by the foral ACs and, finally, the Article 143 ACs. Figures 8 and 9 present a summary of the estimated effect of fiscal decentralization for the AC and Local levels for the different ACs – these effects base on the estimates reported in column (5) of Table 7. As can be seen, the ACs that benefited more in terms of fiscal decentralization are the Basque Country, followed by Andalusia, the Canary Islands, Catalonia, Valencia, Galicia, and, lastly, by Navarre. The effect that has been found for these ACs is consistent with the commitment hypothesis, where central government fiscal policy actions would be controlled by regional governments, so that it becomes more difficult for the central government to violate its commitment to monetary stability. Another possible effect to consider is the attitude of these regional governments for auditing the performance of central government fiscal policy. Controlling, for example, the final assignment of public spending that the central government transfers to the ACs could help to control the price level and keep it stable – if the expenditure results in an increase in the productivity of the economy such that prices can remain stable. The estimates presented in column (5) of Table 7 indicate that a 1% increase in the level of fiscal decentralization at the AC level represents a 1.161% or a 2.597% reduction in the rate of inflation, depending on the type of AC.

For the rest of ACs, fiscal decentralization at the AC level has either no impact on inflation or a positive effect. As mentioned above, if fiscal decentralization has no impact on inflation, we face a situation where levels of inflation (high or low) will tend to persist. Finally, the case where the relationship between inflation and fiscal decentralization is positive is consistent with the problem of collective action, where inflation is not a problem that affects the economy of the ACs, but a general problem affecting the Spanish economy in general.

In all cases we found that fiscal decentralization at the local level helps to keep inflation controlled, showing that the control that local authorities have on the central and regional governments actions helps to reflect spending on activities that do not generate inflation. Here we see that the effects of a 1% increase of the fiscal decentralization at the local level represents a 0.704, 1.664 or 1.416% of inflation reduction, depending on the type of AC.

To sum up, our study has shown that fiscal decentralization helps to reduce inflation in those ACs with more competences transferred. In general, the heterogeneity of the Spanish ACs indicates that the effect of fiscal decentralization at all levels of government can have clear opposed effects depending on the degree of transferred competences of the ACs.

7 Conclusions

The present study has examined whether the relationship between fiscal decentralization and macroeconomic stability focusing on the Spanish regions. The interest of our approach lies in the fact that the Spanish economy has experienced an institutional decentralization process during the democracy period that may have helped to stabilize the economies of each region.

The survey of the economic literature on this topic that we have carried out in the paper reveals the interest on this issue, although most of the contributions are done at an international scope. Our study has pointed out the lack of empirical evidence documenting the case of Spain, making the analysis that is conducted in this paper even more attractive. Apart from the interest that the analysis of the case of Spain could represent, our proposal is based on the use of non-stationary panel data analysis, a methodology so far not been applied in this area, neither on international nor on national studies. The application of these techniques is important if meaningful conclusions are to be obtained from the estimated models. The paper has investigated whether the specification suggested by the economic literature defines a long-run equilibrium relationship (cointegration relationship) or not (spurious relationship). Of special relevance is the issue of the cross-section dependence affecting the units of the panels, something to be expected when we are analyzing regional data that belong to the same country.

The paper evidences that there is indeed a long-run relationship between fiscal decentralization and inflation for the Spanish autonomous communities, although the magnitude and sign of the effect of the degree of fiscal decentralization on the inflation rate depends on the type of autonomous community. It has been shown that the fiscal decentralization at the local level leads to control the inflation, the effect being higher for those ACs with higher competences transferred. The effect of the fiscal decentralization at the AC level is more heterogeneous, since the impact may be negative, positive or null, depending on the type of AC.

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A Appendix. Statistical sources

The data base is defined using the following macroeconomic variables:

- Inflation rate: annual variation in the consumer price index for CCAA base 1992. Source: National Statistics Institute (INE)
- Unemployment rate: percentage of the total unemployed population in the ACs. Source: Valencian Institute of Economic Research (IVIE)
- GVApc: gross value added per capita in constant 1980 million euros. Source: Doménech, Escribá and Murgui (1999) and BDMores database
- AC: ratio of investment carried out by the AC governments (in constant euros in 1980) over the total investment (in constant euros in 1980) made by all government levels. Source: IVIE
- Local: ratio of investment carried out by the local governments (in constant euros in 1980) over the total investment (in constant euros in 1980) made by all government levels. Source: IVIE
- D1: dummy variable that designates those ACs with competences defined by Article 143 of the Spanish Constitution
- D2: dummy variable that designates those ACs with competences defined by Article 151 of the Spanish Constitution
- D3: dummy variable that designates the foral ACs
- D4: dummy variable that designates the single province ACs

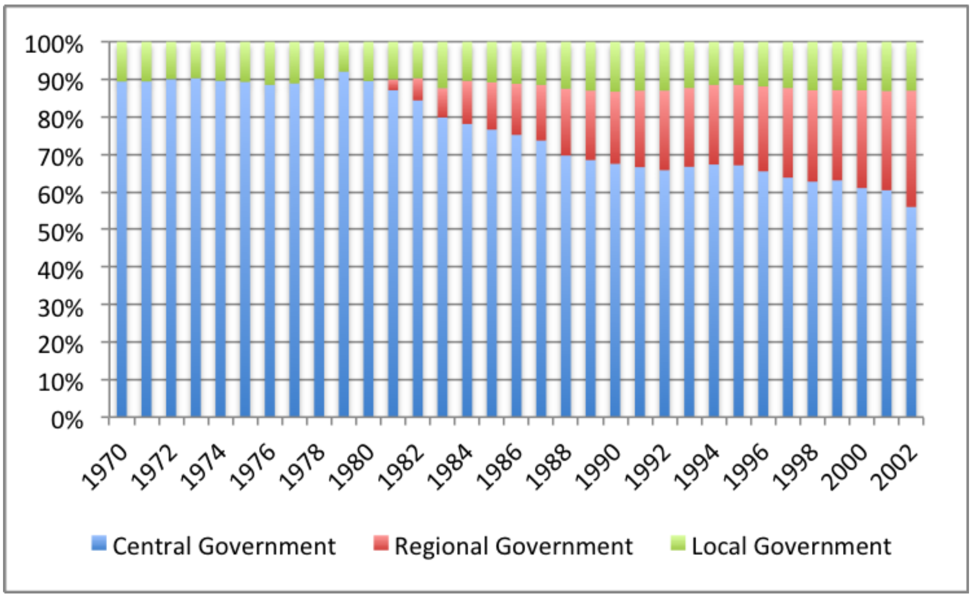

Figure 1: Evolution of public expenditure by level of government

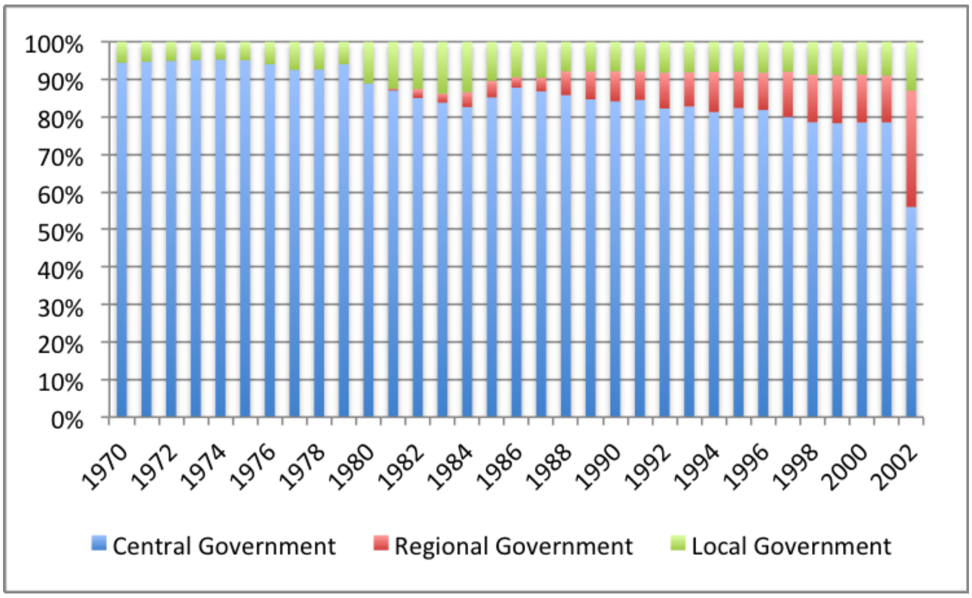

Figure 2: Evolution of public revenues by level of government

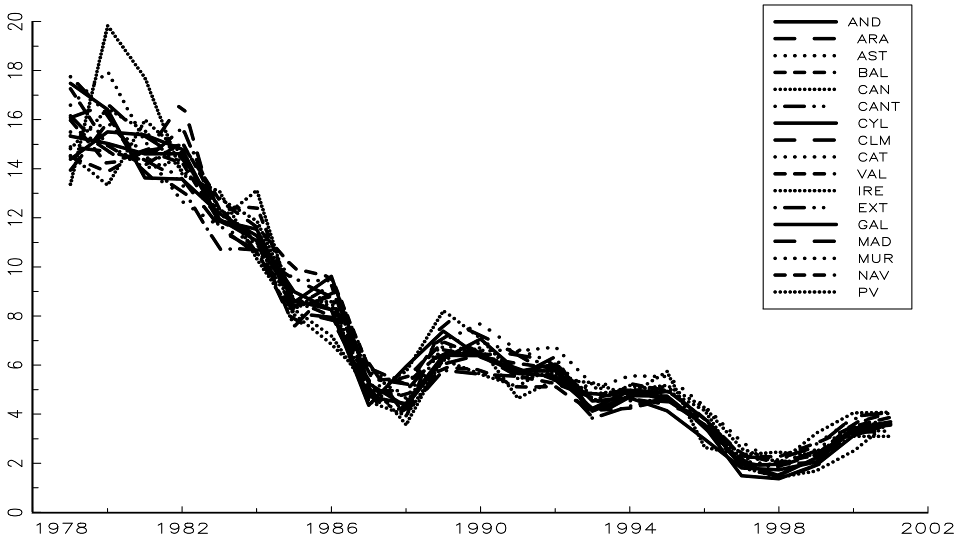

Figure 3: Inflation rate of the Spanish ACs

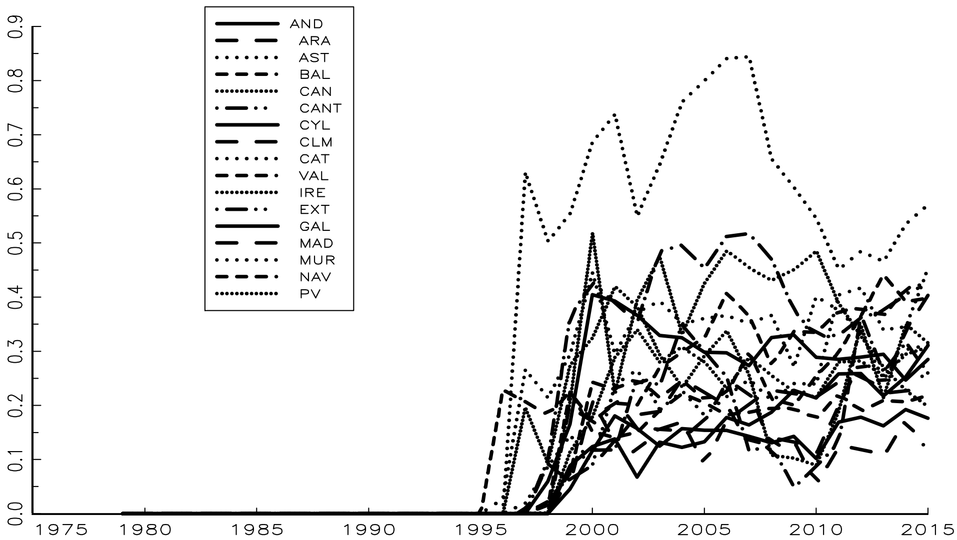

Figure 4: Fiscal decentralization in terms of Spanish ACs

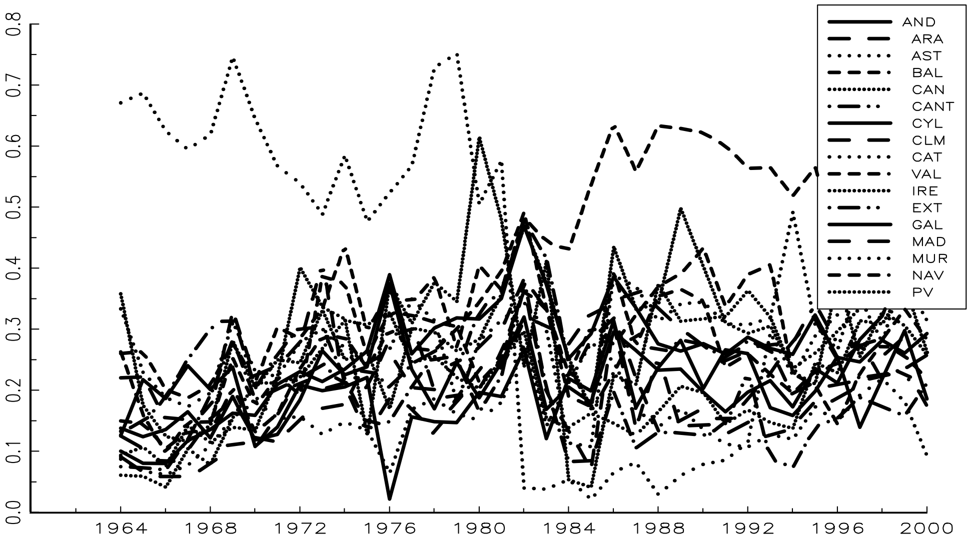

Figure 5: Fiscal decentralization in terms of Spanish local governments

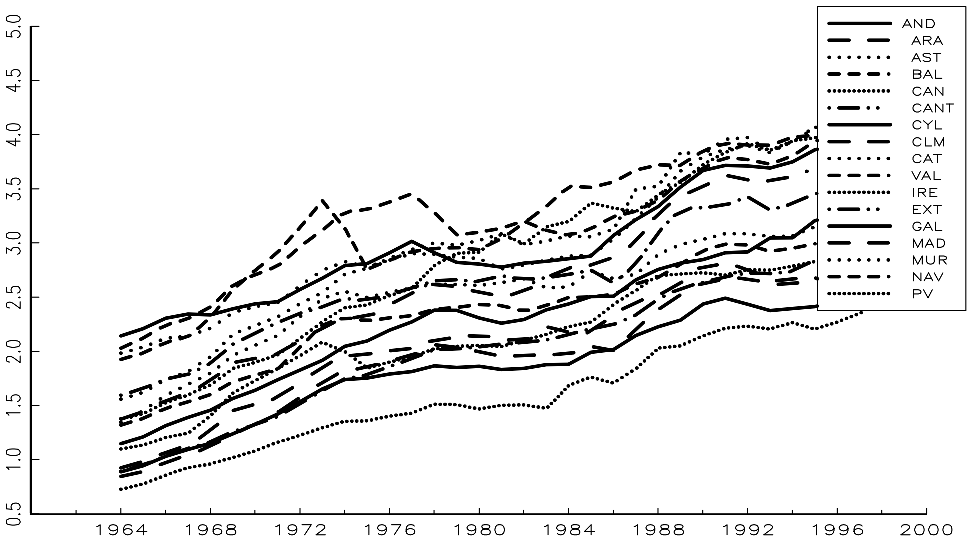

Figure 6: GVA per capita of the CCAA

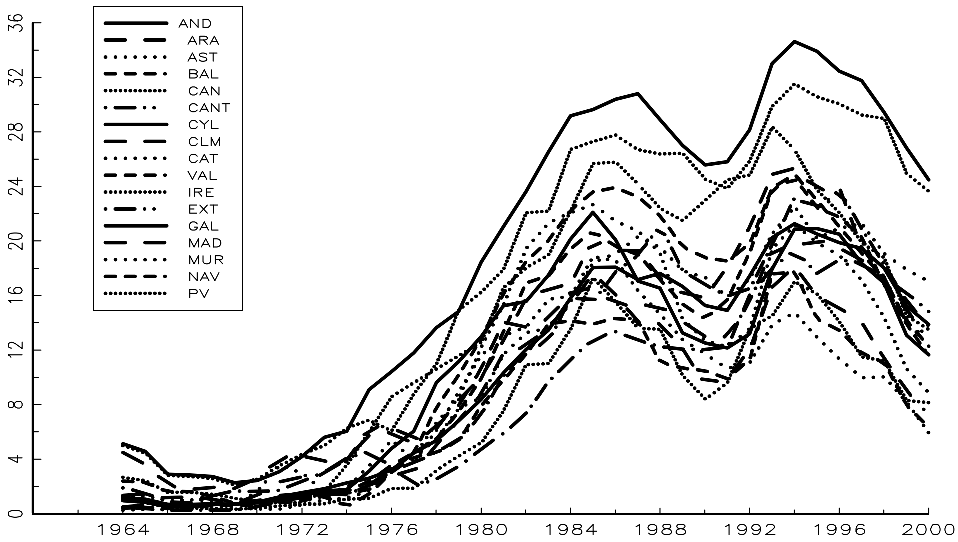

Figure 7: Unemployment rate of the ACs (in percentage)

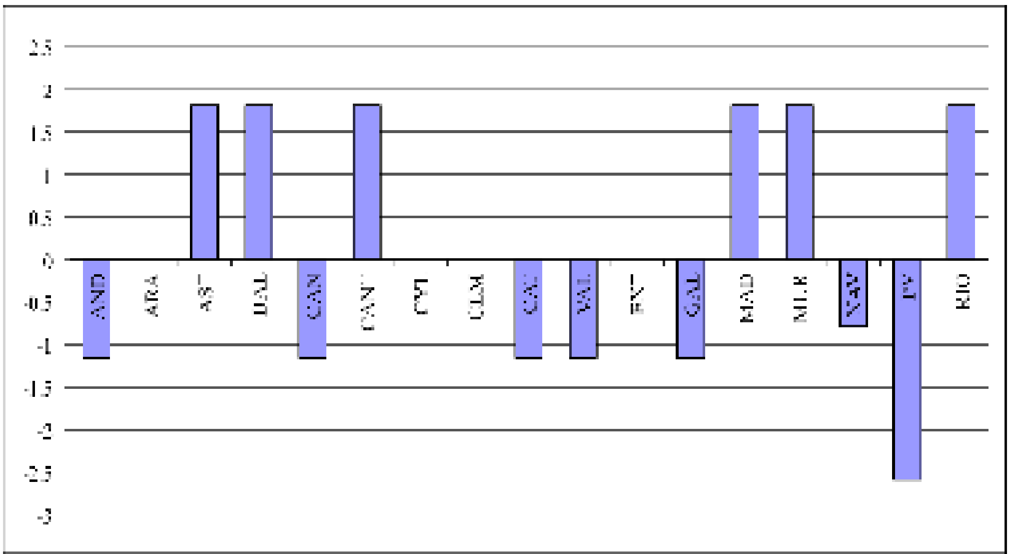

Figure 8: Effects of fiscal decentralization at the AC level

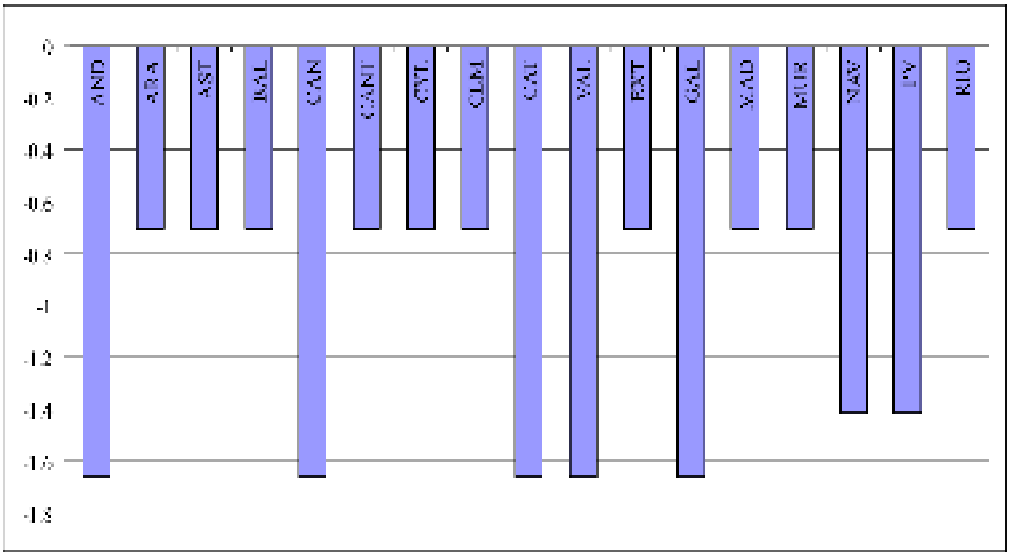

Figure 9: Effects of fiscal decentralization at the local level

Table 1: Cross-section dependence

Deterministic specification: Intercept

	Ng (2006)							Pesaran (2004, 2012)		Bailey et al. (2012)		
	$svr(W)$	p-value	$svr(S)$	p-value	$svr(L)$	p-value	η	WCD_{LM}	WCD	$\mathring{\delta}$	$\mathring{\delta}_L$	$\mathring{\delta}_U$
Inflation	3.232	0.001	-0.106	0.542	3.842	0.000	0.140	66.026	27.539	1.101	0.886	1.134
Unemployment	3.102	0.001	0.231	0.409	1.906	0.028	0.103	62.149	32.624	1.005	0.899	1.112
AC	1.594	0.055	-0.785	0.784	2.242	0.012	0.324	11.362	6.180	0.483	0.467	0.498
Local	1.175	0.120	-0.165	0.565	1.828	0.034	0.228	23.513	13.705	0.846	0.760	0.931

Deterministic specification: Linear trend

	Ng (2006)							Pesaran (2004, 2012)	
	$svr(W)$	p-value	$svr(S)$	p-value	$svr(L)$	p-value	η	WCD_{LM}	WCD
Inflation	-4.169	1.000	-1.031	0.849	8.302	0.000	0.096	117.47	44.913
Unemployment	1.814	0.035	0.079	0.468	6.482	0.000	0.096	63.813	33.165
GVApc	5.766	0.000	0.485	0.314	2.764	0.003	0.235	50.433	28.169
AC	2.921	0.002	1.466	0.071	3.023	0.001	0.390	7.851	4.662
Local	3.547	0.000	0.217	0.414	4.984	0.000	0.154	22.665	14.753

Table 2: Pesaran (2007) panel data unit root test statistics

		Intercept		Linear trend	
		CIPS	CIPS*	CIPS	CIPS*
Inflation	$p = 0$	-3.341^b	-3.329^b	-3.600^b	-3.597^b
	$p = 1$	-3.117^b	-2.946^b	-3.687^b	-3.390^b
	$p = 2$	-2.320^b	-2.320^b	-2.781^b	-2.700^b
	$p = 3$	-1.837	-1.837	-2.649	-2.597
	$p = 4$	-1.243	-1.243	-2.234	-2.196
	$p = 5$	-2.173	-1.717	-2.194	-2.113
Unemployment	$p = 0$	-2.349^b	-2.349^b	-2.473	-2.473
	$p = 1$	-2.423^b	-2.423^b	-2.383	-2.383
	$p = 2$	-2.280^b	-2.280^b	-2.100	-2.100
	$p = 3$	-1.926	-1.926	-1.940	-1.940
	$p = 4$	-1.997	-1.997	-2.025	-2.025
	$p = 5$	-2.150	-2.150	-2.251	-2.251
GVApc	$p = 0$			-2.793^b	-2.768^b
	$p = 1$			-2.530	-2.530
	$p = 2$			-1.865	-1.865
	$p = 3$			-1.726	-1.726
	$p = 4$			-1.321	-1.337
	$p = 5$			-1.319	-1.319
AC	$p = 0$	-2.012	-2.012	-3.148^b	-3.148^b
	$p = 1$	-1.458	-1.458	-2.659	-2.659
	$p = 2$	-0.985	-0.985	-1.891	-1.847
	$p = 3$	-1.043	-0.885	-1.594	-1.601
Local	$p = 0$	-3.021^b	-3.021^b	-3.120^b	-3.120^b
	$p = 1$	-2.271^b	-2.271^b	-2.277	-2.277
	$p = 2$	-1.929	-1.929	-1.893	-1.893
	$p = 3$	-1.860	-1.860	-1.824	-1.824
	$p = 4$	-1.735	-1.735	-1.763	-1.763
	$p = 5$	-1.959	-1.959	-2.078	-2.078

Superscript b denotes rejection of the null hypothesis of unit root at the 5% level of significance

Table 3: Moon and Perron (2004) panel data unit root tests

Intercept				
	t_a	p-value	t_b	p-value
Inflation	-5.649	0.000	-8.299	0.000
Unemployment	0.053	0.521	0.053	0.521
AC	-0.731	0.232	-1.314	0.094
Local	-8.693	0.000	-4.515	0.000

Linear trend				
	t_a	p-value	t_b	p-value
Inflation	-7.786	0.000	-8.840	0.000
Unemployment	0.518	0.698	0.471	0.681
GVApc	-5.077	0.000	-5.381	0.000
AC	-10.077	0.000	-12.153	0.000
Local	-8.792	0.000	-6.865	0.000

Table 4: Bai and Ng (2004) panel data unit root tests

Intercept						
	$ADF_{\hat{e}}^{c}$	p-value	$MQ_f^c(m)$	$\hat{r}_1$	$MQ_c^c(m)$	$\hat{r}_1$
Inflation	0,200	0,579	-21,879	6	-21,558	6
Unemployment	-1,109	0,134	-32,173	5	-32,892	5
AC	-5,282	0,000	-15,484	6	-15,992	6
Local	-6,056	0,000	-35,334	6	-35,867	6

Linear trend						
	$ADF_{\hat{e}}^{\tau}$	p-value	$MQ_f^{\tau}(m)$	$\hat{r}_1$	$MQ_c^{\tau}(m)$	$\hat{r}_1$
Inflation	-0,813	0,208	-21,078	6	-21,864	6
Unemployment	-1,426	0,077	-34,232	5	-31,587	5
GVApc	0,535	0,704	-33,204	6	-31,975	6
AC	-0,454	0,325	-14,591	6	-15,640	6
Local	-1,865	0,031	-35,680	6	-35,604	6

Table 5: Bai and Carrion-i-Silvestre (2013) panel cointegration test statistics. Heterogeneous parameters

Model specification: $\pi_{i,t} = \gamma_i + \beta_{1,i}AC_{i,t} + \beta_{2,i}Local_{i,t} + \beta_{3,i}GVApc_{i,t} + F_t'\lambda_i + e_{i,t}$						
Num. of common factors (r)	Z	p-value	P_m	p-value	P	p-value
1	12.523	1.000	-1.000	0.841	25.750	0.844
2	7.731	1.000	13.425	0.000	144.707	0.000
3	0.130	0.552	1.607	0.054	47.251	0.065
4	2.297	0.989	-0.044	0.518	33.638	0.485
5	5.094	1.000	6.971	0.000	91.485	0.000
6	1.519	0.936	1.195	0.116	43.857	0.120
Model specification: $\pi_{i,t} = \gamma_i + \beta_{1,i}AC_{i,t} + \beta_{2,i}Local_{i,t} + \beta_{3,i}GVApc_{i,t} + \beta_{4,i}U_{i,t} + F_t'\lambda_i + e_{i,t}$						
Num. of common factors (r)	Z	p-value	P_m	p-value	P	p-value
1	2.314	0.990	1.158	0.123	43.552	0.126
2	1.281	0.900	4.227	0.000	68.857	0.000
3	0.410	0.659	-0.883	0.811	26.717	0.809
4	0.600	0.726	6.677	0.000	89.062	0.000
5	-2.052	0.020	2.748	0.003	56.664	0.009
6	0.661	0.746	7.700	0.000	97.498	0.000

Table 6: Bai and Carrion-i-Silvestre (2013) panel cointegration test statistics. Homogeneous parameters

Model specification: $\pi_{i,t} = \gamma_i + \beta_1 AC_{i,t} + \beta_2 Local_{i,t} + \beta_3 GVApc_{i,t} + F_t'\lambda_i + e_{i,t}$						
Num. of common factors (r)	Z	p-value	P_m	p-value	P	p-value
1	12.064	1.000	0.618	0.268	39.097	0.252
2	19.020	1.000	68.042	0.000	595.089	0.000
3	6.213	1.000	1.345	0.089	45.092	0.097
4	5.215	1.000	2.288	0.011	52.871	0.021
5	10.606	1.000	-0.921	0.821	26.406	0.821
6	5.118	1.000	0.846	0.199	40.975	0.191
Model specification: $\pi_{i,t} = \gamma_i + \beta_1 AC_{i,t} + \beta_2 Local_{i,t} + \beta_3 GVApc_{i,t} + \beta_4 U_{i,t} + F_t'\lambda_i + e_{i,t}$						
Num. of common factors (r)	Z	p-value	P_m	p-value	P	p-value
1	13.848	1.000	0.133	0.447	35.095	0.416
2	21.965	1.000	17.605	0.000	179.174	0.000
3	6.074	1.000	61.086	0.000	537.729	0.000
4	6.236	1.000	0.489	0.313	38.029	0.291
5	11.746	1.000	-0.874	0.809	26.791	0.806
6	5.084	1.000	0.854	0.197	41.040	0.189

Table 7: Estimation of the panel data cointegrating vector using DOLS estimation procedure in Bai and Carrion-i-Silvstre (2013)

	CCE estimates		DOLS estimates				
	(*a*)	(*b*)	(1)	(2)	(3)	(4)	(5)
GVApc	0.452	-0.126	1.885	1.525	1.015	1.424	1.417
	(0.178)	(-0.047)	-17.139	-13.424	-8.773	-12.009	-13.935
AC	0.147	-0.056	-0.126	-0.076			
	(0.206)	(-0.067)	(-0.916)	(-0.515)			
Local	-0.157	-0.085	-1.289	-1.199			
	(-0.159)	(-0.081)	(-7.291)	(-6.751)			
Unemployment		-0.027		-0.037	-0.037	-0.032	-0.034
		(-0.749)		(-10.366)	(-10.443)	(-8.528)	(-9.955)
AC*D1					1.061	-0.209	
					-5.363	(-0.67)	
AC*D2					-1.138	-1.233	-1.161
					(-6.355)	(-6.829)	(-6.962)
AC*D3					-0.808	-2.7	-2.597
					(-2.74)	(-6.606)	(-8.351)
AC*D4						1.941	1.797
						-6.398	-9.434
Local*D1					-0.718	-0.867	-0.704
					(-2.787)	(-2.239)	(-2.757)
Local*D2					-1.827	-1.678	-1.664
					(-5.299)	(-4.918)	(-4.865)
Local*D3					-1.594	-1.617	-1.416
					(-4.679)	(-3.87)	(-4.152)
Local*D4						0.254	
						(0.564)	

Individual significance test statistics between parentheses