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EUROPEAN COMMISSION

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Abstract

We provide key stylised facts on fiscal policy developments in Spain over the past three decades using quarterly data (1986Q1-2012Q2). First, we compute stylised facts on the cyclical properties of fiscal policies over that period. Next, we report updated evidence on the macroeconomic effects of non-systematic fiscal policies, including updated estimates of their macroeconomic impact (fiscal multipliers) for alternative datasets. To perform the analysis in the paper we built up a comprehensive database of seasonally adjusted quarterly fiscal variables for the period of interest.

Keywords: fiscal policies, stylised facts, fiscal multipliers, mixed-frequencies, time-series models.

JEL classification: E62; E65; H6; C3; C82.

Resumen

En este trabajo proporcionamos evidencia sobre las propiedades cíclicas y el impacto de la política fiscal en España, a lo largo de las últimas tres décadas, utilizando datos trimestrales (1 TR 1986-2 TR 2012). En primer lugar, analizamos la sincronía cíclica entre los agregados fiscales más relevantes y el crecimiento económico durante este período. En segundo lugar, nos centramos en los efectos macroeconómicos de la política fiscal no sistemática, proporcionando estimaciones actualizadas de un conjunto de multiplicadores fiscales. Para realizar el análisis anterior, en el artículo construimos una base de datos en la frecuencia trimestral que cubre, en particular, los principales agregados de la cuenta de las Administraciones Públicas para el período de interés.

Palabras clave: política fiscal, hechos estilizados, multiplicadores de la política fiscal, modelos de series temporales de frecuencias múltiples.

Códigos JEL: E62; E65; H6; C3; C82.

1 Introduction

Fiscal policies are at the forefront of the economic policy debate in Europe nowadays. Thus it is not surprising to see that an enormous amount of papers have been recently devoted to the analysis of the macroeconomic impact of fiscal policies, the sustainability of public debt or the properties of fiscal consolidations, mostly from a cross-country point of view. The focus on country-specific cases from a broad perspective, though, is more scarce. That is why we focus our paper in one specific case, Spain, which has been until recently in the midst of the euro area sovereign debt crisis. In particular, we concentrate on two specific applications that are relevant from the policy point of view and that have received only partial coverage in existing studies, and/or that deserve an update compared to previous works. On the one hand, we compute stylized facts on the cyclical properties of fiscal policies over the past three decades using quarterly data and focusing on the General Government sector as a whole. On the other hand, we report updated values of the impact of non-systematic fiscal policies on the economy (so-called fiscal multipliers), including the impact of the crisis period, thus covering the historical period that runs from 1986Q1 to 2012Q4.

In the case of Spain, given that quarterly government finance statistics for the General Government sector are only available for the period starting in 2000Q1, in nominal, non-seasonally adjusted terms[1] we decided to adopt the modeling approach of Paredes et al. (2009; 2014) and construct a quarterly fiscal database for Spanish government accounts for the period 1986Q1-2012Q4, solely based on intra-annual fiscal information.[2] As recently claimed by Dilnot (2012) public policy analysis should not be undertaken without thinking carefully and then finding out the numbers. This in itself is the first contribution of our paper, beyond the empirical applications discussed.

The part of the study devoted to the cyclical properties of fiscal policies in Spain is warranted, as only a few studies have dealt, either directly or indirectly, with the hurdle of computing stylized facts on fiscal policies (see Dolado et al., 1993; Marín, 1997; Ortega, 1998; Esteve et al., 2001; André and Pérez, 2005). The topic is clearly relevant from the current, crisis-related perspective, against the background of the renewed support for activist, counter-cyclical fiscal policies that reappeared right after the post-Lehman slump (e.g. Spilimbergo et. al., 2008, Bouthevillain et al.,

[1]See European Commission (2002a, 2002b, 2006).

[2]On the basis of multivariate, state-space mixed-frequencies models, along the lines of Harvey and Chung (2000). The models are estimated with annual and quarterly national accounts fiscal data and government monthly cash and national accounts data.

2009), and that is regaining footage since 2012.[3] In fact, the role of fiscal policies in stabilizing the economy gained policy relevance since the creation of the European Economic and Monetary Union (EMU), given that it was increasingly argued that fiscal policies should take a greater role in demand and output stabilization over the business cycle in euro area countries than before EMU due to the fact that individual countries lost control of their monetary policy tools.[4] In our paper, we analyze the cyclical properties of the main components of the revenue and the expenditure sides of the budget. We look at the unconditional correlation between filtered/detrended series via various ways of filtering. As in Lamo et al. (2013) we distinguish between the fluctuations around the trend that are driven by unpredictable or irregular components of the series (irregular shocks, ad-hoc policy measures, etc.) from those that look at the cyclical components (mixture of systematic autocorrelation properties of the filtered series and irregular factors). We find this particularly relevant as in our case the irregular components are quite likely to reflect policy induced fluctuations, i.e the dynamics of the series due to policy measures.

An alternative way of looking at the correlation of discretionary fiscal policy shocks and macroeconomic variables is to consider the traditional dynamic SVAR approach to the computation of so-called fiscal multipliers. In this literature some identification assumptions are used to compute the macroeconomic effects of fiscal shocks, moving beyond the computation of correlations to the estimation of causal impacts (conditional correlations) of non-systematic fiscal policies. At the current policy juncture, the success of ongoing fiscal consolidations depends crucially on the value of the multiplier, as pointed out e.g. by Boussard et al. (2012). For the case of Spain, previous papers that cover this issue are de Castro (2006), de Castro and Hernández de Cos (2008), de Castro and Fernández-Caballero (2013), European Commission (2012), and Hernández de Cos and

[3]See, e.g. Wren-Lewis (2013), or the discussion around the sizeable Japanese 2013 fiscal stimulus package.

[4]The theoretical literature suggests little consensus as to whether fiscal policies should have, are likely to have, or in fact do have a stabilizing effect on demand. Keynesian economics suggests that governments should and would stabilize demand by behaving counter-cyclically while the normative predictions from a neoclassical perspective depend on the relationship between private and public consumption. Political economy models generally predict pro-cyclical discretionary policies, as interest groups see public spending (Lane and Tornell, 1996) or taxation (Alesina et al., 2008) as a common good to deviate to their benefit, and the pressure to do so, is stronger in economic booms. Some grounds for a-cyclicality can be found in political economy models in which the different status that civil servants enjoy might make public wages and consumption less reactive to the business cycle, or even generate a separate agenda of public employees (rent-seeking behaviour), as in Borjas (1984). Other related papers on the issue of the cyclicality of fiscal policies are Lane (2003), Akitoby et al (2004), Strawczynski and Zeira (2009), Aghion et al. (2009), Afonso et al. (2010), or Coricelli and Fiorito (2013).

Moral-Benito (2013). We update and complement the estimates in those papers, in particular by including the most recent crisis period in a comparable framework, and by analyzing robustness with respect to alternative datasets. Fine tuning country-specific estimates is a crucial issue to draw policy lessons, given that the most recent literature has stressed the significant heterogeneity of estimates derived from general theoretical and empirical models (see European Commission, 2012).

The rest of the paper is organized as follows. In Section 2 we describe the data used in the analysis and provide some descriptive evidence. In that Section we also include a description of the modeling approach used to back-cast part of the sample and the historical input data used for that purpose, leaving specific details to appendices A and B. In addition we discuss from a descriptive point of view the main features of fiscal policies in the sample period covered by our study. In Section 3 we turn to provide stylized facts on cyclical fiscal policies, while in Section 4 we report updated estimates of the macroeconomic impact of non-systematic fiscal policies (fiscal multipliers). Finally, in Section 5 we provide the main conclusions of the paper. The study also includes a number of technical appendices, about the detrending methods used (C) and about the SVAR approach (D).

2 The data

Quarterly General Government figures on an ESA95 basis are available only for the period 2000 onwards, in non-seasonally adjusted terms, and are released by the accounting office IGAE. Unfortunately, this information is not available for previous years. There is one exception to this general pattern about quarterly fiscal data: aggregate public consumption. Nominal and real government consumption expenditure (seasonally and non-seasonally adjusted) are available on a quarterly basis since 1995 in ESA95 terms. These data can be obtained from the Quarterly National Accounts published by the INE. Moreover, the INE also offers the quarterly data for the same variables between 1985 and 1998 on an ESA79 basis.

Two existing databases have been built over the past decade to overcome this lack of official statistics[5]. A first quarterly dataset is the one compiled by Estrada et al. (2004), for the period starting in 1981Q1. This database is the one used to estimate and simulate Banco de España's quarterly macroeconometric model (MTBE henceforth) and thus the interpolation procedure applied

[5] An early contribution along these lines is Corrales and Taguas (1991).

and the indicators used were selected with this specific purpose in mind. Except for public consumption, standard interpolation techniques – Denton method in second relative differences with relevant indicators – were applied to pre-seasonally-adjusted figures.[6] This is a valid approach given the stated uses of the MTBE model and the generated quarterly fiscal dataset is fully consistent with model definitions. A second information source is the REMS database (Boscá et al., 2007), companion to the REMS model (see Boscá et al., 2011) – a DSGE model currently used within the Ministry of Economy and Finance to carry out policy simulations – that includes a quite detailed fiscal block with quarterly variables. The REMS database includes a large set of macroeconomic, financial and monetary variables, and also a group of public sector variables, for the period starting in 1980Q1. The quarterly non-financial fiscal variables in that block are obtained from annual data by quadratic interpolation, ensuring consistency with annual data, though.

In our paper we decide to move one step beyond existing alternatives for a number of reasons. First, we have constructed a new dataset following a proven and transparent methodology, the one used by Paredes et al. (2009; 2014) to build up the euro area fiscal database that is disseminated jointly with ECB's Area Wide Model general macroeconomic database.[7] In this respect, given that we only use publicly available information, our database is to be made freely available upon request.

Beyond this quite relevant transparency consideration, a second reason is related to the nature of the inputs used in the interpolation exercise. Our database makes use of only intra-annual fiscal information. This is a relevant point for further research devoted to the integration of interpolated intra-annual fiscal variables in more general macroeconomic studies, because it allows to capture genuine intra-annual "fiscal" dynamics in the data. This is very important because although government revenues and expenditures (e.g. unemployment benefits) may be endogenous to GDP or any other tax base proxy, the relationship between these variables is at most indirect and extremely difficult to estimate. The decoupling of tax collection from the evolution of standard macroeconomic tax bases (revenue windfalls/shortfalls) is by now a proved stylized fact (see Morris et al., 2009). Thus, we use only fiscal data for interpolation purposes, which overcomes the problem of modeling an indirect relationship which is time-varying.

A third feature of our approach is that, as in Paredes et al. (2009; 2014), we tried to follow to the extent possible some of the principles outlined in the manual on quarterly non-financial

[6] The interpolation relies heavily on Central Government variables. The available quarterly nominal, non-seasonally adjusted General Government series that start in 2000Q1 are not used in the interpolation procedure.

[7] See "The AWM database", September 2010, available at the official AWM site at the Euro Area Business Cycle Network (http://www.eabcn.org/data/awm/index.htm).

accounts for general government: use of direct information from basic sources (public accounts' data), computation of "best estimates", and consistency of quarterly and annual data. As regards the coherence of quarterly data with annual rules, the discussion in European Commission (2002a, 2002b, 2006) shows that there is some room for econometric estimation of intra annual fiscal variables.[8]

Additional methodological issues related to the compilation of our database ("QESFIPDB" henceforth) are detailed in Appendices A and B.[9]

2.1 A first look at the data

Figures 1, 2, 3, and 4 display the main fiscal aggregates of the Spanish General Government sector over the period 1986Q1-2012Q4.[10] Figure 1 presents the public deficit and debt, and the decomposition of the budget balance in revenues and expenditures. The latter aggregates, in turn, are further decomposed in figures 2 and 3. Finally, within total expenditure in Figure 4 we present the components of government consumption.

Overall, the Spanish government sector presented a deficit in 87% of the quarters covered by the dataset (1986Q1-2012Q4). This is quite clear visually from the two lower panels of Figure 1: public expenditure presented a mean value of 41.4% of GDP, and exceed 40% of GDP in 56% of the quarters, while public revenue averaged 38%, and only exceed 40% of GDP in 8% of the quarters, that coincided with the pre-crisis, housing-boom-related windfall revenues. At the same time, public debt increased substantially during crisis periods, and declined significantly between the late nineties and 2007.

[8] This is the case for two main reasons. Firstly, ESA95 does not consider the quarterly aspects of taxes and social payments with sufficient precision to ensure clarity of interpretation in all situations; this is because, when discussing non-financial accounts, the ESA95 guiding documents occasionally take a perspective which assumes an annual reference period is in mind, thus remaining silent on which quarter within a particular annual reference period is involved. Secondly, it is also the case that many accounting or legal events are annual events by definition (e.g. a tax levied in a complete year); this fact does not present a problem for the statistician compiling annual data (there is no need to establish the amount and time of recording to a particular annual reference period), but do pose problems for the compiler of quarterly data, that needs to attribute revenue and expenditure not merely to a reference year but also to quarters within that year.

[9] The QESFIPDB is to be made freely available to interested researchers and policy analysts, in the vein of the Paredes et al (2009; 2014) euro area quarterly fiscal database, that is distributed by the Fiscal Policies Division of the European Central Bank upon request.

[10] Quarterly nominal GDP for the period 1986Q1-2012Q4 has been taken from the Banco de España database.

Figure 1: *Main Government finance variables. Percent of nominal GDP.*

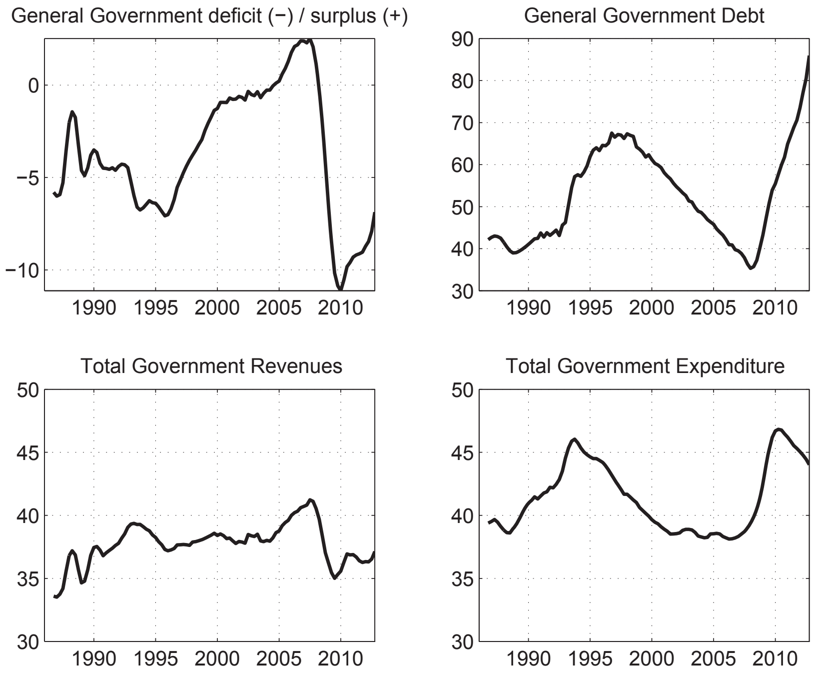
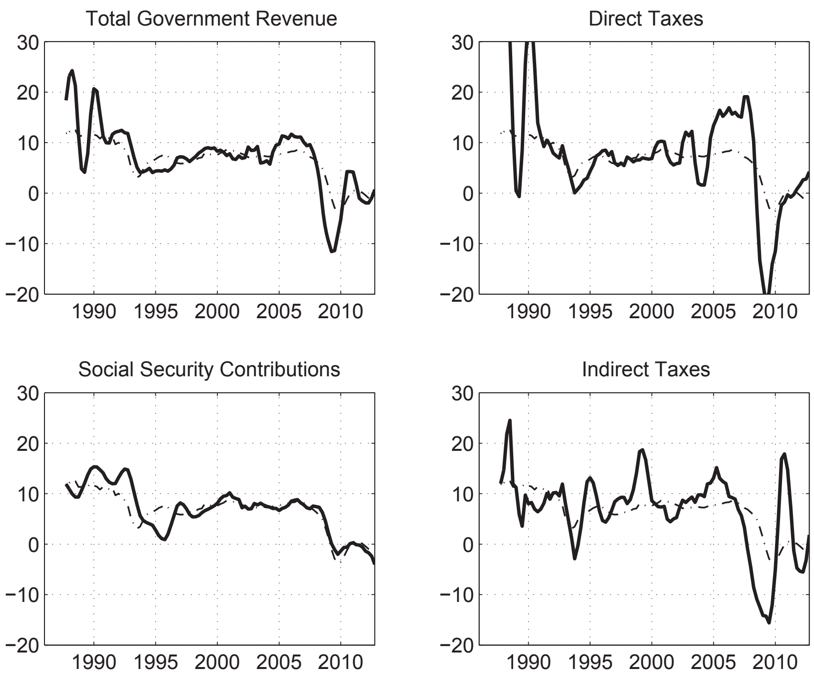


Figure 2: *General government revenues and nominal GDP (dashed line). Year-on-year growth rates of 4-quarter-moving-sums (seasonally-adjusted, nominal terms).*

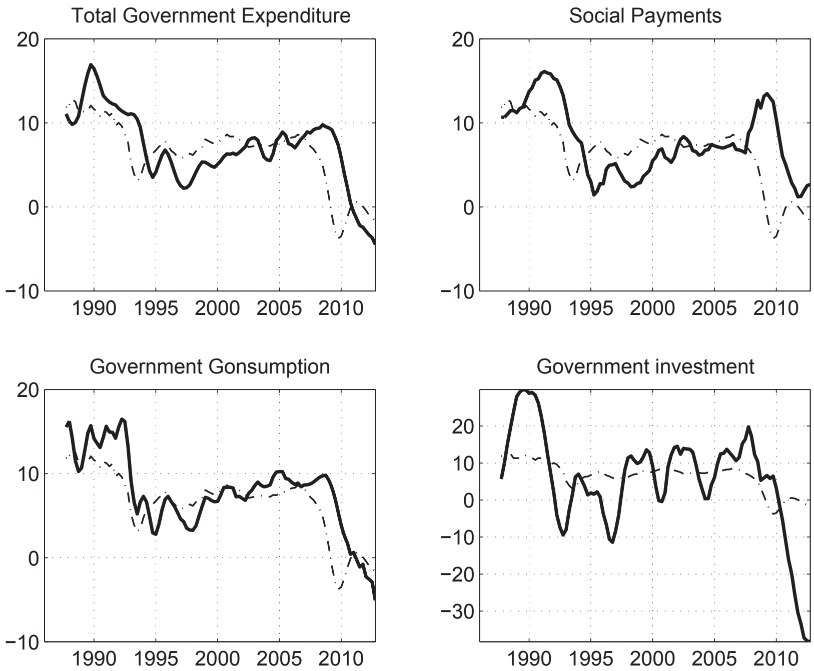


Turning to a chronological exposition, between 1986 to 1988, following Spain's accession to the European Community and the commencement of a new cyclical expansion, there was a change in direction in Spanish fiscal policy. This period was characterized by the reduction of the budget deficit from 5.8% in 1986 to 3.3% in 1988, essentially due to the growth of government revenue. In fact, public revenue as a percentage of GDP increased by 2.1 percentage points while public expenditure fell by some 0.5 percentage points. Moreover, there was a significant improvement in the primary balance, that enabled public debt to be whittled down to 39% of GDP at the end of 1988, down from the local maximun of 43% reached in 1987Q3. Despite the reduction in the expenditure-to-GDP ratio, public outlays registered very dynamic growth rates which prevailed for some years, until the early nineties (see Figure 3). Such expansion is linked to the phasing-in of the Welfare State in Spain.

This period of limited fiscal restraint came to an end in 1989, when the budget deficit started growing again to reach 7% at the height of the economic crisis that started in 1993. The primary balance followed a similar path to the deficit. After small surpluses between 1987 and 1989, it moved into deficit in 1990. Finally, there was only a slight increase in public debt, to 45.6% of GDP in 1993, primarily as a consequence of the strong growth in GDP between 1989 and 1991 (11% in nominal terms), and despite the increase in the cost of debt during this period. In the following years however, public debt rocketed to exceed 60% of GDP in 1994, as a consequence of the sizeable budget deficits and the fall in nominal GDP growth due to the economic crisis and the prohibition on monetary financing of the deficit as of 1994, under the Treaty on European Union. At the same time, the interest burden rose, reaching almost 5% of GDP in 1994.

The second half of the nineties, especially since 1996, is characterized by a protracted period of fiscal consolidation due to the commitment to meet the convergence criteria set out in the Treaty on European Union to regulate access to the Third Stage of EMU. Accordingly, public deficits displayed a declining trend that spread until 2007, when an unprecedented surplus of some 2% of GDP was recorded. Such steady and protracted reduction of general government deficits came hand in hand with a prolonged period of economic expansion. However, not all the years between 1996 and 2007 can be duly labeled as years of "fiscal consolidation". As Figure 1 shows, the reduction in the public deficit was mainly the result of a drop in spending in the second half of the nineties, which fell by more than 6 points of GDP. However, this trend of expenditure retrenchment was reverted in the following years. In fact, public expenditure in nominal terms registered elevated growth rates, only masked by even higher nominal GDP growth, partly due to high inflation. Moreover,

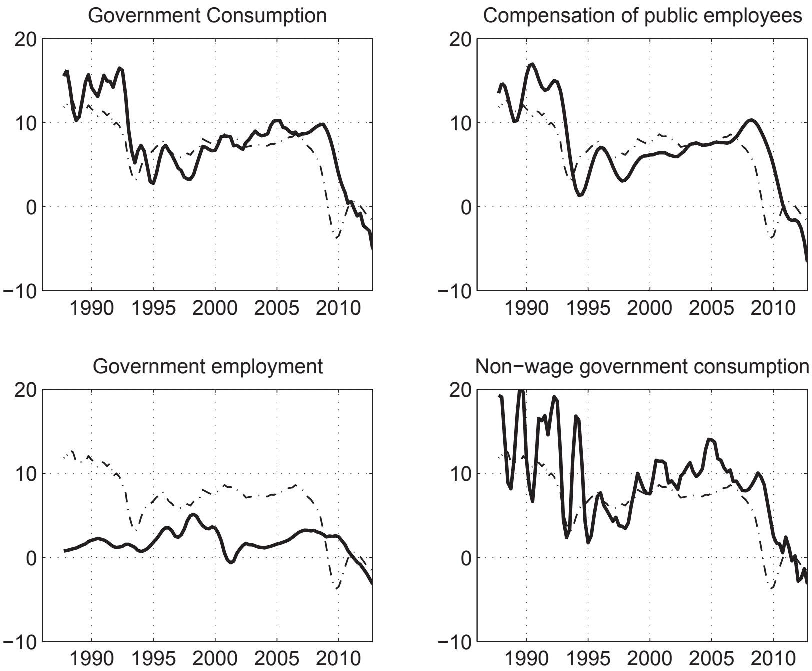
Figure 3: *General government expenditures and nominal GDP (dashed line). Year-on-year growth rates of 4-quarter-moving-sums (seasonally-adjusted, nominal terms).*



after 2004 nominal government expenditure displayed growth rates above those of nominal GDP (Figure 3). Still, the deficit reduction continued as a result of the buoyancy of tax revenues (Figure 2), which benefited largely from the tax-friendly growth composition to a large extent linked to the disproportionate development of the construction sector, especially related to the construction of dwellings.

The figures also show the strong drop in revenues since the onset of the financial crisis, in particular on indirect taxes and direct taxes, while social security contributions were more resilient. In fact, as it is apparent from Figure 2, the most recent crisis was the only period since 1986 in which total revenues (annualized) entered into negative territory in nominal terms. Total revenues contracted for seven consecutive quarters, namely from 2008Q1 till 2009Q3, presenting an average drop of 8.5% per quarter in year-on-year terms. The "double-dip" that the Spanish economy suffered between 2011Q4 and 2013Q4 also implied some negative registers of nominal government revenue in year-on-year terms (five consecutive quarters, from 2011Q2 to 2012Q2), even though the tax-increasing measures enacted over that period were conductive to containing the downward pressures arising from depressed tax bases.

Figure 4: *Government consumption components. Year-on-year growth rates of 4-quarters moving sums (seasonally-adjusted, nominal terms).*



Looking at the most recent decade, within total public expenditure, government consumption and investment were the most dynamic components in the pre-crisis period, and also the ones that were contracted more in the most recent fiscal retrenchment episode, as evident from Figure 3. Between 2008Q1 and 2010Q3 public consumption increased on average by 6% per quarter on a year-on-year basis, though on a decreasing path, in particular since mid-2010. As of 2010Q4, government consumption displayed negative rates of change till the end of 2012, with the exception of the first quarter of 2011 (a quarter preceding the local and regional elections of May 2011). Within public consumption (Figure 4) the main part of the adjustment was taken by the wage bill component, including public employment. In turn, government investment displayed a distinctive cycle-like pattern, typically described as being synchronized with electoral cycles. During the financial crisis, between 2008Q1 and 2009Q4 a number of fiscal stimulus packages made public investment grow by close to 10% in cumulative terms (6% on average per quarter in year-on-year terms), but since 2010Q1 and till 2012Q4 it dropped by almost 70%, what implied that more than 50% of the overall expenditure adjustment done by the Spanish public administrations over the period 2010-2012 was due to public investment reduction.[11] Finally, as regards social payments, the significant

[11]Between 2010Q1 and 2012Q4 total expenditure fall by 2.7% of GDP. Bearing in mind that Social payments increased by 1.3 points of GDP and interest payments by 1.2 points, the rest of components of spending fall by 5.2%

increase of 2008-2009 was due to massive unemployment spending, even though the rest (including, most notably, pensions' spending) showed a cumulated increase of some 10%. In 2010-2012, social payments decelerated considerably, mainly due to the cumulated fall in unemployment-related spending, partly due to some cost-cutting measures, and the deceleration of pensions and others (that presented a nominal increase of 5.6% between 2010Q1 and 2012Q4).

In this Section we have shown some stylized facts of fiscal policies over 1986Q1-2012Q4, that will be complemented, from more analytical angles, in the next two Sections.

3 Cyclical properties of fiscal variables

In tables 2 and 3,we report dynamic cross-correlation functions. We look at the unconditional correlations between detrended series at the standard business cycle frequencies. The underlying assumption to detrending filters is that aggregate seasonally-adjusted economic time series can be decomposed into a trend component, T_t, the so-called cyclical component, C_t, that fluctuates around the trend, and an unpredictable random component (or irregular component), ϵ_t, i.e. $y_t = T_t + C_t + \epsilon_t$. Most of the detrending filters take out the trend component from the original time series, so that both the cyclical and irregular components $C_t + \epsilon_t$ are taken as a measure of the cycle. To try to isolate the systematic autocorrelation properties of the filtered series C_t from the irregular fluctuations or nonsystematic behavior of the series, ϵ_t, we use univariate ARIMA filters to extract ("pre-whiten") the later from the detrended components $C_t + \epsilon_t$. [12]

Following standard practice we measure the co-movement between two series using the cross correlation function (CCF thereafter). Each row of the table displays the CCF between a given detrended fiscal variable at time t+k, and detrended GDP at time t.[13] In selecting this statistic, we follow the common practice in the related literature that shows results for Spain (Dolado et al., 1993; Marín, 1997; Esteve et al., 2001; André and Pérez, 2005; Lamo et al., 2013), and other works in the general literature of fiscal policies' stylized facts. Following the standard discussion in the literature, it is said that the two variables co-move in the same direction over the cycle if the maximum value in absolute terms of the estimated correlation coefficient of the detrended series

of GDP. Of the latter amount, government investment was responsible for 2.8 points, i.e. 53.4% of the total.

[12]See André and Pérez (2002; 2005) and the references quoted therein. See also den Hann (2000) for a similar procedure in a VAR framework.

[13]In the case of pre-whitened variables, the corresponding row of the table shows the CCF between the irregular components of the fiscal variable fiscal variable at time t+k, and the irregular component of GDP at time t.

(call it dominant correlation) is positive, that they co-move in opposite directions if it is negative, and that they do not co-move if it is close to zero. A cut-off point of 0.20 roughly corresponds in our sample to the value required to reject at the 5% level of significance the null hypothesis that the population correlation coefficient is zero. Finally, the fiscal variable is said to be lagging (leading) the real economic activity variable if the maximum correlation coefficient is reached for negative (positive) values of k. For the sake of robustness, and following Lamo et al. (2013), we show results for the mean of a set of standard filters[14] [15] as applied to seasonally-adjusted time series. In Appendix C we describe the methods used. For ease of exposition, in Table 1 we present the list of variables (and acronyms) used.

Turning to the results, Table 2, Panel 1, shows the strong and pro-cyclical behavior of total government revenue in Spain, a result that is consistent with the evidence based on annual data for Spain for the 1960-1990 period (see Esteve et al., 2001), and also with the existing results for the euro area aggregate (see Paredes et al., 2009; 2014), and G7 countries (Fiorito, 1997). Total revenue mimics the business cycle behavior in upturns and downturns, reflecting the operation of automatic stabilizers, in a broadly contemporaneous manner – the dominant correlation is contemporaneous when the crisis period is excluded, and when the systematic part of the correlation is cleaned up (pre-whitened series). The dominant correlation is high, ranging between 0.5 (pre-crisis) and 0.6 (full sample). When we filter out the dynamics of the series due to systematic autocorrelation to end up with the irregular component, the correlation is somewhat weaker but high (0.3 and 0.4, respectively). This indicates that the unpredictable component of the series is responsible for an important part of the pro-cyclicality of the real public revenue series. As regards relative volatility, government revenues are much more volatile than GDP, close to three times, on average, a figure higher than the one for the euro area. This may reflect the fact that a number of taxes, most notably corporate taxes, property taxes and other indirect taxes, tend to follow boom-bust dynamics and

[14] We have selected a number of filters that are standard in the related, macroeconomic literature. The filters are: (i) first difference filter; (ii) linear trend; (iii) Hodrick-Prescott filter for two alternative values of the band-pass parameter: the standard 1600, and a higher value of 8000 the one suggested by Marcet and Ravn (2004) for countries with more volatile cycles (like Spain); (iv) The Band-Pass filter of Christiano and Fitzgerald (2003), with two different pairs of band-pass parameters $[p_l, p_u]$, capturing fluctuations between 1.5 and 8 years ($[p_l = 6, p_u = 32]$) and between 1.5 and 12 years ($[p_l = 6, p_u = 48]$).

[15] An alternative to the combination of filters would have been to focus on just one specific filter, for example, the Band-Pass filter. We prefer to follow the more agnostic approach of Lamo et al. (2013) and show average results for a number of standard filters. In any case, this is done only for the sake of exposition, and all individual filters' results are available from the authors upon request.

Table 1: *List of General Government sector variables covered, and acronyms for them.*

Name of the variable	Acronym	Name of the variable	Acronym
Net lending (+)/net borrowing (-) (public surplus/deficit)	DEF	Public debt	MAL
Total government revenue	TOR	Total expenditure	TOE
Direct taxes	DTX	Social payments	THN
Social security contributions	SCT	of which Unemployment benefits	UNB
Total indirect taxes	TIN	Interest payments	INT
Value Added taxes	VAT	Subsidies	SIN
Other indirect taxes	OTIN	Government investment	GIN
	= TIN - VAT	Nominal government consumption	GCN
		Government wage consumption expenditure	COE
		Non-wage final consumption expenditure	OGCN
		Other expenditure	OTOE = residual from TOE
		Real government consumption	GCR
		Government employment	LGN

do react to the cycle more than proportionally (Morris and Schuknecht, 2007). In addition, the progressive structure of the income tax also causes excess volatility of income taxes relative to the business cycle.

This is clear from Table 2, Panel 2, were we show how the standard deviation of the cyclical component of direct taxes (mixture of household and corporate taxes) are around 5 times higher than the one of real GDP, while that of indirect taxes is around 4 times as high, while the relative volatility of Social Security contributions with respect to GDP ranges between 1.5 (full sample) and 2 (pre-crisis). In this respect, it is not surprising that indirect taxes are less volatile than direct taxes, given that, for example, tax bases such as households' income and corporate profits are more volatile than private sector consumption. Also, social contributions rely on personal income, but at proportional rates and applied to censored tax bases. This creates a lower relative volatility. In the case of Spain, the fact that unemployment benefits pay social contributions may create some extra volatility relative to real GDP in downturns. Compared to the results of Esteve et al. (2001) and Marín (1997), with data up the 1990s, our results by components show robust evidence of pro-cyclicality of revenue components. In fact, when looking at the whole sample versus the pre-crisis sample, co-movements between cyclical components are stronger in all the cases considered. Interestingly, as regards, the leading/lagging structure of public revenue components when the whole sample is considered, the overall contemporaneous observation of Panel 1 as regards total revenues is the combination of a leading behavior of direct and indirect taxes, while social contributions display a lagging behavior, more consistent with the traditional role of government revenues as automatic stabilizers.

Most existing studies look at the cyclical properties of government spending (see Frankel, Vegh and Vuletin, 2013, and the references quoted therein). Indeed, an important reason for the usual finding of pro-cyclical spending is precisely that government receipts get increased in booms, typically beyond expectations, and thus governments use the surplus to increase spending proportionately as a consequence of political pressure or just following certain social-welfare-improving objectives. We show the cyclical properties of total government expenditure in Table 2, Panel 1, and those of its components in Table 3. As expected, in Table 3 total expenditure appears pro-cyclical as well, but lagged; this behavior can be rationalized on the basis of the political economy arguments mentioned before. The lag detected with quarterly data implies that total expenditure follows GDP with a -minimum- delay of one year. Budgetary patterns on the spending side tend to be quite persistent, in particular as regards sizeable items like public wages or public employment. For example, only in the period following an economic downturn are fiscal consolidation measures implemented, while in expansions, fresh government revenues tend to expand the public sector wage bill with some delay. When the pre-whitened cyclical co-movements are considered, the correlation between the shocks is barely significant.

The pro-cyclical pattern of total expenditures is due to the government consumption component (GCR), in line with available evidence for the euro area obtained with annual data (see Lamo et al., 2013). The pro-cyclicality of public consumption in the case of Spain is a result already obtained in previous works for samples covering up to the beginning of the 2000s (Dolado et al., 1993; André and Pérez, 2005; Marín, 1997; Ortega, 1998; Esteve et al., 2001). Co-movements among unanticipated components are also pro-cyclical and explain a significant portion of the co-movement among cyclical components. Within government consumption (Table 4) the pro-cyclical, lagged co-movement is explained by the wage bill, while the correlation of the cycle of real GDP and that of non-wage government consumption is weaker, though still positive. For the whole sample, there is a positive cyclical co-movement of public employment and real GDP, but this is weaker that the one of the wage bill, suggesting that public wages (per employee) would be the part of the wage bill displaying the strongest correlation. Interestingly, when the inertia of the series is removed (pre-whitening), the cyclical correlation of public wage bill/public employment shocks/discretionary policies and real GDP shocks tend to be non-significant. At the same time, the correlation of shocks to non-wage government consumption with real GDP shocks is negative (counter-cyclicality). Again, these results as regards pre-whitened series might be an indication that the components of government consumption are usually used to adjust budgetary outcomes, with no clear pattern of individual components, while at the end the aggregate ends up being pro-cyclical. Government investment (again in Table 3) also displays a marked pro-cyclical co-movement. Nevertheless, when the inertia of the cycles is taken out, there remains no clear cyclical pattern. As in the case of government consumption, this might be an indication that shocks to government investment are used to meet budgetary outcomes, as also highlighted by Esteve et al. (2001).

By contrast, social payments reflects a counter-cyclical pattern, mainly due to the properties of unemployment benefits; unemployment-related spending increase in downturns and decrease in upturns, reflecting a role of automatic stabilization. The latter evidence is consistent with an interpretation whereby employment losses at the beginning of a cyclical downturn tend to be associated with new unemployed receiving full-entitlement benefits (given that downturns do occur after a good times period), coupled with the fact that the average duration of the entitlement tends to be lower than the number of quarters the economy is below trend. In Table 4 it is possible to see that besides the counter-cyclical pattern, unemployment spending leads the cycle, i.e. employment destruction may start somewhat ahead of a real GDP downturn. The dominant correlation for the whole sample is higher than the one obtained with the pre-crisis sample, given the sizeable increase of unemployment-related spending during the most recent crisis. As regard other social payments in cash, that include contributory and non-contributory pensions as well as other social transfers like temporary disability, the estimated cyclical correlation is weaker than the one estimated for unemployment benefits. Given the different nature of the components of that aggregate, results on co-movement with the business cycle are not clear cut. In fact, while in the pre-crisis sample a positive dominant correlation is found, consistent with the hypothesis that social spending is increased in good times and reduced in bad times, the inclusion of the crisis years (2008Q1-2012Q4) weakens the dominant correlation and turns it to a negative value (counter-cyclicality). This is consistent with the observation that during the whole crisis social payments gained weight over GDP, as they were conceived as a tool to guarantee social cohesion in the midst of the crisis and the fiscal consolidation process that came hand-in-hand with it. In any case, when the inertia of the series is removed, the correlation among pre-whitened series shows no particular pro- or counter-cyclical behavior, most likely reflecting the fact that decisions on pensions are relatively erratic: in good times they are increased above determinants like inflation or wages due to equity and/or electoral considerations, while in bad times their growth tend to be moderated on the back of fiscal consolidation needs.

As regards the other components of spending, interest payments exhibit a counter-cyclical behavior. In downturns interest payments increase on the back of increased public debt and, typically in the case of Spain, rising interest rates on newly issued debt. Correlations among shocks explain a significant portion of the cyclical co-movement. Subsidies seems to be also a counter-cyclical policy item, even thought the correlations are relatively low. Finally, other public expenditures, and aggregate of very different spending items given that gathers the rest of current and capital expenditures not discussed above, displays a positive co-movement, as total expenditure, even though it is weak.

As regards the relative volatility of real public expenditure over real GDP, as in the case of public revenue it is higher than one, but of an order of magnitude considerably lower, being around 1.5 as compared to 3-4 in the case of total revenues. In both cases the magnitude of the relative volatility is considerably higher that in the euro area aggregate case (see Paredes et al., 2014). This is an issue that may deserve further attention, given the findings in the literature on the detrimental effects of fiscal policies volatility on economic growth (Afonso and Furceri, 2010).

By component, social payments other that unemployment benefits, and government consumption, are the most stable components, with relative cyclical variabilities only slightly above one, while unemployment spending and government investment display volatilities six or more times the volatility of real GDP. In addition, subsidies (transfers to firms), being presumably decided on an occasional basis, are the most volatile component within the public expenditure categories considered. The relative stability of government consumption is consistent with the fact that the provision of public goods should be broadly independent of the business cycle. Yet, as signalled by Fiorito (1997), government consumption does not only include such public goods as defense, justice, etc., but also accommodates merit goods such as health or education that are partially supplied by private units and that also involve purchases from private firms. Within government consumption, the standard deviation of the wage component is half that of the non-wage component, and within the wage component public employment is the only government spending item with a relative volatility below one. The latter reflects the broad stability of most public employees in Spain, that enjoy a per-life civil-servant status.

Table 2: Stylized facts on fiscal policies I. Correlations between cyclical components of the fiscal variable (in real terms) and real GDP.[‡]

Panel A. Main aggregates

		$\frac{\sigma_y}{\sigma_x}$	-4	-3	-2	-1	0	1	2	3	4	
Total government revenues (TOR)												
Detrended	86Q1-12Q4	2.7	0.51	0.53	**0.59**	0.55	0.58	0.48	0.29	0.17	0.06	Pro-cyclical Lead
Detrended	86Q1-07Q4	2.6	0.24	0.20	0.29	0.29	**0.47**	0.46	0.31	0.26	0.19	Pro-cyclical Cont.
Pre-whitened	86Q1-12Q4	3.8	0.19	0.06	0.08	0.05	**0.41**	0.23	0.03	0.00	-0.06	Pro-cyclical Cont.
Pre-whitened	86Q1-07Q4	3.6	0.15	-0.10	-0.13	-0.17	**0.26**	0.20	0.13	0.17	0.18	Pro-cyclical Cont.
Total government expenditure (TOE)												
Detrended	86Q1-12Q4	1.5	-0.27	-0.31	-0.25	-0.23	-0.11	-0.05	0.04	0.19	**0.36**	Pro-cyclical Lag
Detrended	86Q1-07Q4	1.3	0.00	-0.08	-0.02	-0.05	0.07	0.11	0.17	0.31	**0.46**	Pro-cyclical Lag
Pre-whitened	86Q1-12Q4	1.4	-0.03	**-0.20**	**-0.14**	**-0.16**	**0.05**	0.01	**-0.15**	-0.02	0.04	A-cyclical
Pre-whitened	86Q1-07Q4	1.4	**0.14**	-0.05	-0.06	**-0.19**	-0.02	-0.03	-0.08	0.09	0.11	A-cyclical
Net lending (+) / net borrowing (-) (DEF)												
Detrended	86Q1-12Q4	1.4	0.64	0.69	**0.70**	0.63	0.55	0.41	0.20	0.03	-0.15	Pro-cyclical Lead
Detrended	86Q1-07Q4	0.9	0.35	0.36	0.39	0.37	**0.44**	0.40	0.22	0.10	-0.04	Pro-cyc. cont./lead
Pre-whitened	86Q1-12Q4	1.3	0.17	0.20	0.20	0.16	**0.25**	0.16	-0.10	-0.17	-0.18	Pro-cyc. cont.
Pre-whitened	86Q1-07Q4	1.3	0.01	-0.10	-0.03	0.03	0.25	**0.33**	0.12	0.07	0.07	Pro-cyc. cont./lag.

Panel B. Government revenue components.

		$\frac{\sigma_y}{\sigma_x}$	-4	-3	-2	-1	0	1	2	3	4	
Direct taxes (DTX)												
Detrended	86Q1-12Q4	5.0	0.48	0.51	**0.62**	0.58	0.61	0.54	0.39	0.27	0.15	Pro-cyclical lead
Detrended	86Q1-07Q4	5.3	0.37	0.31	0.41	0.37	**0.46**	0.42	0.30	0.20	0.10	Pro-cyclical cont.
Pre-whitened	86Q1-12Q4	6.9	0.14	-0.01	0.19	0.12	**0.33**	0.18	0.07	0.06	0.05	Pro-cyclical cont.
Pre-whitened	86Q1-07Q4	6.7	0.13	-0.18	0.01	-0.04	**0.21**	0.15	0.10	0.11	0.16	Pro-cyclical cont.
Indirect taxes (TIN)												
Detrended	86Q1-12Q4	4.3	0.52	**0.54**	0.47	0.41	0.41	0.21	0.07	-0.06	-0.26	Pro-cyclical lead
Detrended	86Q1-07Q4	3.8	0.25	0.30	0.26	0.28	**0.44**	0.30	0.25	0.18	-0.08	Pro-cyclical cont.
Pre-whitened	86Q1-12Q4	5.9	0.18	0.06	-0.06	0.09	**0.42**	0.15	0.12	-0.04	-0.19	Pro-cyclical cont.
Pre-whitened	86Q1-07Q4	5.8	0.13	-0.05	-0.24	-0.09	**0.35**	0.11	0.25	0.19	0.07	Pro-cyclical cont.
Social Security Contributions (SCT)												
Detrended	86Q1-12Q4	1.4	0.03	0.08	0.18	0.24	0.36	0.37	0.38	0.42	**0.46**	Pro-cyclical lag.
Detrended	86Q1-07Q4	1.8	-0.07	-0.07	0.01	0.04	0.19	0.22	0.29	0.39	**0.51**	Pro-cyclical lag.
Pre-whitened	86Q1-12Q4	1.8	-0.22	-0.15	-0.04	0.13	**0.38**	0.16	0.03	-0.04	-0.03	Pro-cyclical cont.
Pre-whitened	86Q1-07Q4	2.1	**-0.25**	**-0.22**	-0.14	-0.05	**0.21**	0.07	0.06	0.06	0.05	Borderline

[‡] Notes: Variables are transformed in real terms (GDP deflator) before filtering. Quarterly real GDP and GDP deflator are taken from the Banco de España database. The table shows averages over the five detrending methods used.

Table 3: *Stylized facts on fiscal policies II (Main expenditure aggregates). Correlations between cyclical components of the fiscal variable (in real terms) and real GDP.*‡

		$\frac{\sigma_y}{\sigma_x}$	-4	-3	-2	-1	0	1	2	3	4	
Social transfers other than unemployment benefits (THN - UNB)												
Detrended	86Q1-12Q4	1.1	**-0.35**	**-0.39**	-0.30	-0.20	-0.12	-0.02	0.01	0.11	0.20	Count.-cyc. lead
Detrended	86Q1-07Q4	1.3	-0.06	-0.14	-0.07	0.04	0.12	0.23	0.23	0.34	**0.43**	Pro-cyclical lag.
Pre-whitened	86Q1-12Q4	1.7	0.00	-0.16	-0.16	-0.06	0.01	0.11	0.10	0.08	0.11	A-cyclical
Pre-whitened	86Q1-07Q4	2.0	0.08	-0.18	-0.18	-0.07	0.04	0.16	0.13	0.09	0.10	A-cyclical
Unemployment benefits (UNB)												
Detrended	86Q1-12Q4	6.7	-0.52	-0.56	-0.54	**-0.59**	-0.47	-0.41	-0.27	-0.10	0.04	Count.-cyc. lead
Detrended	86Q1-07Q4	6.1	**-0.41**	**-0.36**	**-0.28**	**-0.33**	-0.18	-0.19	-0.12	-0.02	0.05	Count.-cyc. lead
Pre-whitened	86Q1-12Q4	6.5	0.01	-0.06	-0.14	**-0.37**	-0.09	-0.22	-0.16	0.04	-0.08	Count.-cyc. lead
Pre-whitened	86Q1-07Q4	6.7	-0.02	0.05	0.08	-0.14	0.05	-0.23	**-0.28**	-0.13	-0.17	Count.-cyc. lag.
Government consumption (GCR)												
Detrended	86Q1-12Q4	1.2	-0.13	-0.05	0.14	0.20	0.31	0.25	0.26	0.26	**0.37**	Pro-cyclical lead
Detrended	86Q1-07Q4	1.4	-0.05	0.08	0.33	0.39	**0.48**	0.34	0.30	0.24	0.38	Pro-cyclical cont.
Pre-whitened	86Q1-12Q4	1.7	-0.13	-0.20	0.05	-0.03	**0.30**	0.05	-0.04	-0.07	0.19	Pro-cyclical cont.
Pre-whitened	86Q1-07Q4	2.0	-0.17	-0.10	0.28	0.15	**0.38**	0.02	-0.07	0.02	0.35	Pro-cyclical cont.
Government investment (GIN)												
Detrended	86Q1-12Q4	6.6	0.13	0.18	0.23	0.25	0.26	0.25	0.28	0.30	**0.38**	Pro-cyclical lag.
Detrended	86Q1-07Q4	6.1	0.38	0.41	0.45	**0.46**	0.45	0.42	0.42	0.38	0.41	Pro-cyc. cont./lead
Pre-whitened	86Q1-12Q4	5.4	**0.18**	0.12	0.15	0.10	-0.03	**-0.23**	**-0.22**	-0.15	0.16	Borderline
Pre-whitened	86Q1-07Q4	6.3	**0.27**	0.23	0.23	0.14	0.04	-0.16	-0.13	-0.08	0.20	Pro-cyclical lead

‡ Notes: Variables are transformed in real terms (GDP deflator) before filtering. Quarterly real GDP and GDPdeflator are taken from the Banco de España database. The table shows averages over the five detrending methods used.

Table 4: *Stylized facts on fiscal policies III (Other expenditure items). Correlations between cyclical components of the fiscal variable (in real terms) and real GDP.*[‡]

Panel A. Government consumption components.

		$\frac{\sigma_y}{\sigma_x}$	-4	-3	-2	-1	0	1	2	3	4	
Compensation of public employees (COE)												
Detrended	86Q1-12Q4	1.6	-0.30	-0.30	-0.19	-0.09	0.07	0.15	0.23	0.31	**0.43**	Pro-cyclical lag.
Detrended	86Q1-07Q4	1.7	-0.11	-0.11	0.04	0.14	0.32	0.40	0.45	0.51	**0.58**	Pro-cyclical lag.
Pre-whitened	86Q1-12Q4	1.5	-0.16	**-0.20**	-0.14	-0.04	**0.17**	0.12	-0.02	0.12	0.12	Borderline
Pre-whitened	86Q1-07Q4	1.4	-0.15	-0.24	-0.19	-0.12	0.15	0.20	0.15	**0.28**	0.20	Pro-cyclical lag.
Government employment (LGN)												
Detrended	86Q1-12Q4	0.8	0.19	0.20	0.24	0.26	0.28	0.27	0.26	**0.29**	0.27	Pro-cyclical lag.
Detrended	86Q1-07Q4	0.9	0.38	0.38	0.40	0.42	**0.43**	0.41	0.36	0.36	0.27	Pro-cyclical cont.
Pre-whitened	86Q1-12Q4	0.7	0.03	-0.02	-0.04	0.05	**0.10**	-0.04	**-0.11**	0.02	-0.06	A-cyclical
Pre-whitened	86Q1-07Q4	0.9	0.04	0.03	-0.01	0.09	**0.13**	0.00	-0.05	0.08	-0.03	A-cyclical
Non-wage government consumption (OGCN)												
Detrended	86Q1-12Q4	2.5	-0.07	-0.06	0.08	0.09	**0.19**	0.05	0.07	0.03	0.18	Borderline
Detrended	86Q1-07Q4	3.2	-0.06	-0.03	0.16	0.16	**0.24**	0.01	-0.02	-0.10	0.09	Pro-cyclical cont.
Pre-whitened	86Q1-12Q4	4.5	-0.06	**-0.31**	-0.02	-0.07	**0.23**	-0.03	-0.08	-0.11	0.10	Count.-cyclical lead
Pre-whitened	86Q1-07Q4	5.2	-0.26	**-0.38**	0.12	0.10	**0.34**	0.02	-0.06	-0.05	0.19	Count.-cyclical lead

Panel B. Other government expenditure components.

		$\frac{\sigma_y}{\sigma_x}$	-4	-3	-2	-1	0	1	2	3	4	
Interest payments (INP)												
Detrended	86Q1-12Q4	5.4	-0.24	-0.28	-0.33	**-0.39**	**-0.35**	**-0.39**	-0.33	-0.30	-0.27	Count.-cyc. lag.
Detrended	86Q1-07Q4	6.5	-0.17	-0.25	-0.34	**-0.42**	**-0.37**	**-0.44**	-0.35	-0.28	-0.20	Count.-cyc. lag.
Pre-whitened	86Q1-12Q4	6.3	0.02	-0.15	-0.17	-0.15	0.00	**-0.18**	0.07	0.14	0.12	Borderline
Pre-whitened	86Q1-07Q4	7.2	-0.04	-0.21	-0.23	-0.21	-0.03	**-0.26**	0.09	0.13	0.14	Count.-cyc. lag.
Subsidies (SIN)												
Detrended	86Q1-12Q4	10.2	0.08	0.01	-0.05	-0.06	-0.10	0.02	0.02	0.11	0.10	A-cyclical
Detrended	86Q1-07Q4	11.9	0.05	-0.06	-0.18	**-0.23**	**-0.31**	-0.11	-0.11	0.02	-0.06	Count.-cyc. cont.
Pre-whitened	86Q1-12Q4	19.3	0.05	0.11	0.01	-0.05	**-0.22**	0.01	-0.05	0.02	-0.05	Count.-cyc. cont.
Pre-whitened	86Q1-07Q4	21.2	0.17	0.11	-0.06	-0.17	**-0.30**	0.05	0.05	0.07	-0.11	Count.-cyc. cont.
Other public expenditure (OTOE)												
Detrended	86Q1-12Q4	4.7	0.20	0.20	0.24	**0.24**	0.22	0.23	0.19	0.17	0.20	Pro-cyc. Lead
Detrended	86Q1-07Q4	6.0	**0.36**	0.29	0.26	0.19	0.13	0.12	0.06	0.04	0.09	Pro-cyc. Lead
Pre-whitened	86Q1-12Q4	7.4	0.04	0.00	0.12	0.11	0.13	0.07	-0.14	-0.08	0.11	A-cyclical
Pre-whitened	86Q1-07Q4	8.9	**0.26**	0.13	0.10	-0.07	-0.10	-0.09	-0.18	-0.02	0.20	Pro-cyc. Lead

[‡] Notes: Variables are transformed in real terms (GDP deflator) before filtering. Quarterly real GDP and GDP deflator are taken from the Banco de España database. The table shows averages over the five detrending methods used.

Finally, as regards government net lending, i.e. the difference between total revenues and total expenditures, it appears clear from Table 2 that the pro-cyclical correlation has increased with the most recent crisis, as well as its volatility. The dominant correlation ranges from 0.4 (pre-crisis) to 0.7 (full sample). In general, it can be said that government balances act as a shock absorber, but are not more volatile that several expenditure or receipt components. The pro-cyclicality found is in line with the results for the G7 countries of Fiorito (1997), and of Galí and Perotti (2003) for Europe. As discussed in Lamo et al. (2013), from a theoretical point of view, our results would render some empirical support to models that predict pro-cyclical fiscal policies. Among these are political economy models (see the references in the introductory section, Fernández de Córdoba et al., 2012, and the references quoted therein) that tend to rationalize pro-cyclicality on the grounds that bureaucrats, or governments in general, maximize the available budget for wage-related public spending, which creates boom-bust-like dynamics as government revenue windfalls in good times are spent in full, while fiscal consolidation needs in adverse economic circumstances force a cut in wage and non-wage government spending. Another branch of models exploit market imperfections to justify the existence of pro-cyclicality in fiscal policies (see, among others, Mendoza and Oviedo, 2006). Also, a classical explanation is the one of Alesina and Tabellini (2005), in which "less-than-benevolent" governments would have incentives to appropriate rents and as such voters would demand more public goods and fewer taxes to prevent the latter from happening when the economy is doing well.

4 The macroeconomic impact of non-systematic fiscal policy

4.1 The existing evidence for Spain

For the case of Spain, as mentioned in the introduction, some recent papers have dealt with the matter of estimating fiscal multipliers within standard VAR frameworks. On the one hand, the papers by de Castro (2006), de Castro and Hernández de Cos (2008), de Castro and Fernández-Caballero (2013), and European Commission (2012), follow a linear SVAR approach and use similar identification approaches. On the other hand, Hernández de Cos and Moral-Benito (2013) move one step forward and estimate STVAR models to compute regime-dependent fiscal multipliers. The latter approach is in line with the traditional Keynesian view that, given slack resources in the economy, fiscal policy may be more effective at increasing output in recessions than during normal times. Under this hypothesis, most studies averaging the multiplier over the cycle would under-estimate

Table 5: *Cumulative government spending ($G = GCN + GIN$) output multipliers: summary of selected empirical results (linear SVARs) from the literature.*

	Sample	Database	Cumulative multiplier[a]			Variables included in the VAR[b]
			H=1	H=4	H = 8	
de Castro (2006)	81Q1-01Q4	MTBE	–	1.14*	1.04*	G, NT, GDP, i^{3m}, P
de Castro and Hdez de Cos (2008)	81Q1-04Q4	MTBE	–	1.31*	1.33*	G, NT, GDP, i^{3y}, P
de Castro and Fndez-Caballero (2013)	81Q1-08Q4	MTBE/IGAE[c]	0.41*	0.94*	0.95*	G, NT, GDP, i^{3y}, P, E
	89Q1-08Q4	MTBE/IGAE[c]	0.49*	1.36*	1.98*	G, NT, GDP, i^{3y}, P, E
European Commission (2012)	86Q1-10Q4	QESFIP[d]	0.30*	1.15*	1.79*	G, NT, GDP, i^{3y}, P
Hdez de Cos and Moral-Benito (2013)	86Q1-11Q4	QESFIP[e]	0.65*	0.60*	0.30	G, NT, GDP
"Turbulent times": crisis[f]	86Q1-11Q4	QESFIP[e]	0.84*	1.48*	1.30*	G, NT, GDP
"Turbulent times": no crisis[f]	86Q1-11Q4	QESFIP[e]	0.79*	0.64*	0.60*	G, NT, GDP

[a] The cumulative multiplier at each horizon H is defined as $\frac{\sum j=1^H \Delta GDP_i}{\sum j=1^H \Delta G_i}$. Baseline specifications are taken. An asterisk indicates significance within a 68% confidence interval.

[b] NT: net taxes; $i^{3y/m}$: 3 year/month interest rate on gov. bonds; P: GDP def.; E: real effec. exchange rate.

[c] MTBE series for the period 1981Q1-1999Q4. IGAE for 2000Q1-2008Q4 (seasonal adjustment performed by the authors).

[d] Beta version of the database with cut-off date March 2012. National accounts' revisions occurred since then, in particular in Oct. 2013 (revision of 95Q1-12Q4 tax series due to the change in the statistical criteria to compute refunds of taxes).

[e] Beta version of the database with cut-off date January 2013.

[f] STVAR model. Principal components index PCA II, Panel A of Table 4.

its size in recessions. Being an interesting and suggestive idea from a theoretical point of view, its empirical implementation via STVARs along the lines of Auerbach and Gorodnichenko (2012; 2013) has been recently subject also to a number of critiques, linked in particular to robustness issues (large number of parameters to be estimated that limits, as a consequence, the set of variables considered).[16]

[16] The STVAR approach involves highly nonlinear estimation of a large number of parameters, which is very challenging in terms of numerical computation. Auerbach and Gorodnichenko (2012; 2013) overcome the challenge using quarterly US data over the postwar period 1947-2008, i.e., 248 observations. For the case of Spain the sample lenght is more limited (108). The shortcomings of data might represent a concern in terms of the convergence properties of the numerical methods employed for the likelihood maximization required by the STVAR approach. Moreover, the fact that a large number of parameters is estimated tends to restrict the number of variables that are included in the VAR, which may bias the results if, for example, interest rates or inflation are excluded, as it is typically the case. Finally, as discussed by Ramey and Zubairy (2013), constructing impulse responses in nonlinear models is far from straightforward. In the STVAR approach the impulse response functions are constructed under the assumption that the shocks cannot modify the state of the economy, i.e. it is implicitly assumed that the economy remains in the same regime, either expansion or recession, over the entire horizon for which IRFs are computed. These assumptions might lead to biases in the estimated IRFs.

Fine tuning country-specific estimates is a crucial issue to draw policy lessons, given that the most recent literature has stressed the significant heterogeneity of estimates derived from general theoretical and empirical models (see European Commission, 2012; Favero et al., 2011). In Table 5 we summarize some selected results from the aforementioned papers as regards the impact on real GDP of a government spending (defined as GCN + GIN) shock. Most papers look at the pre-crisis period, and find cumulative government spending multipliers of around 1 to 1.4 after four quarters, that increase up to 1 to 1.9 after eight quarters. Papers including post 2008Q1 figures show somewhat divergent results for linear multipliers. Results may depend, basically, on the sample considered, the dataset, and the variables included in the empirical model, given that identification methods are quite homogeneous across papers.

In the next subsection we update the estimates of fiscal multipliers available for Spain, in order to deepen our understanding of a number of issues. First, we include the most recent financial crisis episode (2008Q1-2012Q4), in order to assess its impact on estimated linear multipliers, by comparing the full sample with the sample excluding the post 2008Q1 period. Second, we test the sensitivity of estimated multipliers to the exclusion of relevant variables like prices and interest rates, that tend to be mandatory in non-linear SVAR approaches, given that they are subject to "the curse of dimensionality". Finally, we test the sensitivity and robustness of the estimated multipliers to alternative datasets (QESFIPDB, REMS, MTBE), to grasp some intuition on whether empirical results may end up being influenced by that choice.

4.2 Empirical exercises

The baseline VAR includes quarterly data on public expenditure ($G_t = GCN + GIN$), net taxes ($NT_t = TOR - THN - INP$) and GDP (y_t), all in real terms,[17] the GDP deflator (P_t) and the three-year interest rate of government bonds (i_t). All variables are seasonally adjusted and enter in logs except the interest rate, which enter in levels. In addition, we also add in certain specifications the level of public debt, given that this variable has been signalled as being of relevance for the estimation of multipliers, in particular in periods of fiscal stress (see Favero and Giavazzi, 2007). The VARs are estimated for the period 1986Q1 to 2012Q4. The GDP volume and its deflator have been taken from the Quarterly National Accounts (National Institute of Statistics, INE) while the three-year bond rate has been obtained from the Banco de España database. The SVAR approach in this paper is completely standard and follows as such the seminal contributions of Blanchard

[17]The nominal variables have been deflated by the GDP deflator in order to obtain the corresponding real values.

and Perotti (2002) and Perotti (2004). For further details see Appendix D. We show the results of our empirical exercises[18] in Table 6, and figures 5 and 6.

When comparing the full sample (1986Q1-2012Q4) with the sample excluding the post 2008Q1 period, the following results are worth highlighting. First, government spending shocks' cumulative multipliers (panel A of Table 6 and Figure 5) are marginally higher after one and two years, but not in the medium term. At the same time, when public debt is included in the models as an additional variable, the responses are stronger than in the baseline case, and also the increase in the quantitative point estimates when comparing samples. Allowing net taxes and government spending to respond to the public debt level, in a period of fiscal stress, induces a more muted effect on the interest rates in response to the fiscal shock. This is a channel put forward by Favero and Giavazzi (2007) in a similar set-up to ours. Second, turning to the components of "government spending" (panel B of Table 6, and Figure 5), the cumulated output multiplier associated to public consumption decreases after one year when the whole sample is considered. Interestingly, this aggregated result unveils a different behavior of output in response to the non-wage and the wage components. On the one hand, non-wage government consumption multipliers are higher when estimated over the whole sample. This is consistent with the presence of a higher output multiplier in bad times, given also the direct effect on private demand of this component. On the other hand, though, the responses of output to wage bill shocks change substantially when the 2008Q1-2012Q4 quarters are included. The channel through which government personnel expenditure affects the economy is not only linked to the direct effect on private demand, but also to the effect through the labor market. In this respect, after an initial positive impact in the first year (multiplier below one), a negative and significant effect on output is observed, which can be explained by the potentially negative effects on private investment profitability stemming from higher personnel spending by the general government sector (Alesina et al., 2002) or by signaling effects on private sector wages

[18]The potential endogeneity of structural shocks may affect our results. To address this potential problem, we test that indeed there is no co-movement between the variables included in the VAR and the estimated structural shocks. Along the lines of Favero et al. (2011) and Auerbach and Gorodnichenko (2013), among others, we have then regressed the structural shocks on lags of the full information set used in our VAR specification. The results (available upon request) show that indeed the shocks are orthogonal to lags of the full information set, and mostly orthogonal if contemporaneous information is included. Additionally, we have looked at alternative concepts of exogeneity. Particularly, those referred as "block-exogeneity" tests which analyze, on the basis of Granger causality tests, whether or not a variable is independent of the others (see Engle et al., 1983, for a formal description). Our results confirm the exogeneity of the estimated structural shocks.

Table 6: *Cumulative output multipliers of fiscal shocks.*[a]

A. GOVERNMENT SPENDING SHOCK	Quarters 1	4	8	12	16	20
– Sample 1986Q1 - 2012Q4						
Baseline	1.23*	2.12*	2.06*	1.58*	1.15*	0.74*
Baseline: 3 variables	1.11*	1.73*	1.54*	1.17*	0.72*	0.21
Model with debt	1.34*	2.67*	2.76*	2.20*	1.68*	1.27*
– Sample 1986Q1 - 2007Q4						
Baseline	1.25*	1.94*	1.87*	1.51*	1.13*	0.80*
Baseline: 3 variables	1.25*	1.84*	1.50*	1.17*	0.76*	0.31
Model with debt	1.20*	2.03*	2.14*	1.86*	1.50*	1.08*
REMS database	-0.64*	-0.51*	-0.24	-0.56	-1.27	-2.30
MTBE database	0.51*	1.03*	1.39*	1.06*	0.64	0.20

B. SPENDING COMPONENTS' SHOCKS	Quarters 1	4	8	12	16	20
– Sample 1986Q1 - 2012Q4						
Public consumption	1.45*	2.09*	1.58*	0.52	-0.59	-1.95
– Purchases	1.64*	3.20*	3.23*	1.54	-0.75	-3.9
– Personnel expenditure	1.36*	0.73*	-0.56	-2.24*	-4.36*	-8.07*
Public investment	-0.20	0.47*	1.02*	0.84	0.01	-1.58
– Sample 1986Q1 - 2007Q4						
Public consumption	1.47*	2.04*	1.87*	1.36*	0.84*	0.39
– Purchases	1.63*	2.55*	2.02*	1.15*	0.2	-0.78
– Personnel expenditure	1.52*	2.23*	2.73*	3.63*	3.81	3.60
Public investment	0.02*	0.78*	1.57*	1.58*	0.97*	0.28

C. NET TAXES SHOCK	Quarters 1	4	8	12	16	20
– Sample 1986Q1 - 2012Q4						
Baseline	0.25*	0.26*	0.43*	0.59*	0.72	0.87
Baseline: 3 variables	0.26*	0.26*	0.43*	0.64*	0.84*	1.01*
– Sample 1986Q1 - 2007Q4						
Baseline	0.25*	0.19*	0.71*	0.92*	1.20*	1.23*
Baseline: 3 variables	0.17*	0.02	0.15	0.21	0.29	0.3
REMS database	-0.32*	0.43*	1.18*	1.33*	1.08*	0.83*
MTBE database	0.52*	0.99*	1.01*	0.64*	0.26	-0.14

[a] The cumulative multiplier at each horizon H is defined as $\frac{\sum j=1^H \Delta GDP_i}{\sum j=1^H \Delta G_i}$. Baseline specifications are taken. An asterisk indicates significance within a 68% confidence interval.

(Lamo et al., 2012). So far, the results are in line with the flavor of those obtained in previous papers (de Castro, 2006, de Castro and Hernández de Cos, 2008) though much more clear-cut, in particular as regards the evidence on the short- to medium-term output effect of the wage bill.[19] Finally, it is worth mentioning that public investment shocks lead to a weaker reaction by GDP when the 2008-2012 years are added to the sample. In fact, the sizeable investment packages implemented to smooth the effects of the crisis in 2009-2010, mainly targeted to projects with a limited or null impact on potential growth, were accompanied by a collapse in economic activity, which may explain the lower multipliers. When the sample period is constrained to end in the last quarter of 2007 public investment shocks lead to more significant GDP increases, in accordance with previous empirical evidence for Spain.

A second issue we wanted to explore is the sensitivity of estimated multipliers to the exclusion of relevant variables, in particular prices and interest rates, given that this is typically a must in non-linear SVAR approaches, in order to reduce the dimensionality of the number of parameters to be estimated. The results of a tentative exercise are presented in panel A of Table 6. In the three-variables canonical SVAR model government spending multipliers are lower than those obtained with five-variables, standard, SVARs. Lower multipliers arise mainly because of the feedback effects/responses of interest rates and prices to the fiscal shock. This may help in rationalizing part of the differences in linear SVAR estimates observed in Table 6 between the multipliers reported by Hernández de Cos and Moral-Benito (2013) and, for example, European Commission (2012), and may also give an indication of potential biases in estimates obtained with non-linear models that do exclude relevant variables.

Finally, we test the sensitivity and robustness of the estimated multipliers to alternative datasets (QESFIPDB, REMS, MTBE) in a few cases, to grasp some intuition on whether empirical results may end up being influenced by that choice. We restrict the comparison to the 1986Q1-2007Q4 period for which we have coverage from the three datasets at hand. First, it is worth noticing that the government spending multipliers estimated with our dataset are significantly higher and more persistent than those obtained with the MTBE or the REMS (in the short-term) databases (panel A of Table 6.In addition, it seems that using REMS data always lead to less precise responses of all variables, in view of the width of confidence bands. As regards the responses of GDP to a shock to net taxes, they seem to be in line in the three datasets, and also similar to other studies

[19] In the latter respect, though, the evidence with the sample until 2007Q4 regarding shocks to the public sector wag bill is in contrast with the aforementioned evidence for Spain.

for Spain and other OECD countries, even thought the MTBE response is less persistent. GDP rises in response to higher net taxes, with the exception of the contemporaneous reaction in the REMS case. Admittedly, the increase is at odds with almost any theoretical model. This pattern is observed in the short term by Perotti (2004), de Castro and Hernández de Cos (2008) or Heppke-Falk et al. (2006), among others, which probably reveals the difficulty to identify net tax shocks properly. In all cases, government spending declines and recovers later on, standing significantly above the baseline in the medium term. This positive reaction of expenditure seems in line with the tax-and-spend view of fiscal policy.[20]

5 Conclusions

In this paper we provide a comprehensive database of quarterly fiscal variables suitable for macroeconomic analysis built up on the basis of state-of-the-art macroeconometric models. All models are multivariate, state space mixed-frequencies models, estimated with available national accounts fiscal data (mostly annual) and, more importantly, monthly and quarterly information taken from all available sources of fiscal data. The database spans over the period 1986Q1-2012Q4, and covers a wide number of fiscal aggregates, suitable for macroeconomic analysis. All the time series included are presented in gross (non-seasonally adjusted) and seasonally adjusted terms. We focus solely on intra-annual fiscal information for interpolation purposes. This approach allows us to capture genuine intra-annual "fiscal" dynamics in the data, so that we avoid two important problems that are present in fiscal time series interpolated on the basis of general macroeconomic indicators: (i) the endogenous bias that arises if the so interpolated fiscal series were used with macroeconomic variables to assess the impact of fiscal policies; (ii) the well-known decoupling of tax collection from the evolution of macroeconomic tax bases (revenue windfalls/shortfalls). On the basis of our quarterly fiscal database we provide in the paper a number of applications that highlight its usefulness for macroeconomic analysis and policy.

Firstly, we provide some stylized facts on the cyclical properties of fiscal policies. We find that total revenues in Spain display a pro-cyclical behavior, that can be to a large extent explained by discretionary changes in policy (unpredictable component), and are much more volatile than GDP, most likely due to the fact that a number of taxes, most notably corporate taxes, property taxes and other indirect taxes, tend to follow boom-bust dynamics, and also to the progressive structure

[20] de Castro et al. (2004) provide some evidence supporting this view for Spain.

Figure 5: *Response to government spending shocks: government consumption (GCN) and investment (GIN) aggregate, government consumption – including components: COE and OGCN – and government investment. Baseline sample: 1986Q1-2012Q4.*

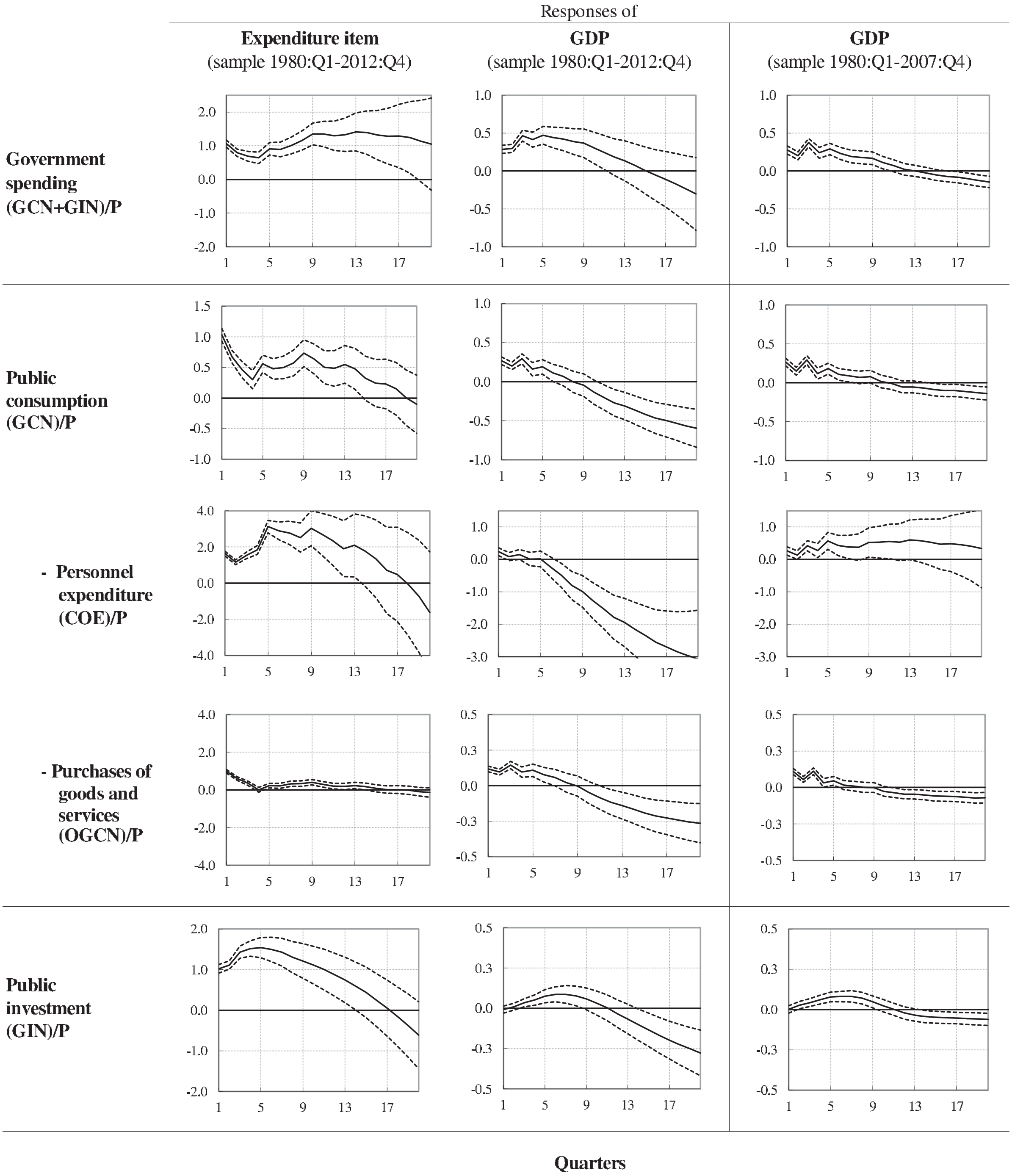
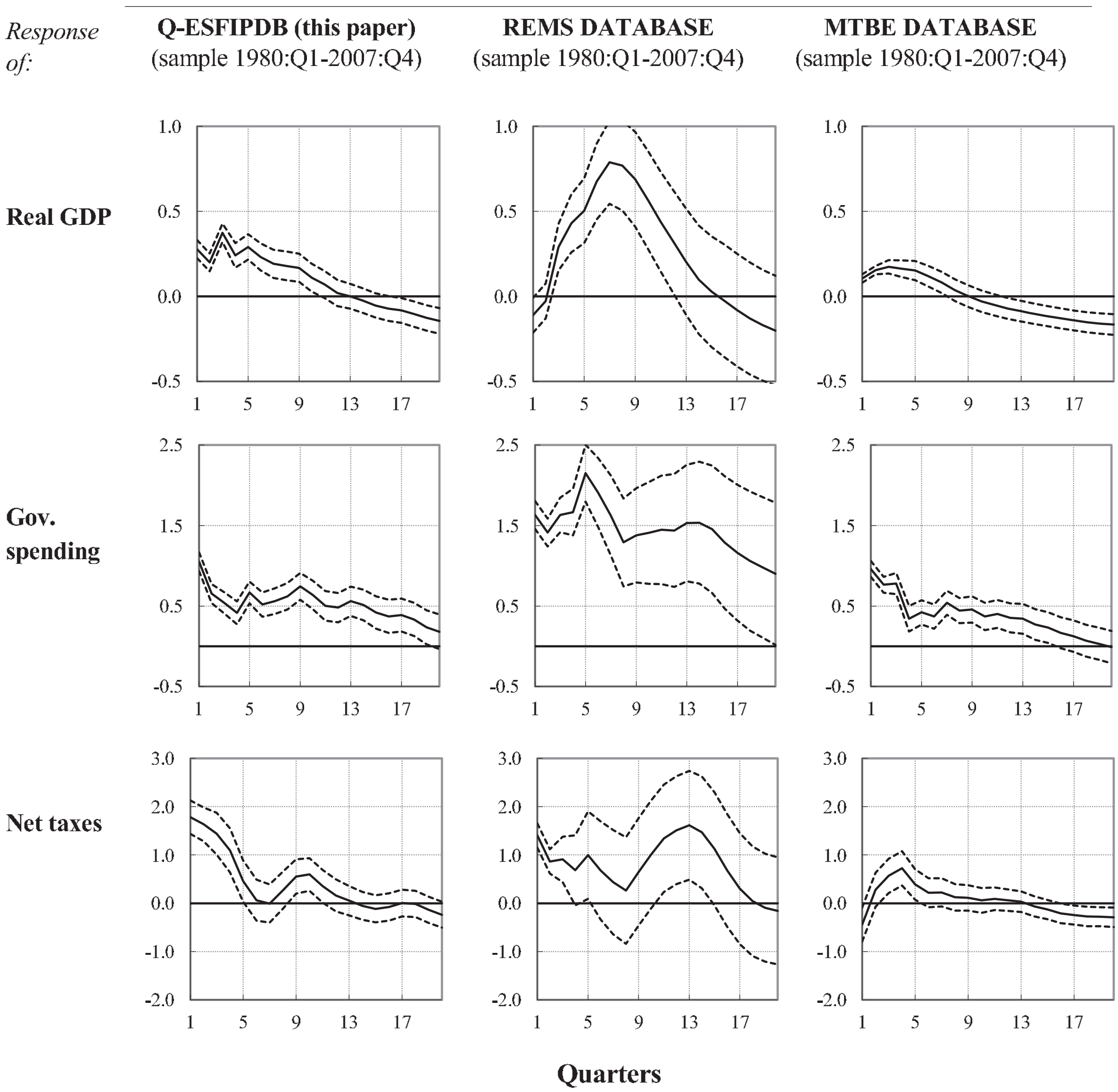


Figure 6: *Response to government spending shocks, alternative databases.*



of the income tax. Some studies have hinted at pro-cyclical revenues as a source of pro-cyclical government spending. In fact, we find that total expenditure appears pro-cyclical as well, but lagged. The pro-cyclical pattern of total expenditures is due to the government consumption and investment components. Social transfers, particularly unemployment-related expenditure, on the contrary, present a distinct counter-cyclical behavior. Public spending, overall, is found to be more volatile that real GDP, and higher than comparable euro area reference variables.

Secondly, we run standard SVAR models and provide some updated estimates and insights on the impact of changes in fiscal aggregates on macroeconomic variables. First, when including the most recent financial crisis episode (2008Q1-2012Q4), government spending shocks are marginally higher, in particular when taxes and spending are allowed to respond to public debt. As regards components, the negative medium-term effect of wage bill shocks gets clearly profiled, compared to less clear-cut results in other studies, while at the same time the positive impact of non-wage government shocks increased with the crisis. Second, we show that estimated multipliers may be sensitive to the exclusion of relevant variables like prices and interest rates, and as a consequence results from models more computationally-intensive models that are force to drop relevant variables, would have to be read from this perspective. Finally, we test the sensitivity and robustness of the estimated multipliers to alternative datasets (QESFIPDB, REMS, MTBE): it seems that the choice of the dataset might not be neutral. As regards the latter, thus, we feel reassured with our approach of trying to advance the frontier of data availability for the profession, by improving in a transparent way the methodology and inputs used to overcome the limitations of available official statistics to carry out studies on the macroeconomic impact of fiscal policies.

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A Main elements of the quarterly fiscal database

Data inputs. The variables of interest are quarterly general government accounts on an ESA95 basis. Quarterly, non seasonally adjusted figures are available from 2000 onwards, as mentioned above. Annual ESA95 data fully consistent with the quarterly figures start in 1995. Nevertheless, annual ESA79 figures are available from 1985 to 1998. Therefore, we extended our annual ESA95 dataset backwards by means of the annual growth rates of ESA79 figures. The corresponding annual values are used as an anchor for the quarterly interpolation procedure. As regards short-term indicators, we use national accounts and cash data for different revenue and expenditure items available for the different sub-sectors and public entities, at quarterly and monthly frequencies as indicators to interpolate the quarterly ESA95 missing values. Quarterly and monthly fiscal variables (indicators) are also taken from the IGAE and the Ministry of Employment (State Secretary of the Social Security). For the Central government and the Social Security subsectors, short-term public finance statistics in Spain are published timely, with a broad coverage of budgetary categories. For the former, monthly figures in ESA95 (NA) standards covering all the relevant revenue and expenditure details are published within one month while for the latter monthly cash figures are made available with a short delay and cover both the Social Security System and the Public Employment System. We also use some derived indicators of regional governments' spending. On available data for the sub-national governments and the developments of indicators suitable for short-term forecasting in Spain, see Fernández-Caballero, Pedregal and Pérez (2012).

Empirical methodology. The basic model is of the Unobserved Component Model class known as the Basic Structural Model (Harvey, 1989), that decomposes a set of time series in unobserved though meaningful components from an economic point of view (mainly trend, seasonal and irregular). Given that the data used are sampled at different frequencies (annual, quarterly and monthly) we use a mixed-frequencies formulation, along the lines of Harvey and Chung (2000), Moauro and Savio (2005), and Proietti and Moauro (2006). The modeling approach is described in detail in Appendix B.

Output database. The list of variables' names is shown in Table 1. We provide a quite disaggregated set of nominal fiscal variables for the General Government sector in ESA95 terms, seasonally and non-seasonally adjusted. The issue of seasonal adjustment of quarterly fiscal variables in Europe is an important one, as signalled in European Commission (2007).[21] On the revenue side of government accounts the database covers total government revenue, direct taxes (with a proxy for the breakdown between direct taxes paid by households and firms), social security contributions (with a proxy for the breakdown between contributions paid by employers and others), and total indirect taxes. On the expenditure side, it covers total expenditure, social payments (of which also unemployment benefits), interest payments, subsidies, government investment and government consumption. Given the relevance of the latter variable (part of the demand side of GDP), we provide the breakdown between nominal and real government consumption, the breakdown between government wage and non-wage consumption expenditure, and government employment. The net lending of the government, a key policy variable, can be computed as the difference between total revenues and total expenditures. We also provide quarterly public debt for the period of reference.

B Methodology to compute the quarterly fiscal database

The exposition in this section follows closely Pedregal and Pérez (2010). The starting point of the modeling approach is to consider a multivariate Unobserved Components Model known as the Basic Structural Model (Harvey, 1989). A given time series is decomposed into unobserved components which are meaningful from an economic point of view (trend, T_t, seasonal, S_t, and irregular, e_t). Equation (B1) displays a general form, where t is a time sub-index measured in quarters, z_t denotes the variable in ESA95 terms expressed at an annual and quarterly sampling interval (depending on availability) for our objective time series, and u_t represents the vector of quarterly indicators.

$$\begin{bmatrix} \mathbf{z}_t \\ \mathbf{u}_t \end{bmatrix} = \mathbf{T}_t + \mathbf{S}_t + \mathbf{e}_t \tag{B1}$$

The general consensus in this type of multivariate models in order to enable identifiability is to build SUTSE models (Seemingly Unrelated Structural Time Series). This means that components of the same type interact among them for different time series, but are independent of any of the components of different types. In addition, statistical relations are only allowed through the covariance structure of the vector noises, but never through the system matrices directly. This

[21] The type of models that we use encompass the estimation of seasonal components, and so it is possible to recover model-consistent seasonally-adjusted series. Nevertheless, the companion database we provide in Excel format includes seasonally-adjusted series based on the well-known program TRAMO-SEATS (Gómez and Maravall, 1996), given that this is the standard method used by many National Statistical Institutes to seasonally-adjust macroeconomic time series.

allows that, trends of different time series may relate to each other, but all of them are independent of both the seasonal and irregular components. The full model is a standard BSM that may be written in State-Space form as (see Harvey, 1989)

$$\mathbf{x}_t = \mathbf{\Phi x_{t-1}} + \mathbf{Ew_t} \tag{B2}$$

$$\begin{bmatrix} \mathbf{z}_t \\ \mathbf{u}_t \end{bmatrix} = \begin{bmatrix} \mathbf{H} \\ \mathbf{H}^u \end{bmatrix} \mathbf{x}_t + \begin{bmatrix} \epsilon_t \\ \mathbf{v}_t \end{bmatrix} \tag{B3}$$

where $\epsilon_t \sim N(0, \Sigma_\epsilon)$ and $\mathbf{v}_t \sim N(0, \Sigma_{\mathbf{v}_t})$. The system matrices $\mathbf{\Phi}$, $\mathbf{E}$, $\mathbf{H}$ and $\mathbf{H}^u$ in equations (B2)-(B3) include the particular definitions of the components and all the vector noises have the usual Gaussian properties with zero mean and constant covariance matrices (ϵ_t and $\mathbf{v}_t$ are correlated among them, but both are independent of $\mathbf{w}_t$). The particular structure of the covariance matrices of the observed and transition noises defines the structures of correlations among the components across output variables. The structure of system matrices in (B2)-(B3) is as follows. Let m be the number of variables $\mathbf{z}_t$ and $\mathbf{u}_t$ altogether; $\mathbf{I}$ and $\mathbf{0}$ are $m \times m$ identity and zero matrices, respectively. Then $\mathbf{\Phi}$ is a matrix formed by the block concatenation of the matrices

$$\mathbf{\Phi}_0 = \begin{bmatrix} \mathbf{I} & \mathbf{I} \\ \mathbf{0} & \mathbf{I} \end{bmatrix}, \mathbf{\Phi}_k = \begin{bmatrix} \cos\frac{2\pi k}{12} & \sin\frac{2\pi k}{12} \\ sin\frac{2\pi k}{12} & \cos\frac{2\pi k}{12} \end{bmatrix} \otimes \mathbf{I},\ k = 1, 2, \ldots, 6 \tag{B4}$$

$\mathbf{E}$ is an identity matrix of appropriate dimension and

$$\begin{bmatrix} \mathbf{H} \\ \mathbf{H}^u \end{bmatrix} = \begin{bmatrix} \mathbf{I} & \mathbf{0} & \mathbf{I} & \mathbf{0} & \ldots & \mathbf{I} & \mathbf{0} \end{bmatrix} \tag{B5}$$

The mixture of frequencies, and the estimation of models at the quarterly frequency, implies combining variables that at the quarterly frequency can be considered as stocks with those being pure flows. Thus, given the fact that our objective variables are observed at different frequencies, an accumulator variable has to be included

$$C_t = \begin{cases} 0, & t = \text{first quarter} \\ 1, & \text{otherwise} \end{cases} \tag{B6}$$

so that the previous model turns out to be

$$\begin{bmatrix} \mathbf{z}_t \\ \mathbf{x}_t \end{bmatrix} = \begin{bmatrix} \mathbf{I} \otimes C_t & \mathbf{H\Phi} \\ \mathbf{0} & \mathbf{\Phi} \end{bmatrix} \begin{bmatrix} \mathbf{z}_{t-1} \\ \mathbf{x}_{t-1} \end{bmatrix} + \begin{bmatrix} 1 & \mathbf{HE} \\ \mathbf{0} & \mathbf{E} \end{bmatrix} \begin{bmatrix} \epsilon_t \\ \mathbf{w}_t \end{bmatrix} \tag{B7}$$

$$\begin{bmatrix} \mathbf{z}_t \\ \mathbf{u}_t \end{bmatrix} = \begin{bmatrix} \mathbf{I} & \mathbf{0} \\ \mathbf{0} & \mathbf{H}^u \end{bmatrix} \begin{bmatrix} \mathbf{z}_t \\ \mathbf{x}_t \end{bmatrix} + \begin{bmatrix} \mathbf{0} \\ \mathbf{I} \end{bmatrix} \mathbf{v}_t \tag{B8}$$

Given the structure of the system and the information available, the Kalman Filter and Fixed Interval Smoother algorithms provide an optimal estimation of states. Maximum likelihood in the time domain provides optimal estimates of the unknown system matrices, which in the present context are just covariance matrices of all the vector noises involved in the model. The use of the models selected and the estimation procedures described in the previous paragraph, allows the estimation of models with unbalanced data sets, i.e. input variables with different sample lengths. This is a feature of relevance for the construction of the database at hand, given occasional differences in temporal coverage of indicators.

In our case, particular empirical specifications for each variable will be considered in the light of the available information (fiscal indicators). For instance, for the case of total government revenues, $\mathbf{z}$ comprises total government revenues in National Accounts terms, a variable that is available at the annual frequency from 1986-1999 and at the quarterly frequency from 2000Q1-2010Q4, while $\mathbf{u}$ is a matrix composed of three series (available at the quarterly frequency for the whole sample period): (i) a proxy to general government total revenues in public accounts (cash) terms; (ii) Central government total revenues and (iii) Social Security (SSS+SPEE) sector's total revenues. In order to reduce the dimensionality of our models and somewhat avoid the "curse of dimensionality" we opted for variable-by-variable models. By this we mean that, in all cases,$\mathbf{z}$ encompasses just one time series (annual/quarterly), and $\mathbf{u}$ the set of indicators corresponding to the latter variable, with a maximum of five indicators.

C Description of detrending methods

We consider a quite standard set of detrending methods:

- First order differencing takes the cycle to be the variable in first differences. Thus it assumes that the trend is the lagged variable, or similarly the series is a random walk with no drift. Therefore y_t can be represented as: $y_t = y_{t-1} + C_t + \epsilon_t$ where the trend is $T_t = y_{t-1}$ and an estimate of the detrended component is obtained as $y_t - y_{t-1}$.

- An alternative method of detrending is by removing a deterministic trend; the usual procedure is to take the least squares residual after regressing the series on a constant and a polynomial function of time. The implicit assumption is that the trend and cyclical components are orthogonal, and that T_t is a deterministic process which can be approximated with polynomial functions of time. These assumptions imply a model for y_t of the form: $y_t = T_t + C_t + \epsilon_t$, $T_t = f(t)$, and we take $f(t) = a_0 + a_1\ t + a_2\ t^2$. Even though the disturbance may be serially correlated, it can be shown that the unknown parameters in $f(t)$ can be estimated efficiently by ordinary least squares.

- The Hodrick and Prescott filter (HP Filter) extracts a stochastic trend that moves smoothly over time and is not correlated with the cycle component. The HP filter depends on a smoothing parameter (λ) that penalizes large fluctuations. A large λ implies a higher penalty and, therefore, a smoother cycle.

- The band pass filter is a frequency domain based filter. It assumes that the trend component has the power at lower frequencies of the spectrum. The choice in this procedure is to define the limits of the frequency band, say p_l and p_u, to isolate the cyclical component with a period of oscillation between p_l and p_u. We use an "optimal" finite sample approximation for the band pass filter as proposed by Christiano and Fitzgerald (2003). We make two choices for the cycle length between 1.5 and 8 years, $[p_l, p_u] = [6, 48]$, and between 1.5 and 6 years, $[p_l, p_u] = [6, 32]$, removing thus all the fluctuations that have a periodicity larger than 8(6) or smaller than 1.5 years.

- As regards the procedure to isolate the pure irregular component of the time series of interest, we follow the traditional approach of pre-whitening the series of interest, by means of ARIMA specifications, as in André and Pérez (2002; 2005). Let's assume a given cyclical component C_t is representable by a linear model of the general ARIMA class $\phi(B)C_t = \theta(B)\epsilon_t$ where ϵ_t is a white noise variable, and $\phi(B)$, $\theta(B)$ are polynomials in the lag operator B. Pre-multiplying C_t by an estimate of $\theta(B)^{-1}\phi(B)$ provides a pre-whitened version of C_t, which is an estimate for ϵ_t, a white noise variable representing the purely stochastic component of C_t. If the series y_t follows the above mentioned ARIMA process, the dynamic properties of the detrended series, call it y_t^F can be studied by means of expression $y_t^F = F(B)y_t = F(B)\frac{\theta(B)}{\phi(B)}\epsilon_t$ $= \Pi(B)\epsilon_t$, where F(B) is the filter applied to detrend the series. Thus, obtaining an estimate $\hat{\Pi}(B)$ of $\Pi(B)$ it is possible to generate the pre-whitened series $\hat{\epsilon}_t = \hat{\Pi}^{-1}(B)y_t^F$. If properly applied this method should be independent of the filtering procedure as correlations would be computed among irregular components.

D The SVAR approach

The reduced-form baseline VAR is specified in levels. $X_t \equiv (g_t, t_t, y_t, p_t, r_t)$ is the vector of endogenous variables and $D(L)$ is an autoregressive lag-polynomial. The benchmark specification includes a constant and a deterministic time trend. The vector $U_t \equiv (u_t^g, u_t^t, u_t^y, u_t^p, u_t^r)$ contains the reduced-form residuals, which in general will present non-zero cross-correlations. The baseline VAR includes four lags of each endogenous variable according to the information provided by LR tests, the Akaike information criterion and the final prediction error.

We apply the identification strategy proposed by Blanchard and Perotti (2002) and Perotti (2004), which exploits decision lags in policy making and information about the elasticity of fiscal variables to economic activity. Their strategy relies on the assumption that the reduced-form residuals of the g_t and t_t equations, u_t^g and u_t^t, can be thought of as linear combinations of three types of shocks: a) the automatic responses of spending and net taxes to the rest of macroeconomic variables in the system, b) systematic discretionary responses of fiscal policy to the same set of macro variables and c) random discretionary fiscal policy shocks, which are the truly uncorrelated structural fiscal policy shocks whose effects are the purpose of our analysis.

The innovations model can be written as $\Gamma U_t = BV_t$, where $V_t \equiv (e_t^g, e_t^t, e_t^y, e_t^p, e_t^r)$ is the vector containing the orthogonal structural shocks. Accordingly, the reduced-form residuals are linear combinations of the orthogonal structural shocks of the form $U_t = \Gamma^{-1} BV_t$. The respective matrices Γ and B can be written as:

$$\Gamma = \begin{pmatrix} 1 & 0 & -\alpha_{g,y} & -\alpha_{g,p} & -\alpha_{g,r} \\ 0 & 1 & -\alpha_{t,y} & -\alpha_{t,p} & -\alpha_{t,r} \\ -\gamma_{y,g} & -\gamma_{y,t} & 1 & 0 & 0 \\ -\gamma_{p,g} & -\gamma_{p,t} & -\gamma_{p,y} & 1 & 0 \\ -\gamma_{r,g} & -\gamma_{r,t} & -\gamma_{r,y} & -\gamma_{r,p} & 1 \end{pmatrix}$$

and

$$B = \begin{pmatrix} 1 & \beta_{g,t} & 0 & 0 & 0 \\ \beta_{t,g} & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix}$$

As we are interested in analysing the effects of "structural" discretionary spending shocks e_t^g on the rest of the variables of the system, estimations for the $\alpha_{i,j}$'s and $\beta_{i,j}$'s are needed. In general, approving and implementing new measures in response to specific economic circumstances typically takes longer than three months. Hence, one key assumption in this approach is that quarterly variables allow setting discretionary contemporaneous responses of fiscal variables to changes in underlying macroeconomic conditions to zero. Therefore, the coefficients $\alpha_{i,j}$'s only reflect the automatic responses of fiscal variables to the rest of the variables of the system, the first source of innovations aforementioned.

The way fiscal variables are defined allows making further assumptions concerning the values of the $\alpha_{i,j}$'s. Specifically, the semi-elasticities of fiscal variables to interest rate innovations are set to zero given that interest payments on government debt are excluded from both definitions. Moreover, the automatic responses of public expenditure to economic activity and the real exchange rate are also set to zero. The case of the price elasticity is different because some share of purchases of goods and services is likely to respond to the price level. Thus, we set the price elasticity of government expenditure to -0.5.

Output and price elasticities of net taxes, $\alpha_{t,y}$ and $\alpha_{t,p}$, are estimated at 0.64 and 0.87, respectively, fully in line with those in de Castro and Hernández de Cos (2008). These are obtained as weighted averages of the elasticities of the different net-tax components, including transfers, computed on the basis of information like statutory tax rates and estimations of the contemporaneous responses of the different tax-bases and, in the case of transfers, the relevant macroeconomic aggregate to GDP and price changes.

Furthermore, given that our main interest lies on expenditure shocks we assume that spending decisions are prior to tax ones, which implies a zero value for $\beta_{g,t}$. This allows us to retrieve e_t^g directly and use it to estimate $\beta_{t,g}$ by OLS, which completes the identification of the first two equations. For the remaining shocks the sequential ordering u_t^y, u_t^p and u_t^r is imposed. The corresponding structural shocks are estimated by instrumental variables in turn, using e_t^g and e_t^t as instruments for u_t^g and u_t^t, respectively. In any case, since we are interested in studying the effects of fiscal policy shocks, the ordering for the remaining variables is immaterial to the results. Impulse response functions are reported jointly with 68 % confidence bands obtained by Monte Carlo integration methods with 1000 replications.

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