COLLANA DEL
DIPARTIMENTO DI ECONOMIA

THE ITALIAN NOMINAL INTEREST RATE CONUNDRUM: A PROBLEM OF GROWTH OR PUBLIC FINANCE?

Nicola Caravaggio and Giovanni Carnazza

ISSN 2279-6916 Working papers

(Dipartimento di Economia Università degli studi Roma Tre) (online)

Working Paper n° 265, 2021

I Working Papers del Dipartimento di Economia svolgono la funzione di divulgare tempestivamente, in forma definitiva o provvisoria, i risultati di ricerche scientifiche originali. La loro pubblicazione è soggetta all'approvazione del Comitato Scientifico.

Per ciascuna pubblicazione vengono soddisfatti gli obblighi previsti dall'art. l del D.L.L. 31.8.1945, n. 660 e successive modifiche.

Copie della presente pubblicazione possono essere richieste alla Redazione.

REDAZIONE:
Dipartimento di Economia
Università degli Studi Roma Tre
Via Silvio D'Amico, 77 - 00145 Roma
Tel. 0039-06-57335655   fax 0039-06-57335771
E-mail: amm.economia@uniroma3.it
https://economia.uniroma3.it/

ROMA TRE
UNIVERSITÀ DEGLI STUDI

DIPARTIMENTO DI ECONOMIA

THE ITALIAN NOMINAL INTEREST RATE CONUNDRUM: A PROBLEM OF GROWTH OR PUBLIC FINANCE?

Nicola Caravaggio and Giovanni Carnazza

The Italian nominal interest rate conundrum: a problem of growth or public finance?[1]

Nicola Caravaggio (*)
Giovanni Carnazza (*)

Abstract

In the economic literature, there has been a large heterogeneity of results in relation to the impact of fiscal variables on interest rates. Focusing on the Italian economy and considering the nature of our interest rate determinants (public finance variables and nominal GDP growth), we decided to undertake a cointegration analysis relying on the Autoregressive Distributed Lag (ARDL) bound test approach, a particular suitable procedure within this peculiar framework, able to disentangle short-run and long-run dynamics. Our results are quite controversial, shedding new light on the role of gross debt and primary balance as a share of GDP in relation to the long-term Italian nominal interest rate. In this context, the ECB has probably played a crucial role, especially in the most severe phases of the Sovereign debt crisis. The European fiscal framework then shows further critical issues in relation to the new role that fiscal variables play within our econometric analysis.

Keywords: Italian economy; Sovereign bond yield; European Monetary Union; Public finance

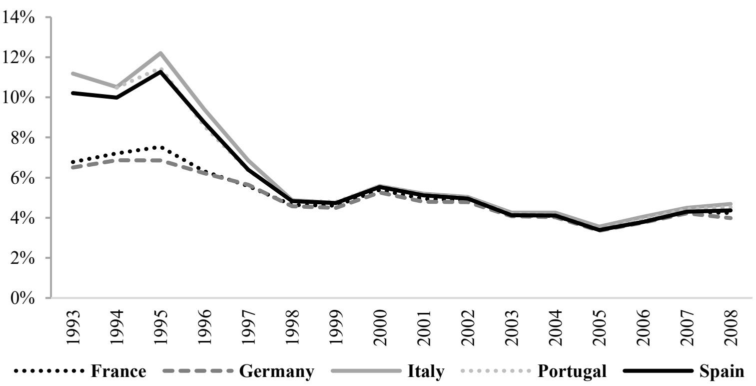
JEL Classification: E430; E580; E620; G120; C130; C220

[1] We thank Alessia Naccarato for having read the paper and for useful observations.
(*) Università degli Studi Roma Tre, Department of Economics – nicola.caravaggio@uniroma3.it
(*) Università degli Studi Roma Tre, Department of Economics – giovanni.carnazza@uniroma3.it

1. Introduction: setting the issue from an historical point of view

From a historical perspective, it is possibile to identify two different dynamics in the sovereign bond market of the countries of the European Monetary Union (EMU). In the years before and after the initial adoption of the single currency (1999), a rapid convergence of the interest rates within the euro-area sovereign bond market can be generally observed. At this regard, Figure 1 shows this convergence phenomenon in some selected European countries. This behavior has been considered fairly natural since there were no longer devaluation risks and the inflation was under the direct control of a single central bank (European Central Bank – ECB), whose main objective was only to maintain price stability. Losing each country the ability to set its inflation rate independently, the EMU formally eliminates inflation risk on government debt (Alesina *et al*., 1992). This shift is particularly relevant for those countries, such as Italy, which had experienced in the past high inflation and large public sector imbalances.

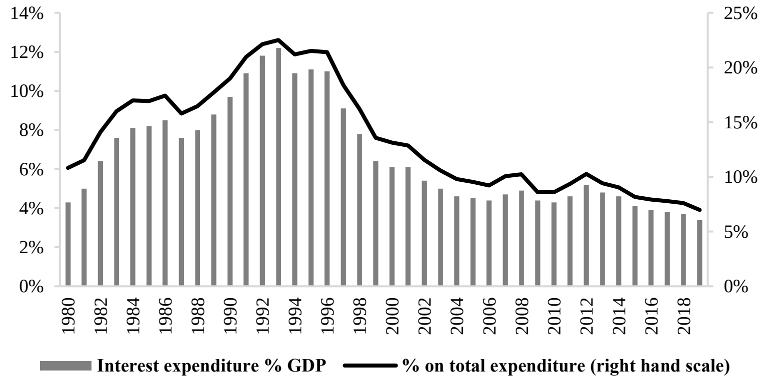
Figure 1 – 10-year nominal interest rates in selected European countries (1993-2008)



Source: authors' elaborations on *ECB* data

In relation to the period before the adoption of the single currency and focusing our attention on the Italian economy, the detachment path between fiscal and monetary policies can be actually traced back much earlier the offical institution of the ECB and it can be seen as a first attempt to deal with the high inflation rates mainly determined by the oil price shocks of the seventies. In particular, in 1981 the Bank of Italy gains full autonomy to decide wheter or not purchase Treasury bills not taken up by brokers at auctions (the so-called "divorce"). Some authors have interpreted this radical change, on the one side, as a way to give the Bank of Italy full control of the monetary base in order to moderate the sharp rise in prices – which can be seen as a legacy of the seventies – but, on the other side, also as a way to stimulate the governments of the eighties to implement a public debt stabilization policy. This second aspect soon turns out to be in vain, leading only to a change in the way the Treasury finance itself from the Bank of Italy to citizens' savings. In this context, the aforementioned divorce puts an end to the fiscal dominance regime which had characterized the previous period. The end of the accomodative policy of monetary financing of the public deficit determines a significant increase in the interest rate that the Treasury has to pay to the subscribers of newly issued public debt securities. This aspect, together with the increasingly large share of public debt held by market, determines an unprecedented increase in interest expenditure, which represents more than one fifth (22.5%) of the total public expenditure in 1993; in the same year, the share of interest expenditure on GDP reaches its historical maximum since the Italian unification of 1861 (12.2%) (Figure 2).[2]

Figure 2 – Interest expenditure as a % of GDP and of total expenditure in Italy (1980-2019)

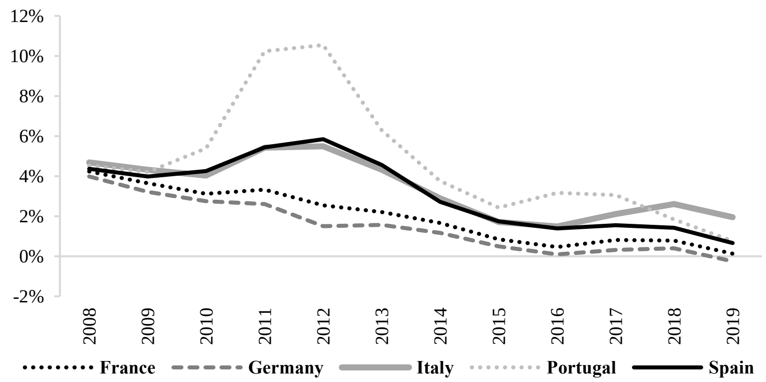


Source: authors' elaborations on *Bank of Italy* and *Istat* data

[2] In any case, we must also keep in mind that in the eighties the progressive liberalization of capital movements tends to eliminate the advantages enjoyed by the Treasury in accessing high national savings, exposing more and more each new issue to the judgment of financial markets.

In this scenario, since the beginning of the EMU, which includes the path convergence started – and imposed – after the Treaty of Maastricht, and until the first half of 2008, sovereign bond yields have been substantially the same with an average bond spread, relative to the benchmark German bund, close to zero. On the contrary, after the bankruptcy of Lehman Brothers in September 2008, it begins to emerge a significant divergence of the sovereign yields with debt markets experiencing a severe financial stress since mid-2010 (Figure 3). This represents the second different dynamic in the EMU sovereign bond market. Since then, nominal interest rates begins to deeply diverge, creating a new era in relation to the interrelationships among the ECB, the national sovereign debt markets, and the fiscal policies of each member state.

Figure 3 – 10-year nominal interest rates in some European countries (2008-2019)



Source: authors' elaborations on *ECB* data

In this new scenario, financial markets have become more sensitive to macroeconomic fundamentals after the 2008 crisis, especially after the Eurozone sovereign debt crisis and in the European periphery (Beirne and Fratzscher, 2013). In particular, over the past few years, the following aspects seem to have characterized the attitude of financial markets towards the Eurozone sovereign debt. At the outset of the crisis, investors have not attributed a significant importance to macroeconomic fundamentals, whose level has evolved very smoothly; this phase has instead been driven mainly by financial panic, concerning the most vulnerable countries in terms of public finance. The significant subsequent increase of the interest rates has then taken place against only marginally deteriorated economic conditions. If, at the peak of the crisis, the analysis of macroeconomic fundamentals had experienced a great attention, afterwards the sovereign spreads have returned to be driven by market sentiment (Alessi *et al.*, 2019). The new sensitivity of investors for the risk of sovereign default has increased financing costs of many euro member countries – for a certain period, Greece has even lost the access to the market. This significant change in risk awareness has boosted the interest rate costs associated with a deteriorating fiscal position; in particular, in the post crisis period the fiscal framework gains much more importance in relation to the determination of the level of the interest rates of sovereign bonds in the EMU (Heinemann *et al*., 2014). At this regard, it is interesting to note that the same European Commission methodology for estimating the fiscal room allowed for each country has *de facto* undermined the effective application of countercyclical fiscal policies in the most serious stages of the crisis, creating a vicious circle between restrictive fiscal policies and the fall of GDP (Fantacone *et al*., 2015; Carnazza *et al*., 2020). Most of the problems emerging after the financial crisis of 2008 were probably hidden in the same monetary union structure. In particular, Baldwin and Giavazzi (2015) identify three main critical issue: first of all, the lack of a central bank as a lender-of-last-resort; secondly, the prevalence of banking sector; the third factor is closely connected to the previous one, *i.e.*, the vicious circle between banks and governments. In this scenario, the absence of fiscal risk sharing and of a banking union make each member states of the EMU more vulnerable to sovereign debt crisis (Wyplosz and Sgherri, 2016).

Given these premises, our paper focuses its attention on the Italian nominal interest rate with a maturity of 10-year and on its determinants in the period between 1999 and 2019 considering quarterly data. We decide to not include 2020 in our sample since the health and the subsequent economic crisis seem to represent an anomaly in relation to the evolution of the sovereing interest rates and the debt-to-GDP ratio. This conclusion comes from two observations: first of all, the dynamics of sovereign interest rates at a national level have been strongly affected by a new asset purchase programme of private and public sector securities, called Pandemic Emergency Purchase Programme (PEPP), launched by the ECB on the evening of March 18, 2020 (Carnazza and Liberati, 2021); secondly, the significant fall of GDP have affected the debt-to-GDP ratio in a misleading way, taking into consideration that its growth path should soon return to normality during 2021 after the probable end of the health crisis. In this context, our main purpose will be to understand what are the main drivers of the nominal 10-year sovereign interest rate, especially in relation to public finance variables and the contribution of nominal GDP growth.

The paper is structured as follows. Section 2 discusses from a theoretical perspective the main determinants of sovereign bond yields with a specific focus on the role played by the primary balance and the debt-to-GDP ratio. Section 3 introduces the data and describes the methodology used to estimate the significance and the impact of selected variables on the 10-year Italian government bond interest rate. Section 4 shows and discusses the main results, introducing also some robustness checks. Section 5 concludes and provides some policy implications.

2. The main determinants of sovereign bond yields and the impact of fiscal policy: a brief review of the literature

Generally speaking, the main determinants of sovereign bond yields are three: credit risk, liquidity consideration, and changes in risk aversion. First of all, credit risk can be divided into three different types of risk: the default risk, defined as the possibility that the issuer does not repay either the coupon or the principal; the credit spread risk, interpreted as the danger that the interest rate on a bond turns out to be too exiguous relative to an investment with a lower default risk; the downgrade risk, which reflects the chance of a downgrade by a credit rating agency. Secondly, liquidity risk refers to the probability that a market is not characterized by a sufficient volume of buy and sell orders, as well as to the hazard that a large-scale transaction can strongly affect prices. Thirdly, risk aversion represents the propensity of investors to take risk, which – during times of financial uncertainty – usually translates into a ‘flight’ to the risk-free sovereign market. In the EU, this role is played by Germany, whose bonds are perceived as safe-haven assets (Barrios *et al.*, 2009). This general characterization has to be framed within the EMU, which changes the relative importance of these three determinants. In particular, the loss of monetary sovereignty and the impossibility of devaluing the national exchange rate expose each member state to more likely liquidity crisis and not just solvency crisis as in the case to be characterized by a national lender-of-last-resort. From a theoretical point of view, this is to say that the sovereign risk tends to be higher within a monetary union (Lemmen and Goodhart, 1999).

If we contextualize the topic of the main determinants of national interest rates in a global perspective, *i.e.*, outside the domestic market, here two different strands emerge: some authors take the complete integration of the capital markets as given, which implies the idea of the presence of a single pool of funds for OECD countries (Barro and Sala-i-Martin, 1990; Ford and Laxton, 1999); while some authors stress the key importance of national factors (Christiansen and Pigott, 1997; Breedon *et al*., 1999). In relation to the first aspect, it should be noted that the financial globalization and the following easier access to foreign savings has particularly helped emerging sovereign markets (Baldacci and Kumar, 2010). These aspects have also contributed to keep sovereign bond yields low in industrial countries, such as the United States (Hauner *et al*,, 2010). In this context, it is important to take into consideration a third aspect, *i.e.*, the possibility of potential spillovers among bond markets. This possibility comes from the observation that investors tend to react, particularly in the short-run, to rate movements in the most important markets rather than to domestic fundamentals. The phenomenon has been called by Summers (1986) ‘noise trading’ and it is mostly related to financial speculation based on the past movements of price changes rather than on the basis of economic fundamentals.

Our analysis follows the second strand of the literature as the first one tends to deny any impact of national factors, such as fiscal variables at the national level, on national interest rates.[3] At this regard, it is important to underline that the most recent theoretical and empirical literature stresses the importance of fiscal variables as the key drivers of sovereing bond yields, both on a long-run and a short-run basis (Gruber and Kamin, 2012; Lam and Tokuoka, 2013; Akram and Li, 2020). As highlighted by Afonso and Leal (2017), after the bankruptcy of Lehman Brothers, recent economic literature seems to confirm that macroeconomic determinants, such as the debt-to-GDP ratio and the budget balance, begin to explain much of the variation of sovereign bond yields in the long term (Laubach, 2009; Poghosyan, 2014). Since our main purpose is not to identify the main determinants of sovereign bond yields in the short-run or in the long-run but to measure the related role of fiscal policy, we decide to focus our attention on the effects of fiscal policy on interest rates. At this regard, theory does not offer a clear-cut answer. If we refer to the traditional IS-LM framework, an expansionary fiscal policy will not increase interest rates only in two extreme cases whose occurrence is substantially rare: first of all, when the economy is caught in the so-called liquidity trap, any movement of the IS curve does not influence the level of the interest rate; secondly, if the expansionary fiscal policy is determined by a tax cut and the marginal propensity to consume out of disposable income is zero, the consequent increase in the budget deficit will have no effect on interest rates. The previous reasoning implies that, within this conventional framework, the potential positive impact of fiscal policy on interest rate – *i.e.*, an increase of the interest rate – is generally true. In this traditional neoclassical context, this implies that the

[3] In any case, the same authors who underline the importance of global conditions in influencing government bond yields often recognize the important role national deficits, debts and other country specific factors may play (see, for example, Baldacci and Kumar, 2010).

higher deficits and debt levels the higher expected increase of the interest rate, with the size of this increasing mainly depending on the budget level, the economic environment and the impact on financial markets (Afonso and Leal, 2017). At this regard, Reinhart and Rogoff (2010) emphasize the damaging consequences of high levels of public debt on sovereign bonds due to potential distrust of investors in relation to any default risk.

In this context, Faini (2006) identifies three different aspects which can affect the potential impact of public debt and deficits on domestic interest rates: how Ricardian is the economy under consideration, the level of substitutability of privately versus publicly provided goods, and the degree of opennes of the economy. In relation to the first aspect, Barro (1974) suggested another reason why budget deficits arising from tax cuts may no have effect on the level of interest rates: individuals interpret a tax cut as equivalent to a postponed tax liability, which implies an unaltered intertemporal (and intergenerational) budget constraint and no change in their level of consumption. This is known in the literature as 'Ricardian equivalence,' according to an expression coined by Buchanan (1976), but it requires to be valid a number of questionable assumptions. First of all, the life horizon of individuals must be infinite (for this reason, it is important to take into consideration the concept of intergenerational legacy); secondly, taxes must not be distortionary (in other words, a government should introduce lump sum taxes, which represents the only case where the relative prices are not influenced); finally, individuals must be able to borrow and lend at the same rate the Treasury issues its debt. The second aspect highlighted by Faini (2006) hinges on the idea that if private and publicly provided goods are perfect substitutes, an increase in public spending will be fully offset by an equivalent fall in private consumption, leaving interest rates unchanged. Ultimately, traditionally in a small open economy, fiscal policy will be totally ineffective in influencing the level of interest rates: any change in domestic saving would be offset by international capital flows. However, it should be always took in mind that fiscal policy can always influence the first two determinants of sovereign bond yields, *i.e.*, the credit and the liquidity risk, increasing the risk premium demanded by financial markets.

Given the previous theoretical considerations, it is not surprising that there has been a large heterogeneity of results in relation to the impact of fiscal variables on interest rates. In particular, from an empirical perspective, in most of the literature, the potential impact of fiscal policy on interest rate is extremely linked to the way fiscal policy is measured. There exist two different channels which can affect the interest rate: on the one side, a flow variable, *i.e.*, the budget balance; on the other side, a stock variable, *i.e.*, the level of debt. Both variables are mostly considered in relation to GDP, making the previous definition more confusing: if in the first case the decifit/surplus to GDP ratio is comparing two flow variables, in the second case we are relating a stock variable at the numerator (the level of debt) to a flow variable at the denominator (the level of nominal GDP).

Concerning the first variable, *i.e.*, the budget balance, almost all major macroeconomic models underline the presence of an economically significant connection between changes in budget deficit and long-term interest rates (Gale and Orszag, 2003). This is especially true for the post crisis period (Afonso and Leal, 2017), but the conclusions of the literature are quite discordant, mostly when considering the second way fiscal policy can be measured, that is the debt-to-GDP ratio. Regarding the United States, for instance, Feldstein (1986) highlights that expected deficits, but not their current levels, have a clear impact on long-term interest rates while the same characterization seems not to be valid for public debt. Conversely, Goldstein and Woglom (1992) overturn the previous conclusions, finding evidence of a positive impact of the debt level of U.S. states on their bond yield. According to the two authors, there exists a 'credit punishing hypothesis': if a government runs persistent and excessive fiscal deficits, then it will face progressively an increased cost of borrowing. In other words, financial markets tend to automatically provide an incentive to correct irresponsible fiscal behaviours. Using data from 12 OECD countries, Alesina *et al*. (1992) show that the spread between public and private bond yields is positively associated to public debt-to-GDP ratio. Lemmen (1999) confirms the same outcomes in relation to yields of bond issued by state governments in Australia, Canada, and Germany. In relation to the EMU, Faini (2006) focuses his attention on the potential spillovers that a national fiscal policy may create to the rest of the Eurozone, demonstrating that fiscal policy, interpreted as budget balance-to-GDP ratio, tends to determine larger negative effect (*i.e*., a fall in the primary surplus or an increase in the primary deficit increase interest rates) at the Eurozone level than for individual countries. This implies the existence of significant spillovers between a national fiscal policy and the rest of the area. Taking as a benchmark the German macroeconomic fundamentals, Bernoth *et al.* (2004) find that EMU membership changes the relation between yield spreads and the fiscal variables significantly. In particular, before the adoption of a single currency, an increasing debt ratio relative to Germany amplifies the interest rate spread with an almost linear approximation. The foundation of the monetary union changes radically the previous characterization: on the one side, the linear effect of debt on interest rates tends to be lower; on the other side, a nonlinear relationship with growing marginal effects appears as the debt ratio relative to Germany increases. This new characterization implies that the risk premium is lower for those countries with a debt ratio not too distant from Germany's ratio and higher for the rest of the countries. According to the authors, the explanation is quite obvious: financial markets anticipate fiscal support for EMU countries in financial distress but not for those countries that have been characterized by a long history of fiscal undiscipline. In this context, Afonso and Leal (2017) examine the existence of a possible shift among the determinants of sovereign bond spreads of eleven EMU countries in relation to Germany before and after the formal beginning of the financial crisis of 2008, that is the bankruptcy of Lehman Brothers: the two fiscal variables of our interest, *i.e*., the budget balance and the public debt, become statistically significant only in the second period – a worsening of the budget balance or an increase in the public debt in relation to GDP tend to influence positevely the related spread – while in the first period the only determining factor of the current spread of a country is the spread of the previous year, which confirms the very low volatility during this latter period.

3. Data and methodology

Our analysis aims to investigate the Italian nominal 10-year sovereign interest rates (*INT*) by means of a set of core determinants: domestic and foreign gross debt as share of GDP (*GDD* and *GDF,* respectively), gross issue as percentage of GDP (*GI*), primary balance as percentage of GDP (*PB*), interests' expenditure as percentage of GDP (*IE*), and nominal GDP growth (*GDPG*). In relation to these variables, it appears important to preliminarily underline the reasons beyond their use. First of all, following Akram and Li's approach (2020), being only interested on the Italian sovereign market and not on the influence among the other Euro area member countries, we decide not to use the spread as a dependent variable but simply the absolute level of the nominal interest rate. Unlike other previously mentioned papers, we then divide the debt-to-GDP ratio between the domestic and the foreign component. Given the growing importance of such indicator, whose relevant influence has been previously highlighted both in historical and economic terms, our aim has been to investigate any eventual difference between the two components. We then consider gross issue in order to understand if the amount of issue may influence the evolution of the nominal interest rate.[4] As formerly noted, the primary balance can be then interpreted as the key indicator of the fiscal stance of a country, capturing better than total balance autonomous changes in fiscal policy (Ardagna *et al.*, 2007). Interest expenditure has been taken into consideration for completeness. Ultimately,

[4] At this regard, Carnazza (2019) shows that in Italy, on the one side, the issuance of new sovereign bond has tended to be more concentrated during phases of lower yield; in any case, on the other side, this shrewdness was not able to influence the trend of interest expenditures, mainly determined by the trend of the average return. The first result seems to be confirmed by our conclusions where an increase of gross issue leads to a decrease of the long-term nominal interest rate.

as highlighted by the previous review of the literature, we also consider the role of nominal GDP growth as a fundamental regressor of our dependent variable.

We used quarterly data from 1991 to 2019 for a total of 84 observations. Each series has been used in its raw form without seasonal or calendar adjustment. This decision has been influenced by the aim to better capture the high frequency linkage with our dependent variable and not to mislead the subsequent results.[5] Descriptive statistics and sources of our variables are reported in Table 1.

Table 1 – Descriptive statistics

	Source	Obs.	Mean	Std. Dev.	Min	Max	Variance	Skewness	Kurtosis
INT	*ECB*	84	3.912	1.312	1.213	6.613	1.722	-0.535	2.368
GDD	*Bank of Italy and Istat*	84	78.899	10.963	62.186	96.179	120.184	0.168	1.502
GDF	*Bank of Italy and Istat*	84	41.114	4.304	29.821	48.182	18.523	-0.404	2.512
GI	*ECB*	84	28.595	7.956	14.119	51.811	63.295	0.336	2.752
PB	*Istat*	84	1.621	4.252	-7.901	8.921	18.082	-0.618	2.332
IE	*Istat*	84	4.708	0.854	2.832	7.117	0.729	0.429	2.981
GDPG	*Istat*	84	0.785	7.369	-12.190	12.786	54.308	-0.329	1.780

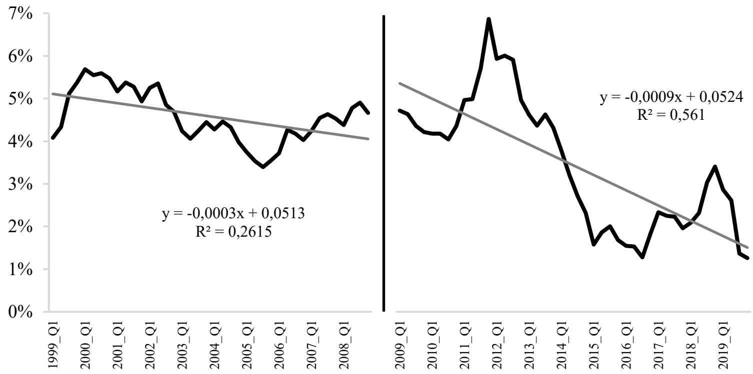
Source: authors' elaborations on *ECB, Bank of Italy* and *Istat* data

Before starting our main analysis, it may be useful to examine the dynamics of the dependent variable over the sample period. In order to show them in a more illustrative way, we decided to annualize the data with the exception of the nominal interest rate. The evolution of the latter variable has been divided taking the first quarter of 2009 as a watershed (Figure 4).[6] This decision is consistent with what has been done later in our model specification where a dummy variable has been set to the first quarter of 2009 in order to take into account the beginning of the Financial crisis and the following Sovereign debt European crisis.[7] The evolution of such variable is characterized by a general decreasing trend but the two subperiods show a different

[5] On the other hand, we are interested in identifying the immediate influences among the 10-year nominal interest rate and the others regressors. At this regard, a seasonal or calendar adjustment may hide significant information enclosed in the raw movements of each series.
[6] This structural break has been tested through a Chow test within a linear specification of the underlying trend, showing the highest level of significance.
[7] We tested this specific point with a Chow (1960) and a Wald (Quandt, 1960) test. Both confirmed the first quarter of 2009 as a structural break in the regression.

path: while during the first one the nominal interest rate displays a low variability, after the arrival of the Financial crisis in Europe its dynamics is much more fluctuating with a significant upsurge during the Sovereign debt crisis where high-public debt countries suffered most the financial stress. In this context, a crucial role has been played by the ECB which has provided wide support to the issuance of national debts: the tools progressively implemented have been successful in lowering the interest rates of the more exposed countries (Ghysels *et al.*, 2016; Krishnamurthy *et al.*, 2018) and in reducing the risk premium arising from liquidity concerns (De Pooter *et al.*, 2015).

Figure 4 – 10-year nominal interest rate in Italy



Source: authors' elaborations on *ECB* data

Before proceeding with the proper model specification, it is necessary to preliminary assess the unit root properties of the considered series by means of two different tests: the Augmented Dickey Fuller (ADF) (Dickey and Fuller, 1979, 1981) and the Phillips-Perron (PP) (Phillips and Perron, 1988) tests. Both tests have the null hypothesis H_0 "the series has a unit root" and results for our variables are reported in Table 2. The variable *INT* is clearly I (1) which means it must be integrated of order one to become stationary. Similarly, statistics from the two-unit root tests conclude for the presence of a unit root even for the variables *GDD* and *GDF*. Concerning the remaining variables, the two tests show mixed if not opposite results. The ADF test tends to identify as I (1) all remaining variables apart from *GDPG*, where the conclusion is for stationarity between 5% and 10% level of significance (corresponding to the test with and without the constant, respectively). With the PP test, instead, all remaining variables, especially *GDPG*, result to be stationary, or rather I (0). Therefore, we tested our variables with an additional unit root rest, the so-called KPSS test, proposed by Kwiatkowski *et al*. (1992) characterized by an opposite null hypothesis, that is "the series is stationary." Results are reported in Table 3 where the showed lag order is derived from the sample size following Schwert (1989).

Table 2 – Unit root tests (ADF and PP tests)

		Augmented Dickey-Fuller		**Phillips-Perron**	
		Statistic	p-value	Statistic	p-value
INT	Constant	-1.154	0.693	-0.858	0.801
	Constant + Trend	-2.22	0.479	-2.288	0.441
ΔINT	Constant	-4.068	0.001	-7.052	0
	Constant + Trend	-4.066	0.007	-7.106	0
GDD	Constant	-0.634	0.863	-0.583	0.875
	Constant + Trend	-2.426	0.366	-2.551	0.303
ΔGDD	Constant	-4.658	0	-6.557	0
	Constant + Trend	-4.717	0.001	-6.741	0
GDF	Constant	-2.366	0.152	-2.778	0.061
	Constant + Trend	-2.379	0.391	-2.692	0.240
ΔGDF	Constant	-5.001	0	-7.747	0
	Constant + Trend	-5.021	0	-7.774	0
GI	Constant	-1.873	0.345	-8.777	0
	Constant + Trend	-2.893	0.165	-9.995	0
ΔGI	Constant	-4.293	0.001	-20.803	0
	Constant + Trend	-4.279	0.003	-20.623	0
PB	Constant	-3.558	0.007	-14.886	0
	Constant + Trend	-4.571	0	-15.323	0
ΔPB	Constant	-3.284	0.069	-42.567	0
	Constant + Trend	-4.669	0.001	-42.524	0
IE	Constant	-0.853	0.803	-3.64	0.005
	Constant + Trend	-1.931	0.639	-6.716	0
ΔIE	Constant	-3.335	0.013	-27.278	0
	Constant + Trend	-3.284	0.069	-27.273	0
GDPG	Constant	-2.959	0.039	-37.434	0
	Constant + Trend	-4.957	0	-74.905	0
ΔGDPG	Constant	-3.332	0.061	-40.797	0
	Constant + Trend	-4.941	0	-74.408	0

Source: authors' elaborations on *ECB, Bank of Italy* and *Istat* data

Table 3 – Unit root tests (KPSS)

		Levels		First differences			Levels		First differences	
		Lag order	*Test statistic*	*Lag order*	*Test statistic*		*Lag order*	*Test statistic*	*Lag order*	*Test statistic*
INT	Constant	0	4.39***	0	0.209	Constant + Trend	0	0.729***	0	0.0682
		1	2.27***	1	0.17		1	0.385***	1	0.056
		2	1.56***	2	0.155		2	0.27***	2	0.0514
		3	1.21***	3	0.143		3	0.213**	3	0.0473
GDD	Constant	0	5.39***	0	0.681**	Constant + Trend	0	1.6***	0	0.243***
		1	2.73***	1	0.532**		1	0.828***	1	0.197**
		2	1.84***	2	0.487**		2	0.568***	2	0.187**
		3	1.4***	3	0.465**		3	0.438***	3	0.185**
GDF	Constant	0	4.21***	0	0.183	Constant + Trend	0	0.974***	0	0.0397
		1	2.22***	1	0.16		1	0.522***	1	0.035
		2	1.55***	2	0.162		2	0.371***	2	0.0359
		3	1.21***	3	0.161		3	0.295***	3	0.0364
GI	Constant	0	1.18***	0	0.0159	Constant + Trend	0	0.0678	0	0.012
		1	1,1***	1	0.0268		1	0.072	1	0.0203
		2	1.15***	2	0.0415		2	0.0888	2	0.0315
		3	1.19***	3	0.187		3	0.144	3	0.143
PB	Constant	0	0.0939	0	0.00552	Constant + Trend	0	0.0527	0	0.00552
		1	0.177	1	0.0208		1	0.101	1	0.0208
		2	0.175	2	0.0158		2	0.102	2	0.158
		3	0.328	3	0.172		3	0.204	3	0.172
IE	Constant	0	4.84***	0	0.0278	Constant + Trend	0	0.562***	0	0.0174
		1	2.83***	1	0.124		1	0.408***	1	0.0783
		2	1.96***	2	0.0701		2	0.176***	2	0.0442
		3	1.53***	3	0.217		3	0.216***	3	0.139*
GDPG	Constant	0	0.0111	0	0.00577	Constant + Trend	0	0.00837	0	0.00538
		1	0.0495	1	0.0489		1	0.0375	1	0.0456
		2	0.0302	2	0.0169		2	0.0229	2	0.0157
		3	0.229	3	0.272		3	0.182	3	0.253***

Note: *, **, *** indicate that statistics are significant at the 10%, 5%, and 1% level of significance, respectively
Source: authors' elaborations on *ECB, Bank of Italy* and *Istat* data

The variables *INT*, *GDP*, *GDF*, and *IE* clearly confirm the presence of unit root while *GI* results to be trend stationary. The variables *PG* and *GDPG*, instead, area clearly stationary. Eventually, we also performed the DF-GLS test, proposed by Elliott *et al.* (1996) which is a modified Dickey-Fuller test where the series has been transformed through a generalized least-squares (GLS) regression which makes this test greater in power compared with the ADF. In this final case, results confirm the presence of unit root for *INT*, *GDD*, and *GDF*. Furthermore, stationarity is confirmed for *GI*, *PB*, and *GDPG*, but not for *IE*, where results tend to confirm the presence of unit root, instead.[8] Accordingly, we identified *INT*, *GDD*, *GDF*, and *IE* as I (1) variables and *GI*, *PG*, and *GDPG* as I (0) variables.

The preliminary data analysis showed a mixed order of integration for our right-hand variables. Therefore, to investigate the presence of cointegration among them, we cannot rely on approaches such as Engle and Granger (1987), Johansen (1988, 1991, 1995), or Johansen and Juselius (1990), since they are constructed to assess cases where variables showed the same order of integration. Therefore, to test a long-run causal relationship between nominal interest rates (*INT*) and the group of variables introduced in the previous section, following the work of Akram and Li (2020), we implemented the Autoregressive Distributed Lag (ARDL) bound test procedure proposed by Pesaran *et al.* (1999) which is particularly suited when the variables of interest are either I (0) or I (1). The advantage of the bound testing procedure lies not only on the unnecessary equal order of integration among variables but also for the ability to deal with serial correlation and endogeneity with uneven lags order providing unbiased and efficient estimates. Following the two-step procedure of Pesaran *et al.* (1999), we first considered the following unrestricted Error Correction Model (ECM):

$$\begin{aligned} \Delta INT_t = \mu_0 &+ \beta_1 INT_{t-1} + \beta_2 GDD_{t-1} + \beta_3 GDF_{t-1} + \beta_4 GI_{t-1} + \beta_5 PB_{t-1} + \beta_6 IE_{t-1} + \beta_7 GDPG_{t-1} \\ &+ \sum_{i=1}^{p} \gamma_i \Delta INT_{t-i} \\ &+ \sum_{j=1}^{q_1} \delta_j \Delta GDD_{t-j} + \sum_{k=1}^{q_2} \zeta_k \Delta GDF_{t-k} + \sum_{l=1}^{q_3} \eta_l \Delta GI_{t-l} + \sum_{m=1}^{q_4} \lambda_m \Delta PB_{t-m} \\ &+ \sum_{r=1}^{q_5} \varphi_r \Delta IE_{t-r} + \sum_{s=1}^{q_6} \xi_s \Delta GDPG_{t-s} + \alpha DUM_t + \varepsilon_t \end{aligned} \tag{1}$$

where, in addition to the variables introduced in the previous section, μ_0 is the constant term, *DUM* is a dummy variable set to the first quarter of 2009 to account for the beginning of the

[8] Results from the DF-GLS tests are available upon request.

economic crisis's overspread in the Eurozone, while ε_t represents the idiosyncratic error. By estimating equation 1, we tested the null hypothesis of no cointegration among the variables (H_0: $\beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = \beta_6$) against the alternative of cointegration (H_1: $\beta_1 \neq \beta_2 \neq \beta_3 \neq \beta_4 \neq \beta_5 \neq \beta_6$). The F-statistic computed from the estimation are then evaluated to obtain the upper and lower critical values. If the result of the F-statistics falls outside the bounds, we can conclude for the presence of cointegration regardless of the variables' order of integration – as long as they are not greater than I (1) – and *vice versa*. Once established the cointegrating relationships among the considered variables, we estimate the long-run coefficient of our ARDL (*p*, *q1*, *q2*, *q3*, *q4*, *q5*, *q6*) model through equation (2) with restricted intercept and no trend. Furthermore, we estimated the short run dynamics through equation (3), where *ECT* is the error correction term, derived from the residuals of the long-run equation, which represents the speed of adjustment of the model to the equilibrium level after a shock. The parameter θ represents the speed of adjustment of the model which should results to be statistically significant with a range included between -1 and 0, meaning instantaneous equilibrium convergence and no convergence after a shock, respectively. The lag order of the model has been selected through the Akaike Information Criterion (AIC) selection assuming a White-type coefficient covariance matrix estimator which is consistent in the presence of both heteroskedasticity and autocorrelation (Newey and West, 1987).

$$\begin{aligned} INT_t = \mu_0 + \sum_{i=1}^{p} \gamma_i INT_{t-i} \\ + \sum_{j=1}^{q_1} \delta_j GDD_{t-j} + \sum_{k=1}^{q_2} \zeta_k GDF_{t-k} + \sum_{l=1}^{q_3} \eta_l GI_{t-l} + \sum_{m=1}^{q_4} \lambda_m PB_{t-m} \\ + \sum_{r=1}^{q_5} \varphi_r IE_{t-r} + \sum_{s=1}^{q_6} \xi_s GDPG_{t-s} + \alpha DUM_t + \varepsilon_t \end{aligned} \quad (2)$$

$$
\begin{aligned}
\Delta INT_t = \mu_0 &+ \sum_{i=1}^{p} \gamma_i \Delta INT_{t-i} \\
&+ \sum_{j=1}^{q_1} \delta_j \Delta GDD_{t-j} + \sum_{k=1}^{q_2} \zeta_k \Delta GDF_{t-k} + \sum_{l=1}^{q_3} \eta_l \Delta GI_{t-l} + \sum_{m=1}^{q_4} \lambda_m \Delta PB_{t-m} \\
&+ \sum_{r=1}^{q_5} \varphi_r \Delta IE_{t-r} + \sum_{s=1}^{q_6} \xi_s \Delta GDPG_{t-s} + \alpha DUM_t + \theta ECT_{t-1} + \varepsilon_t
\end{aligned} \tag{3}
$$

After the estimation of the ARDL, we checked its robustness by performing the following tests: the normal distribution of the residuals through the Jarque-Bera test (Jarque and Bera, 1987); the presence of serial correlation in the residuals through the Breusch-Godfrey Lagrange multiplier (LM) test (Breusch, 1978; Godfrey, 1978); the homoskedasticity in the residuals through the Breusch-Pagan-Godfrey test (Godfrey, 1978; Breusch and Pagan, 1979). Moreover, to assess possible omitted variable bias and functional misspecification of the model, we performed the Ramsey RESET test (1969). Eventually, following Pesaran and Shin (1998) we checked the stability of our results by means of the cumulative sum of recursive residuals (CUSUM) and the cumulative sum of squares of recursive residuals (CUSUMSQ) tests (Brown *et al*., 1975).

4. Results and robustness checks

The preliminary objective of our analysis is to test the actual existence of cointegration between the variables of interest. The selected model for the bound test procedure based on equation (1) is the following: ARDL (1, 4, 1, 1, 3, 0, 3). Results from the bound test procedure with the F-statistic and the t-statistic are reported in Table 4. The F-statistic falls above the upper critical value at 1% significance level, then rejecting the null hypothesis of the absence of a long run relationship. Once established, thorough the bound test, the existence of cointegration, we estimated out multivariate Vector Error Correction Model (VECM) through the previous equations (2) and (3). Results for the short-run and the long-run equations are reported in Table 5 and Table 6, respectively.

Table 4 – Bounds test of levels relationship

	k	F-statistic	Significance	I (0)	I (1)
INT	6	4.286	10%	1.99	2.94
			5%	2.27	3.28
			2.5%	2.55	3.61
			1%	2.88	3.99

Source: authors' elaborations on *ECB, Bank of Italy* and *Istat* data

Table 5 – Estimated short-run coefficients

Variable	Coefficient	Std. Error	t-Statistic	Prob.
ΔGDD	0.087	0.032	2.691	0.009
ΔGDD(-1)	0.054	0.022	2.415	0.019
ΔGDD(-2)	0.017	0.022	0.789	0.433
ΔGDD(-3)	-0.055	0.022	-2.473	0.016
ΔGDF	0.035	0.036	0.946	0.348
ΔGI	-0.044	0.010	-4.182	0.000
ΔGDPG	0.028	0.023	1.209	0.232
ΔGDPG(-1)	-0.195	0.035	-5.588	0
ΔGDPG(-2)	-0.059	0.020	-2.941	0.005
ΔPB	0.112	0.022	5.077	0
ΔPB(-1)	-0.096	0.026	-3.664	0.001
ΔPB(-2)	-0.078	0.021	-3.718	0.000
DUM	1.320	0.225	5.871	0
ECT(-1)	-0.253	0.041	-6.193	0

Source: authors' elaborations on *ECB, Bank of Italy* and *Istat* data

Table 6 – Estimated long-run coefficients

Variable	Coefficient	Std. Error	t-Statistic	Prob.
GDD	-0.238	0.052	-4.547	0
GDF	-0.315	0.098	-3.220	0.002
GI	-0.106	0.049	-2.147	0.036
GDPG	1.548	0.486	3.189	0.002
PB	0.620	0.170	3.648	0.001
IE	0.187	0.332	0.562	0.576
C	32.431	8.400	3.861	0

Source: authors' elaborations on *ECB, Bank of Italy* and *Istat* data

From the short run results, we can notice the *ECT* term and its significance at 1% level with an associated coefficient within the expected bound (from -1 to 0) and equal to -0.2528. This means that disequilibrium movements are corrected of about 25.28% from one period to another. Therefore, since we are dealing with quarterly data, a return to a long run equilibrium among the variables after a shock is achieved with one year. Still from the short run dynamics we can observe some different impact of our right-hand variables over *INT* compared to our long run results. *GDD* and *GI* have a general positive impact while *GDPG* and *PB* negative. From the short run estimation, we can observe also a positive impact associated with the international crisis dummy with a relative high coefficient. In fact, we clearly observed an increase in nominal interest rates after the spread in Europe of the financial crisis.

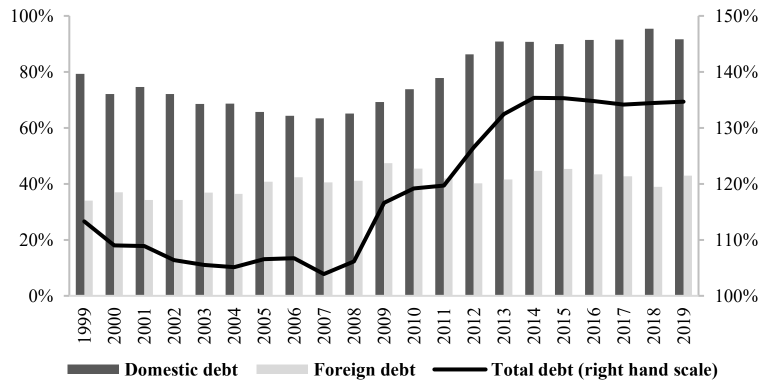
Long run estimations show a negative impact of both *GDD* and *GDF* at 1% level of significance: if *GDD* increases of 1 percentage points, *INT* will decrease by 0.24 while when *GDF* increases of the of the same amount, the negative impact on *INT* is of 0.31 percentage points. *GI* has a negative impact as well, although with a 5% level of significance and a lower coefficient: a 1 percentage point increase of *GI* will lead to a decrease of *INT* of about 0.11 percentage points. *GDPG*, instead, shows a particular high and positive coefficient with 1% level of significance: a 1 percentage point increase in *GDP* will lead to an increase of 1.55 percentage points of *INT*. *PB* as well has a positive and statistically significant coefficient: a 1 percentage point increase of *PG* will increase by 0.62 percentage points *INT*. Conversely, the long-run impact of *IE* is not statistically significant.

The previous results can be combined with three different figures (Figure 5, 6, and 7) in order to better explain the behaviour of the 10-year nominal interest rate. These figures describe the dynamics of the main and significant regressors over the sample period: domestic and foreign government gross debt (*GDD* and *GDF*) and primary public balance (*PB*) as a percentage of GDP and nominal GDP growth (*GDPG*).[9] First of all, it is interesting to analyse the overall evolution of the debt-to-GDP ratio (Figure 5). Its evolution clearly follows the underlying idea of the existence of two different subperiods: on the one side, if until 2008 this

[9] The following figures have been annualized in order to simplify their comprehension and their relationship with the dependent variable.

ratio has known a slight but constant decrease towards the value of 100%, on the other side, it begins to exponentially increase, reaching the maximum of 135.4% in 2014, and then stabilizing around this value. Focusing on this second subperiod and combining this information with the evolution of the 10-year interest rate (Figure 4), it seems evident that the most part of that increase begins when the sovereign bond yield faces a progressive and significant fall after having reached the maximum of 6.6%. If in the short-run *GDD* has a positive impact on the absolute amount of our dependent variable (Table 5), the long-run coefficients confirm how the latter variable has been negatively correlated with the debt-to-GDP ratio (Table 6). In this context, in relative terms *GDF* has grown more than *GDD*: comparing 1999 with 2019, their rates of growth have been respectively of 26.5% and 15.6%.

Figure 5 – Debt-to-GDP ratio: domestic *versus* foreign debt in Italy



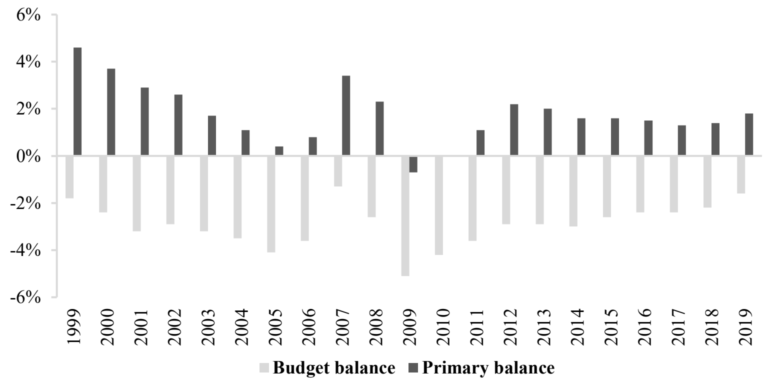
Source: authors' elaborations on *Bank of Italy* and *Istat* data

The primary balance (*PB*) represents the second important regressor of our analysis (Figure 6).[10] Also in this case, the short-run and the long-run effects of this variable are different. In the first case, the worsening of the primary public finance balances tends to increase the sovereign bond yield and this is coherent with the immediate impact of the increase of the debt-to-GDP ratio (Table 5). In the second case, the sign reverses, showing how an improvement of the primary balance translates into an upward push for the nominal interest rate. At this regard,

[10] If the primary balance raises, this translates into a decrease of the deficit or an increase of the surplus. The overall budget balance is reported in order to highlight the role of interest expenditure, which in any case has been found not significant in relation to our dependent variable.

it is interesting to underline that the second subperiod, which has been characterized by an overall reduction of the nominal interest rate after an initial upsurge (Figure 4), has always showed a positive primary balance. Interest expenditure has been found not significant and it is quite interesting to highlight that the overall deficit has been always driven by this kind of expenditure. Focusing only on the primary balance, within the European Union the Italian situation of an overall surplus represents an exceptional virtuous case together with few other Member States (Carnazza, 2018).

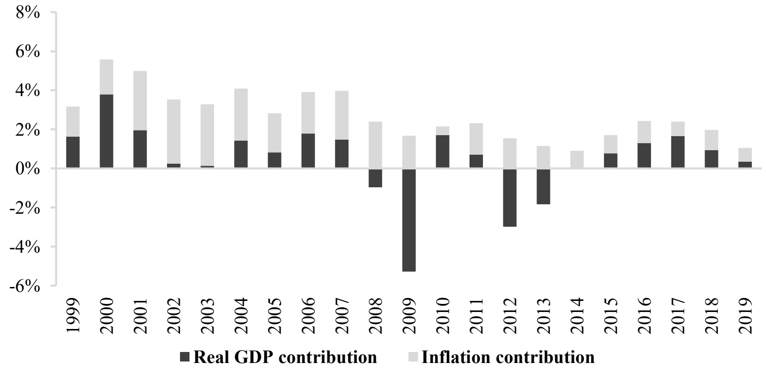
Figure 6 – Budget and primary balance in Italy



Source: authors' elaborations on *Istat* data

The last figure highlights the role of nominal GDP, dividing the related growth rate (*GDPG*) into its two different components: the real GDP contribution and the inflation contribution derived from GDP deflator (Figure 7).

Figure 7 – Nominal GDP growth and its contribution



Source: authors' elaborations on *Istat* data

This subdivision allows us to reflect on the distinctive role played by the two aforementioned factors. In this context, if in the short-run the nominal GDP growth tends to influence negatively the long-term sovereign bond yield, it is important to stress that this regressor represents one of the main significant variables in the long-run. In particular, there exists a significant positive relationship between the nominal GDP growth and the 10-year nominal interest rate. Representing the denominator of the other regressors, its impact is also influenced by the role of the other factors and, therefore, it is not surprising the strong positive impact on our dependent variable. This is particularly evident in the second subperiods when the nominal GDP growth knows an overall contraction mainly due to a fall of the real GDP. On the other side, the role of inflation in influencing the 10-year nominal interest rate is more faceted: if the dynamics of prices represents a crucial element in the remuneration required by financial operators, it is also true that the adoption of a single currency seems to have changed its relative importance. In particular, inflation tends to represent a not significant aspect in explaining sovereign bond yields and this conclusion may derive from the implementation of the single currency and the central bank independence, where price stability and the maintenance of interest rates became the responsibility of a supranational entity (Afonso and Leal, 2017).

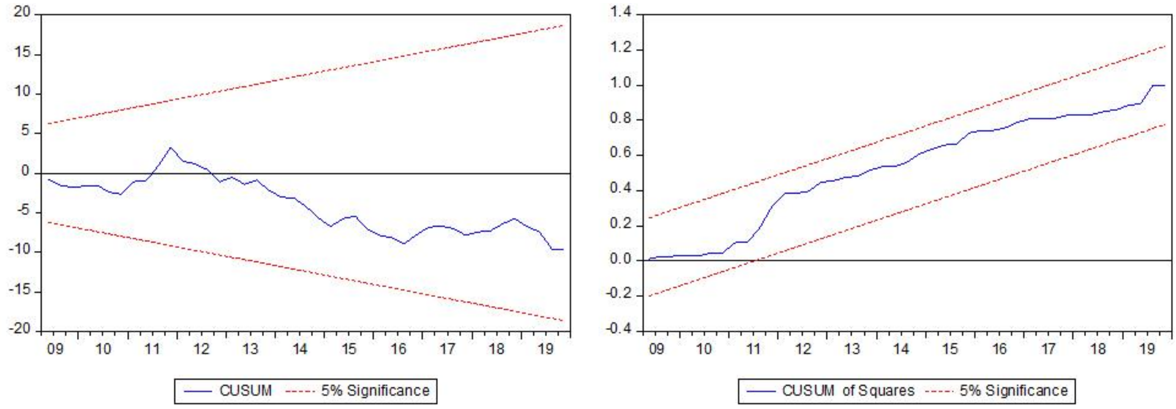
After the estimation of both short and long run dynamics and their implications, we conducted some diagnostic tests to evaluate the possible existence of flaws in our model such as serial correlation, heteroscedasticity, and functional misspecification. Results are reported in Table 7. The mode diagnostics for R^2 and adjusted R^2 are 96% and 95%, respectively, which indicate that the model is well fitted. The Jarque-Bera test concludes for a normal distribution of the residuals (we do not reject the null hypothesis of Normal distribution). They do not show presence of serial correlation and heteroscedasticity according with the Breusch-Godfrey LM test (we do not reject the null hypothesis of no autocorrelation) and the Breusch-Pagan-Godfrey test (we do not reject the null hypothesis of homoskedasticity), respectively. Moreover, the Ramsey RESET test shows the absence of functional misspecification meaning that non-linear combinations of explanatory variables do not have any power in explaining the response variable. Results from the CUSUM and CUSUMSQ are finally reported in Figure 8 and they fall between the 5% critical value bounds. This confirms the stability of our model showing no evidence of statistically significant breaks.

Table 7 – Diagnostic tests of the ARDL-VECM model

Test		Statistic	Prob.	Lags
R^2	0.9644			
Adjusted R^2	0.9523			
Normality: Jarque-Bera				
Jarque-Bera		0.8373	0.6579	
Serial Correlation: Breusch-Godfrey				
F-statistic		0.0900	0.7652	1
		0.3469	0.7084	2
		0.2450	0.8646	3
		0.8714	0.487	4
Heteroskedasticity: Breusch-Pagan-Godfrey				
F-statistic		1.2857	0.2249	
Functional misspecification: Ramsey RESET				
F-statistic*		1.7738	0.1881	

Note: the reported Ramsey RESET test has been performed by considering the squares of fitted values as omitted variables. Nonetheless, we also considered the cubic form obtaining the same conclusion.
Source: authors' elaborations on *ECB, Bank of Italy* and *Istat* data

Figure 8 – Coefficients' stability test



CUSUM CUSUMSQ

Note: The straight dotted lines in red represent the critical bounds at 5% significance level
Source: authors' elaborations on *ECB, Bank of Italy* and *Istat* data

5. Conclusions

The history of the EMU can be divided into two stages: on the one side, in the years before and after the initial adoption of the single currency it has been observed a rapid convergence of the interest rates; on the other side, this rapid convergence loses its importance, and a significant divergence of the sovereign yields begins to appear. Within this new scenario, financial markets have become more sensitive to macroeconomic fundamentals, especially in the European periphery countries, and the fiscal framework seems to represent an important aspect in relation to the determination of the level of the sovereign bond yields.

In the economic literature, there has been a large heterogeneity of results in relation to the impact of fiscal variables on interest rates. Focusing the attention on the Italian nominal interest rate with a maturity of 10-year as a dependent variable, our paper sheds new light on its underlying determinants. Basing our analysis on an ARDL model tailored on the Italian situation, we investigate the role and the impact of public finance variables and the role of nominal GDP growth. The subsequent results put in a different perspective the contribution of fiscal policies in the potential deterioration of financial markets' access. At this regard, we have found an interesting and in some way controversial relationship between public finance variables and the nominal interest rate: an increase of the debt-to-GDP ratio and a decrease of the primary surplus – or an expansion of the primary deficit balance – as a share of GDP tend to decrease the 10-year nominal interest rate in the long-run. From the short run estimation, we can observe a positive impact associated with the international crisis dummy with a relative high coefficient. In fact, we clearly observed an increase in nominal interest rates after the spread in Europe of the financial crisis. In any case, the long-run estimations confirm the previous underlying trends.

In this context, a crucial role has been probably played by the ECB which from 2012 has *de facto* committed itself to unlimited support of the government bond markets. The idea of the

ECB as a new lender of last resort is quite controversial but it is unanimous that the loss of a lender of last resort can make a monetary union very sensitive to changing market sentiments, creating a potential self-fulfilling prophecy: changing market sentiment may determine a sudden stop in government funding, setting in motion a dangerous vicious circle between liquidity and solvency crisis (De Grauwe, 2011). In light of this, it appears very important the ECB announcement in 2015 of the beginning of the Public Sector Purchase Programme (PSPP), which have made the ECB and the participating National Central Banks (NCBs) the dominant investors in the European sovereign bond market (Boermans and Keshkov, 2018).

Generally speaking, as seen, economic literature underlines the relevant effects on sovereign bond yields from unconventional monetary policy actions undertaken by the ECB. At this regard, our results should be contextualized under these general considerations, keeping in any case in mind that the structural change in the expansive stance of monetary policy has occurred in the second sub-sample of our analysis.

This new role of the ECB has been then also confirmed during the recent health crisis through the PEPP announcement, following the path of an important monetary support to relieve financial stress on bond markets (Carnazza and Liberati, 2021). The European fiscal framework, which has been recently suspended in order to cope with the outbreak of Covid-19, shows further critical issues in relation to the new role that fiscal variables play within our quantitative analysis.

References

Afonso, A. and Leal, F. S. (2017). Sovereign yield spreads in the EMU: crisis and structural determinants, Working Papers, 8, Lisbon School of Economics & Management.

Akram, T. and Li, H. (2020). An inquiry concerning long-term U.S. interest rates using monthly data, Applied Economics, 52(24), 2594-2621.

Alesina, A., De Broeck, M., Prati, A. and Tabellini, G. (1992). Default Risk on Government Debt in OECD Countries, Economic Policy, 7(15), 427-463.

Alessi, L., Balduzzi, P. and Savona, R. (2019). Anatomy of a Sovereign Debt Crisis: CDS Spreads and Real-Time Macroeconomic Data, JRC Working Papers in Economics and Finance.

Ardagna, S., Caselli, F. and Lane, T. (2007). Fiscal Discipline and the Cost of Public Debt Service: Some Estimates for OECD Countries, The B.E. Journal of Macroeconomics, 7(1), 1-35.

Baldacci, E. and Kumar, M.S. (2010). Fiscal deficits, public debt, and sovereign bond yields, IMF Working Paper 10/184.

Baldwin, R. and Giavazzi, F. (eds.) (2015). The Eurozone Crisis: A Consensus View of the Causes and a Few Possible Solutions, VoxEU.org Book, London: Centre for Economic Policy Research.

Barrios, S., Iversen, P., Lewandowska, M. and Setzer, R. (2009). Determinants of intra-euro-area government bond spreads during the Financial crisis. European Commission, Directorate General for Economic and Financial Affairs, Economic Papers, 388.

Barro, R. J. (1974). Are government bonds net wealth? Journal of Political Economy 82(6), 1095-1117.

Barro, R. J. and Sala-i-Martin, X. (1990). World real interest rates, NBER Macroeconomics Annual 1990 (Cambridge, MA: MIT Press), 15-59.

Beirne, J. and Fratzscher, M. (2013). The pricing of sovereign risk and contagion during the European debt crisis. ECB, Working Paper Series, 1625.

Bernoth, K., von Hagen, J. and Schuknecht, L. (2004). Sovereign risk premia in the European Government Bond Market. Working Paper Series, European Central Bank, 369.

Boermans, M. and Keshkov, V. (2018). The impact of the ECB asset purchases on the European bond market structure: Granular evidence on ownership concentration. DNB Working Paper, 590.

Breedon, F., Brian, H. and Geoffrey, W. (1999). Long-term Real Interest Rates: Evidence on the Global Capital Market, Oxford Review of Economic Policy 15(2), 128-142.

Breusch, T. S. (1978). Testing for Autocorrelation in Dynamic Linear Models. Australian Economic Papers, 17(31), 334–355.

Breusch, T. S. and Pagan, A. R. (1979). A Simple Test for Heteroskedasticity and Random Coefficient Variation. Econometrica, 47(5), 1287-1294.

Brown, R. L., Durbin, J. and Evans, J. M. (1975). Techniques for Testing the Constancy of Regression Relationships Over Time, Journal of the Royal Statistical Society, Series B (Methodological), 37(2), 149-163.

Buchanan, J. (1976). Barro on the Ricardian Equivalence Theorem. Journal of Political Economy, 84(2), 337-342.

Carnazza, G. (2018). Il saldo strutturale: origini, sviluppi e applicazioni nell'Unione Europea. Argomenti: Rivista di Economia, Cultura e Ricerca Sociale, 11, 15-39.

Carnazza, G. (2019). Spesa per interessi e ciclo economico: un'analisi empirica del caso italiano. Rivista di Diritto Finanziario e Scienza delle Finanze, LXXVIII, 53-75.

Carnazza, G., Liberati, P. and Sacchi, A. (2020) The cyclically-adjusted primary balance: A novel approach for the Euro area, Journal of Policy Modeling, 42(5), 1123-1145.

Carnazza, G. and Liberati, P. (2021). The asymmetric impact of the pandemic crisis on interest rates on public debt in the Eurozone, Journal of Policy Modeling, 43(3), 521-542.

Chow, G. C. (1960). Tests of Equality Between Sets of Coefficients in Two Linear Regressions. Econometrica, 28(3), 591-605.

Christiansen, H. and Pigott, C. (1997). Long-term Interest Rates in Globalised Markets, Economics Department Working Papers, 175, OECD.

De Grauwe, P. (2011). The Governance of a Fragile Eurozone. CEPS Working Documents, Economic Policy.

De Pooter, M., Martin., R. F. and Pruitt, S. (2015). The Liquidity Effects of Official Bond Market Intervention, International Finance Discussion Papers, 1138.

Dickey, D. A., and Fuller., W. A. (1979). Distribution of the Estimators for Autoregressive Time Series with a Unit Root, Journal of the American Statistical Association 74(366), 427-431.

Dickey, D. A., and Fuller, W. A. (1981). Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root, Econometrica 49(4), 1057-1072.

Engle, R. F., and Granger, C. W. J. (1987). Co-Integration and Error Correction: Representation, Estimation, and Testing, Econometrica 55(2), 251-276.

Elliott, G., Rothenberg, T. J., and Stock, J. H. (1996). Efficient tests for an autoregressive unit root, Econometrica, 64(4), 813-836.

Faini, R. (2006). Fiscal policy and interest rates in Europe. Economic Policy, 21(47), 443-489.

Fantacone, S., Garalova, P. G. e Milani, C. (2015). Structural Budget Balance and Fiscal Policy: the Limits of the European Approach. YILDIZ Social Science Review, 1(2), 19-34.

Feldstein, M. (1986). Budget Deficits, Tax rules, and Real Interest Rates, National Bureau of Economic Research, Working Paper, 1970.

Ford, R. and Laxton, D. (1999). World Public Debt and Real Interest Rates, Review of Economic Policy 15(2), 77-94.

Gale, W. G. and Orszag, P. (2003). The Economic Effects of Long-Term Fiscal Discipline, Tax Policy Center, Discussion Paper, 8.

Ghysels, E., Idier, J., Manganelli, S. and Vergote, O. (2016). A High-Frequency Assessment of the ECB Securities Markets Programme, Journal of the European Economic Association, 15(1), 218-243.

Godfrey, L. G. (1978). Testing against general autoregressive and moving average error models when the regressors include lagged dependent variables, Econometrica: Journal of the Econometric Society, 1293-1301.

Goldstein, M., Woglom, G. (1992). Market Based Fiscal Discipline in Monetary Unions: Evidence from the US Municipal Bond Market, in M. Canzoneri, V. Grilli, and P. Masson (eds.) Establishing a Central Bank, Cambridge University Press, 1992.

Gruber, J. W. and Kamin, S. B. (2012). Fiscal Positions and Government Bond Yields in OECD Countries, Journal of Money, Credit and Banking, 44(8), 1563-1587.

Hauner, D., Jonas, J. and Kumar, M. S. (2010). Sovereign Risk: Are the EU's New Member States Different?, Oxford Bulletin of Economics and Statistics, 72(4), 411-427.

Heinemann, F., Osterloh, S. and Kalb, A. (2014). Sovereign risk premia: The link between fiscal rules and stability culture, Journal of International Money and Finance, 41, 110-127.

Jarque, C. M. and Bera, A. K. (1987). A test for normality of observations and regression residuals, International Statistical Review, 55(2), 163-172.

Johansen, S. (1988). Statistical Analysis of Cointegration Vectors, Journal of Economic Dynamics and Control 12(2), 231-254.

Johansen, S. (1991). Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models, Econometrica, 59(6), 1551-1580.

Johansen, S. (1995). Likelihood-Based Inference in Cointegrated Vector Autoregressive Models. Oxford, UK: Clarendon Press.

Johansen, S. and K. Juselius (1990). Maximum Likelihood Estimation and Inference on Cointegration with Applications to the Demand for Money, Oxford Bulletin of Economics and Statistics, 52(2), 169-210.

Krishnamurthy, A., Nagel, S. and Vissing-Jorgensen, A. (2018). ECB Policies Involving Government Bond Purchases: Impact and Channels, Review of Finance, 22(1), 1-44.

Kwiatkowski, D., Phillips, P. C. B, Schmidt, P. and Y. Shin (1992). Testing the null hypothesis of stationarity against the alternative of a unit root: How sure are we that economic time series have a unit root?, Journal of Econometrics, 54(1-3), 159-178.

Lam, R. W. and Tokuoka, K. (2013). Assessing the Risks to the Japanese Government Bond Market, Journal of International Commerce, Economics and Policy, 4(1), 1-15.

Laubach, T. (2009). New evidence on the interest rate effects of budget deficits and debt. Journal of the European Economic Association 7(4), 858-885.

Lemmen, J. (1999). Managing Government Default Risk in Federal States, FMG Special Paper, 116.

Lemmen, J. and Goodhart, C. (1999). Credit risks and Government Bond Markets: A Panel Data Econometric Analysis, Eastern Economic Journal, 25(1), 77-107.

Newey, W. K., & West, K. D. (1987). Hypothesis testing with efficient method of moments estimation. International Economic Review, 777-787.

Pesaran, H. H., and Y. Shin (1998). Generalized Impulse Response Analysis in Linear Multivariate Models, Economics Letters 58(1), 17-29.

Pesaran, M. H., Y. Shin, and R. J. Smith (1999) Bounds Testing Approaches to the Analysis of Long Run Relationships, Edinburgh School of Economics, Discussion Paper Series, 46, Edinburgh: Edinburgh School of Economics.

Phillips, P. C. B., and P. Perron. (1988). Testing for a Unit Root in Time Series Regression, Biometrika, 75(2), 335-346.

Poghosyan, T. (2014). Long-run and short-run determinants of sovereign bond yields in advanced economies, Economic Systems 38(1), 100-114.

Quandt, R. E. (1960). Tests of the hypothesis that a linear regression system obeys two separate regimes. Journal of the American Statistical Association 55(290), 324-330.

Ramsey, J. B. (1969). Tests for Specification Errors in Classical Linear Least Squares Regression Analysis, Journal of the Royal Statistical Society, Series B., 31(2), 350-371.

Reinhart, C. M. and Rogoff, K. S. (2010). Growth in a time of debt, American Economic Review 100(2), 573-578.

Schwert, G. W. 1989. Tests for unit roots: A Monte Carlo investigation. Journal of Business and Economic Statistics 7, 147-160.

Summers, L. H. (1986). Does the Stock Market Rationally Reflect Fundamentals Values? Journal of Finance, 41(3), 591-601.

Wyplosz, C. and Sgherri, S. (2016). The IMF’s Role in Greece in the Context of the 2010 Stand-By Arrangement, Independent Evaluation Office of the International Monetary Fund.