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"Debt-growth linkages in EMU across countries and time horizons"

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Abstract

This paper contributes to the literature by empirically examining whether the influence of public debt on economic growth differs between the short and the long run and presents different patterns across euro-area countries. To this end, we use annual data from both central and peripheral countries of the European Economic and Monetary Union (EMU) for the 1960-2012 period and estimate a growth model augmented for public debt using the Autoregressive Distributed Lag (ARDL) bounds testing approach. Our findings tend to support the view that public debt always has a negative impact on the long-run performance of EMU countries, whilst its short-run effect may be positive depending on the country.

JEL classification: C22, F33, H63, O40, O52

Keywords: Public debt, economic growth, bounds testing, euro area, peripheral EMU countries, central EMU countries.

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1. Introduction

The origin of the sovereign debt crisis in the European Economic and Monetary Union (EMU) goes deeper than the fiscal imbalances in euro area countries. The interconnection between banking, sovereign, and economic crises is obvious: the problems of weak banks and high sovereign debt were mutually reinforcing, and both were exacerbated by weak, constrained growth. Some authors (see Shambaugh, 2012) pointed out that these three interlocking crises (banking, sovereign debt, and economic growth) came together to challenge the viability of the currency union. An analysis of the interrelationship between sovereign and banking risk, an issue of great importance since the development of a "diabolic loop" between sovereigns and banks [see, for example, Alter and Schüler (2012), De Bruyckere *et al.* (2013), Alter and Beyer (2014) or Singh *et al.*, (2016)] is beyond the scope of this paper. Rather, we will focus on the interconnection between sovereign debt and economic growth in 11 EMU countries, both central (Austria, Belgium, Finland, France, Germany and the Netherlands) and peripheral (Greece, Ireland, Italy, Portugal, and Spain)[1]. The recent crisis led to an unprecedented increase in sovereign debt across euro-area countries (by the end of 2013, on average, public debt reached about 100% of GDP – its highest level in 50 years), raising serious concerns about its impact on economic growth.

Overall, the theoretical literature finds that there is cause for taking into account the effects of very high debt on the capital stock, growth, and risk since it tends to point to a negative link between the public debt-to-GDP ratio and the steady-state growth rate of GDP (see, for instance, Aizenman *et al.*, 2007). However, the conventional view is that the impact of debt on output differs depending on the time horizon. While debt may crowd out capital and reduce output in the long run (Salotti and Trecroci, 2016), in the short run it can

[1] Fölster and Henrekson (1999), Romero-Ávila and Strauch (2008), Afonso and Furceri (2010) and Jetter (2014), among others, examine the effects of public finances on economic growth.

stimulate aggregate demand and output [see Barro (1990) or Elmendorf and Mankiw (1999)].

Moreover, some recent studies support the idea that the presence of a tipping point (above which an increase in public debt has a detrimental effect on economic performance) does not mean that it has to be common across countries [see Ghosh *et al.* (2013), Eberhardt and Presbitero (2015) or Markus and Rainer (2016) among them]. Eberhardt and Presbitero (2015) stress that there may be many reasons for the differences in the relationships between public debt and growth across countries. First, production technology may differ across countries, and thus also the relationship between debt and growth. Second, the capacity to tolerate high levels of debt may depend on a number of country-specific characteristics, related to past crises and the macro and institutional framework. Third, vulnerability to public debt may depend not only on debt levels, but also on debt composition (domestic versus external, foreign or domestic currency denominated or long-term versus short term), which may also differ significantly across countries.

However there is hardly any empirical analysis in the literature of the potential heterogeneity in the debt-growth nexus both across EMU countries and across time periods. Indeed, while there is a substantial body of research exploring the interconnection between debt and growth in both developed and emerging countries, few papers to date have looked at this link in the context of EMU. The exceptions, which include Checherita-Westphal and Rother (2012), Baum *et al.* (2013), Dreger and Reimers (2013) and Antonakakis (2014), make use of panel data techniques and obtain average results for euro area countries but do not distinguish between short and long run effects.

In this context, this paper presents a new approach to add to the as yet small body of literature on the relationship between debt accumulation and economic performance in

EMU countries, by examining the potential heterogeneity in the debt-growth nexus both across different euro-area countries and across different time horizons. Therefore, this paper's contribution to the empirical literature is twofold. First, unlike previous studies, we do not make use of panel estimation techniques to combine the power of cross section averaging with all the subtleties of temporal dependence; rather, we explore the time series dimension of the issue to obtain further evidence based on the historical experience of each country in the sample in order to detect potential heterogeneities in the relationship across euro area countries. Second, our econometric methodology is data-driven, and it allows us to select the statistical model that best approximates the relationship between the variables under study for any particular country and to assess both short and long-run effects of public debt on output performance.

The rest of the paper is organized as follows. Section 2 justifies our empirical approach on the basis of a review of the existing literature. Section 3 presents the theoretical framework of the analysis and outlines the econometric methodology. Section 4 describes our data and presents our empirical results. Finally, Section 5 summarizes the findings and offers some concluding remarks.

2. Literature review

Under what conditions is debt growth-enhancing? This challenging question has been studied by economists for a long time, but has recently undergone a notable revival fuelled by the substantial deterioration of public finances in many economies as a result of the financial and economic crisis of 2008-2009[2]. However, the results from the empirical literature on the relationship between public debt and economic growth are far from conclusive [see Panizza and Presbitero (2013) or the technical Appendix in Eberhardt and

[2] During the crisis, public deficits increased not only because economic automatic stabilizers began to work (which meant, for instance, declining revenues) but also because of the launch of fiscal stimulus packages.

Presbitero for two excellent summaries of this literature]. Some authors (Reinhart and Rogoff, 2010 or Pattillo *et al.*, 2011) present empirical evidence to indicate that the relationship is described by an inverted U-shaped pattern (whilst low levels of public debt positively affect economic growth, high levels have a negative impact). In particular, in their seminal paper using a database of 44 countries over a time period spanning 200 years, Reinhart and Rogoff (2010) suggest that the relationship is weak for public debt ratios below 90% of GDP, but that, on average, growth rates decrease substantially above this threshold. However, using data on 20 developed countries Lof and Malinen (2014) find no evidence for a robust effect of debt on growth, even for higher levels of debt; whereas Woo and Kumar (2015), controlling for other factors that also influence growth, detected an inverse relationship between the two variables.

In the EMU context, in a situation in which leverage was already very high[3], the recent economic recession and sovereign debt crisis has stimulated an intense debate both on the effectiveness of fiscal policies and on the possible adverse consequences of the accumulation of public debt in euro area countries. The debate is hotly contested, not only because pundits draw widely different conclusions for macroeconomic policy (in particular, in relation to their positions on economic austerity policies), but also because economists have not reached a consensus (see Alesina *et al.* 2015). Some suggest that now is precisely the time to apply the lessons learnt during the Great Depression and that policymakers should implement expansionary fiscal policies [see, among others, Krugman (2011), Berg and Ostry (2011) or DeLong and Summers (2012)][4] since fiscal austerity may have been the main culprit for the recessions experienced by European countries; others argue that, since

[3] In this regard, Gómez-Puig (2013) attempts to quantify the total level of indebtedness (public and private) in all euro area countries, using a database created with the statistics provided by the European Central Bank. According to her calculations, in September 2012, total leverage (public and private) over GDP recorded levels of 710%, 487%, 413%, 360% and 353% in Ireland, Portugal, Spain, Italy and Greece respectively.

[4] These authors state that deleveraging policies may even prove to be detrimental, depending on the fundamental variables of the economy. Their argument is currently supported by some politicians in southern Europe.

the high level of public sector leverage has a negative effect on economic growth, fiscal consolidation is fundamental to restoring confidence and improving expectations about the future evolution of the economy and therefore its rate of growth [see Cochrane (2011) or Teles and Mussolini (2014)][5].

In our reading of the empirical evidence, few papers have examined the relationship between debt and growth for euro area countries despite the severe sovereign debt crisis in the EMU. Checherita-Westphal and Rother (2012) analyse the empirics of the debt-growth nexus using a standard growth model. They find that the turning point (beyond which government debt negatively affects growth) is at 90–100% of GDP. Baum *et al.* (2013) detect a similar threshold by employing a dynamic approach (while the short-run impact of debt on per capita GDP growth is positive and significant, it decreases to zero beyond debt-to-GDP ratios of 67%, and for ratios above 95% additional debt has a negative impact on output growth). In contrast, Dreger and Reimers (2013) base their analysis on the distinction between sustainable and non-sustainable debt periods and find that the negative impact of the debt-to-GDP ratio on growth in the euro area is limited to periods of non-sustainable public debt; instead, debt will exert a positive impact on growth given that it is sustainable. The studies by Checherita-Westphal and Rother (2012), Baum *et al.* (2013) and Dreger and Reimers (2013) are unified and extended by Antonakakis (2014). Like them, he uses a panel approach, but in addition to debt non-linearities, he also examines the effect of debt sustainability on economic growth in the euro area. Overall, the above-mentioned empirical literature lends support to the presence of a common debt threshold across (similar) countries, like those in the euro-area, and does not distinguish between short- and-long run effects.

[5] The latter approach, which supports austerity measures, has been highly influential among the EMU authorities and is supported by the empirical evidence presented in some influential papers (Reinhart and Rogoff, 2010, among them).

Therefore, to our knowledge, no strong case has yet been made for analysing the effect of debt accumulation on economic growth taking into account the particular characteristics of each euro area economy and examining whether the effects differ depending on the time horizon, in spite of the fact that this potential heterogeneity has been stressed by the literature.

In this context, Eberhardt and Presbitero (2015), who investigate the debt-growth relationship in 118 developing, emerging and advanced economies, find some evidence for nonlinearity and state that there is no evidence at all for a threshold level common to all countries over time. Égert (2015) presents empirical evidence suggesting that 90% is not a magic number since the threshold can be lower and the nonlinearity can change across different samples and specifications. Finally, examining the bi-directional causality between debt and growth in a sample of eleven EMU countries, Gómez-Puig and Sosvilla-Rivero (2015) find that public debt has a negative effect on growth from an endogenously determined breakpoint and above a debt threshold ranging from 56% to 103% depending on the country.

Nor is there any consensus in the literature regarding the sign of the impact of public debt on output in either the short or the long run. The "conventional" view (Elmendorf and Mankiw, 1999) states that in the short-run, since output is demand-determined, government debt (manifesting deficit financing) can have a positive effect on disposable income, aggregate demand, and overall output. Moderate levels of debt are found to have a positive impact on economic growth through a range of channels: improved monetary policy, strengthened institutions, enhanced private savings, and deepened financial intermediation (Abbas and Christensen, 2007). Government debt could be used to smooth distortionary taxation over time (Barro, 1979). This positive short-run effect of budget deficits (and higher debt) is likely to be large when the output is far from capacity.

However, things are different in the long run if the decrease in public savings brought about by a higher budget deficit is not fully compensated by an increase in private savings. In this situation, national savings will decrease and total investment will fall; this will have a negative effect on GDP as it will reduce capital stock, increase interest rates, and reduce labour productivity and wages. The negative effect of an increase in public debt on future GDP can be amplified if high public debt increases uncertainty or leads to expectations of future confiscation, possibly through inflation and financial repression (see Cochrane, 2011). Nonetheless, this "conventional" split between the short and long-run effects of debt disregards the fact that protracted recessions may reduce future potential output (as they increase the number of discouraged workers, with the associated loss of skills, and have a negative effect on organizational capital and investment in new activities). There is, in fact, evidence that recessions have a permanent effect on the level of future GDP (see, e.g., Cerra and Saxena, 2008) which implies that running fiscal deficits (and increasing debt) may have a positive effect on output in both the short and the long run. DeLong and Summers (2012) argue that, in a low interest rate environment, an expansionary fiscal policy is likely to be self-financing. Another strand of literature departs from this "conventional" view and establishes a link between the long-term effect of debt and the kind of public expenditure it funds. The papers by Devarajan *et al.* (1996) and Aschauer (1989), for instance, state that in the long run, the impact of debt on the economy's performance depends on whether the public expenditure funded by government debt is productive or unproductive. Whilst the former (which includes physical infrastructure such as roads and railways, communication, information systems such as phone, internet, and education)[6] may have a positive impact on the economy's growth, the latter does not affect the economy's long-run performance, although it may have positive short-run implications.

[6] Although this sort of investment might not be profitable from the point of view of the single firm (as private costs

3. Theoretical framework and econometric methodology

Our empirical exploration is based upon a standard growth model augmented by adding a debt variable. This allows us to test the impact of debt after controlling for the basic drivers of growth: the stock of physical capital, the labour input and a measure of human capital. The stock of physical capital and the labour input have been the two key determinants of economic growth since Solow's classic model (1956) and many empirical studies have examined their relationship with economic growth (see, e. g. Frankel, 1962). Regarding human capital, Becker (1962) stated that investment in human capital contributed to economic growth by investing in people through education and health, and Mankiw *et al.* (1992) augmented the Solow model by including accumulation of human as well as physical capital (see Savvide and Stengos, 2009).

Therefore, the following equation will be the basis of our empirical analysis:

$$y_t = \alpha_0 + \alpha_1 k_t + \alpha_2 l_t + \alpha_3 h_t + \alpha_4 d_t \qquad (1)$$

As can be seen, equation (1) postulates a long-run relationship between (the log of) the level of production (y_t), (the log of) the stock of physical capital (k_t), (the log of) the labour employed (l_t), (the log of) the human capital (h_t) and (the log of) the stock of public debt (d_t). This relationship can be estimated from sufficiently long time series by cointegration econometric techniques. In this paper we make use of the Autoregressive Distributed Lag (ARDL) bounds testing approach to cointegration proposed by Pesaran and Shin (1999) and Pesaran, Shin and Smith (2001). The ARDL approach involves estimating the conditional error correction version of the ARDL model for the variables under estimation. The existence of an error-correction term among a number of cointegrated variables

exceed private returns), the whole economy would nevertheless benefit enormously, which justifies public provision. For instance, Glomm and Ravikumar (1997), among others, contend that both government infrastructure investment and education expenditures have a significant impact on an economy's long-term growth rate.

implies that changes in the dependent variable are a function of both the level of disequilibrium in the cointegration relationship and the changes in other explanatory variables. This tells us whether any deviation from the long-run equilibrium is feed-backed on the changes in the dependent variable in order to force the movement towards the long-run equilibrium.

This approach presents at least three significant advantages over the two alternatives commonly used in the empirical literature: the single-equation procedure developed by Engle and Granger (1987) and the maximum likelihood method postulated by Johansen (1991, 1995) which is based on a system of equations. First, both these approaches require the variables under study to be integrated of order 1; this inevitably requires a previous process of tests on the order of integration of the series, which may lead to some uncertainty in the analysis of long-run relations. In contrast, the ARDL bounds testing approach allows the analysis of long-term relationships between variables, regardless of whether they are integrated of order 0 [*I(0)*], of order 1 [*I(1)*] or mutually cointegrated. This avoids some of the common pitfalls faced in the empirical analysis of time series, such as the lack of power of unit root tests and doubts about the order of integration of the variables examined. Second, the ARDL bounds testing approach allows a distinction to be made between the dependent variable and the explanatory variables, an obvious advantage over the method proposed by Engle and Granger; at the same time, like the Johansen approach, it allows simultaneous estimation of the short-run and long-run components, eliminating the problems associated with omitted variables and the presence of autocorrelation. Finally, while the estimation results obtained by the methods proposed by Engle and Granger and Johansen are not robust to small samples, Pesaran and Shin (1999) show that the short-run parameters estimated using their approach are $\sqrt{T}-$consistent and the long-run parameters are super-consistent in small samples.

In our particular case, the application of the ARDL approach to cointegration involves estimating the following unrestricted error correction model (UECM):

$$\begin{aligned}\Delta y_t = \beta + \sum_{i=1}^{p} \gamma_i \Delta y_{t-i} + \sum_{i=1}^{p} \omega_i \Delta k_{t-i} + \sum_{i=1}^{p} \varphi_i \Delta l_{t-i} + \sum_{i=1}^{p} \upsilon_i \Delta h_{t-i} + \sum_{i=1}^{p} \phi_i \Delta d_{t-i} + \\ + \lambda_1 y_{t-1} + \lambda_2 k_{t-1} + \lambda_3 l_{t-1} + \lambda_4 h_{t-1} + \lambda_5 d_{t-1} + \varepsilon_t\end{aligned} \quad (2)$$

where Δ denotes the first difference operator, β is the drift component, and ε_t is assumed to be a white noise process. The ARDL approach estimates $(p+1)^k$ number of regressions to obtain the optimal lag length for each series, where p is the maximum number of lags used and k is the number of variables in equation (1). The optimal lag structure of the first differenced regression is selected by the Akaike Information Criterion (AIC) and the Schwarz Bayesian Criterion (SBC) to simultaneously correct for residual serial correlation and the problem of endogenous regressors (Pesaran and Shin, 1999, p. 386). In order to determine the existence of a long-run relationship between the variables under study, Pesaran, Shin and Smith (2001) propose two alternative tests. First, an *F-statistic* is used to test the joint significance of the first lag of the variables in levels used in the analysis (i. e. $\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = \lambda_5 = 0$), and then a *t-statistic* is used to test the individual significance of the lagged dependent variable in levels (i. e. $\lambda_1 = 0$).

Pesaran, Shin and Smith (2001) provide a set of critical values assuming first that the variables under study are *I(1)* and, second, that such variables are *I(0)*. These authors propose a bounds testing procedure: if the calculated *F*-or *t*-statistics exceed the upper critical bound (UCB), they conclude in favour of a long-run relationship, regardless of the order of integration. However, if these statistics are below the lower critical bound (LCB), the null hypothesis of no cointegration cannot be rejected. Finally, if the calculated *F*- and *t*-statistics are between UCB and LCB, then the decision about cointegration is inconclusive.

When the order of integration for all series is *I(1)* then the decision is based on the UCB; if all the series are *I(0)*, it is based on the LCB.

The test statistics based on equation (2) have a different distribution under the null hypothesis of no level relationships, depending on whether the regressors are all *I(0)* or all *I(1)*. Further, in both cases the distribution is non-standard. Pesaran, Shin and Smith (2001) provide critical values for the cases where all regressors are *I(0)* and the cases where all regressors are *I(1)*, and suggest that these critical values be used as bounds for the more typical cases where the regressors are a mixture of *I(0)* and *I(1)*.

If cointegration exists, the conditional long-run model is derived from the reduced form equation (2) when the series in first differences are jointly equal to zero (i. e., $\Delta y=\Delta k=\Delta l=\Delta d=0$). The calculation of these estimated long-run coefficients is given by:

$$y_t = \delta_1 + \delta_2 k_t + \delta_3 l_t + \delta_4 h_t + \delta_5 d_t + \xi_t \quad (3)$$

where $\delta_1 = \frac{-\beta}{\lambda_1}$; $\delta_2 = \frac{-\lambda_2}{\lambda_1}$; $\delta_3 = \frac{-\lambda_3}{\lambda_1}$; $\delta_4 = \frac{-\lambda_4}{\lambda_1}$; $\delta_5 = \frac{-\lambda_5}{\lambda_1}$; and ξ_t is a random error. The standard error of these long-run coefficients can be calculated from the standard errors of the original regression using the delta method.

Finally, if a long-run relation is found, an error correction representation exists which is estimated from the following reduced form equation:

$$\Delta y_t = \sum_{i=1}^{p} \theta_i \Delta y_{t-1} + \sum_{i=1}^{q_1} \varpi_i \Delta k_{t-1} + \sum_{i=1}^{q_2} \pi_i \Delta l_{t-1} + \sum_{i=1}^{q_3} \tau_i \Delta h_{t-1} + \sum_{i=1}^{q_4} \kappa_i \Delta d_{t-1} + \eta ECM_{t-1} \quad (4)$$

4. Data and empirical results

4.1. Data

We estimate equation (4) with annual data for eleven EMU countries: both central (Austria, Belgium, Finland, France, Germany and the Netherlands) and peripheral countries (Greece, Ireland, Italy, Portugal and Spain)[7]. Even though the ARDL-based estimation procedure used in the paper can be reliably used in small samples, we use long spans of data covering the period 1960-2012 (i.e., a total of 52 annual observations) to explore the dimension of historical specificity and to capture the long-run relationship associated with the concept of cointegration (see, e. g., Hakkio and Rush, 1991).

To maintain as much homogeneity as possible for a sample of 11 countries over the course of five decades, our primary source is the European Commission´s AMECO database[8]. We then strengthen our data with the use of supplementary data sourced from International Monetary Fund (International Financial Statistics) and the World Bank (World Development Indicators). We use GDP, capital stock and public debt at 2010 market prices for the level of output, the stock of physical capital and the stock of public debt, as well as civilian employment and life expectancy at birth for indicators of the labour input and human capital[9]. The precise definitions and sources of the variables are given in Appendix 1.

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[7] This distinction between European central and peripheral countries has been used extensively in the empirical literature. The two groups we consider roughly correspond to the distinction made by the European Commission (1995) between those countries whose currencies continuously participated in the European Exchange Rate Mechanism (ERM) from its inception and which maintained broadly stable bilateral exchange rates with each other over the sample period, and those countries whose currencies either entered the ERM later or suspended their participation in the ERM, as well as fluctuating widely in value relative to the Deutschmark. These two groups are also roughly the ones found in Jacquemin and Sapir (1996), who applied multivariate analysis techniques to a wide set of structural and macroeconomic indicators, to form a homogeneous group of countries. Moreover, these two groups are basically the same as the ones found in Ledesma-Rodríguez et al. (2005) according to economic agents' perceptions of the commitment to maintain the exchange rate around a central parity in the ERM, and those identified by Sosvilla-Rivero and Morales-Zumaquero (2012) using cluster analysis when analysing permanent and transitory volatilities of EMU sovereign yields.

[8] http://ec.europa.eu/economy_finance/db_indicators/ameco/index_en.htm

[9] As explained in Appendix 1, following Sachs and Warner (1997), we use life expectancy at birth as the human-capital proxy. Other proxies commonly used for human capital such as years of secondary education and enrolment at secondary

4.2. Time series properties

Before carrying out the ARDL cointegration exercise, we test for the order of integration of the variables by means of the Augmented Dickey-Fuller (ADF) tests. This is necessary just to ensure that none of our variables is only stationary at second differences (i. e., *I(2)*). The results, shown in Table 1, decisively reject the null hypothesis of non-stationarity, suggesting that both variables can be treated as first-difference stationary[10].

Table 1. Augmented Dickey-Fuller tests for unit roots.

Panel A: I (2) versus I (1) (Variables in first differences)				
Country	**Variable**	τ_τ	τ_μ	**T**
AT	Δy	--6.5127	-5.1999*	-2.7422*
	Δk	-4.3308*	-3.6206*	-2.8238*
	Δl	-5.9083*	-5.3123*	-4.5947*
	Δh	-9.9420*	-9.9180*	-2.7413*
	Δd	-5.7918*	-5.6235*	-2.7181*
BE	Δy	-6.7061*	-5.0801*	-2.9577*
	Δk	-4.2892*	-3.7822*	-2.6954*
	Δl	-4.8361*	-4.5554*	-4.1708*
	Δh	-11.0268*	-11.0715*	-3.2521*
	Δd	-7.2830*	-3.7436*	-2.7532*
FI	Δy	-4.8867*	-4.5320*	-3.3071*
	Δk	-3.7701**	-3.8441*	-2.6211*
	Δl	-4.5945*	-4.6448*	-4.6380*
	Δh	-5.9301*	-4.0088*	-3.0615*
	Δd	-4.1571**	-4.2012*	-3.5862*
FR	Δy	-4.8869*	-4.5320*	-3.3071*
	Δk	-3.6816**	-3.0692**	-2.8730*
	Δl	-4.8908*	-4.9177*	-2.9013*
	Δh	-7.0261*	-7.0713*	-3.2521*
	Δd	-4.6158*	-4.6150*	-4.1180*
GE	Δy	-6.6679*	-5.1871*	-3.3196*
	Δk	-3.7030**	-3-6413*	-2.7401*
	Δl	-5.9950*	-5.7201*	-5.2289*
	Δh	-7.9188*	-7.4507*	-2.6810*
	Δd	-4.7909*	-4.4196*	-2.5651**
GR	Δy	-4.9108*	-3.8706*	-3.5100*
	Δk	-4.1123**	-3.6180*	-2.6658*
	Δl	-4.1775*	-3.2877**	-2.7391*
	Δh	-7.5080*	-6.7105*	-2.8612*
	Δd	-9.1968*	-8.5823*	-2.8743*
IE	Δy	-3.9471**	-3.5356*	-2.7748*
	Δk	-4.0129**	-3.7324*	-2.6380*
	Δl	-4.7243*	-3.9504*	-3.1723*
	Δh	-5.2499*	-3.1738**	-2.6364*
	Δd	-3.6018**	-3.6301*	--3.1692*
IT	Δy	-6.9406*	-4.2181*	-2.6475*
	Δk	-4.5159*	-3.5312**	-2.7899*

school were available only from 1980. Additionally, the proxy years of secondary education did not change during the sample period. As shown in Jayachandran and Lleras-Muney (2009), longer life expectancy encourages human capital accumulation, since a longer time horizon increases the value of investments that pay out over time. Moreover, better health and greater education are complementary with longer life expectancy (Becker, 2007).

[10]These results were confirmed using Phillips-Perron (1998) unit root tests controlling for serial correlation and the Elliott, Rothenberg, and Stock (1996) Point Optimal and Ng and Perron (2001) unit root tests for testing non-stationarity against the alternative of high persistence. These additional results are not shown here for reasons of space, but they are available from the authors upon request.

	Δl	-4.0228**	-4.0473*	-4.0761*
	Δ***h***	-5.7923*	-4.0831*	-2.9108*
	Δ***d***	-4.6082*	-3.6530*	-2.9241*
NL	Δ***y***	-4.3834*	-3.4255**	-2.6215*
	Δ***k***	-4.2530*	-3.1562**	-2-6234*
	Δl	-5.7439*	-5.8074*	-4.5647*
	Δ***h***	-9.0270*	-8.5068*	-2.9240*
	Δ***d***	-5.3582*	-4.9341*	-3.8121*
PT	Δ***y***	-4.7999*	-3.5718*	-2.5546**
	Δ***k***	-4.2971*	-2.9443**	-2.5840**
	Δl	-4.7487*	-4.7232*	-4.6853*
	Δ***h***	-5.7846*	-5.4675*	-2.7329*
	Δ***d***	-4.0644**	-3.9994*	-2.8629*
SP	Δ***y***	-3.5807**	-3.6355*	-2.6507*
	Δ***k***	-3.9787**	-3.3918**	-2.7152*
	Δl	-4.4395*	-3.6134*	--2.7684*
	Δ***h***	-7.1213*	-6.9283*	-2.7529*
	Δ***d***	-3.6815**	-3.8129*	--2.8241*

Table 1 (Continued)

Panel B: I (1) versus I (0) (Variables in levels)				
Country	**Variable**	τ_τ	τ_μ	τ
AT	***Y***	-1.3393	-2.4451	2.3954
	K	-0.6238	-2.4602	-0.0349
	L	-2.1348	1.6423	3.5707
	H	-2.2066	-0.2614	1.9615
	D	-3.0156	1.1100	3.5156
BE	***Y***	-2.0986	-2-1541	1.7470
	K	-1.7936	-2.5072	0.6156
	L	-1.3175	0.3671	1.8619
	H	-3.1226	-1.0485	0.6528
	D	-1.3880	-1.2012	1.3224
FI	***Y***	-1.4191	-1.8605	2.5771
	K	-1.7451	-2.3438	0.9656
	L	--3.0428	-2.4541	0.5916
	H	-2.4975	0.2117	2.7514
	D	-1.8771	-0.7870	1.5818
FR	***Y***	-1.5816	-2.0082	1.3944
	K	-1.8122	-2.3024	0.7936
	L	-1.9436	-1.6164	0.9568
	H	-3.1226	0.4458	2.5123
	D	-3-0927	-0.1796	1.8067
GE	***y***	-1.5816	-2.0082	1.3944
	k	-1.8122	-2.3024	0.7936
	l	-1.9436	-1.6164	0.9568
	h	-2.2338	-1.5692	1.3238
	d	-0.3146	-1.6901	2.5730
GR	***y***	-1.0010	-2.3408	1.3569
	k	-1.5597	-2.4808	-0.5418
	l	-2.2558	-2.0543	-0.3281
	h	-1.5812	-0.7191	1.5861
	d	-1.1751	-1.4518	1.1216
IE	***y***	-1.9512	-0.8449	2.2557
	k	-3.0149	-1.6303	0.9326
	l	-1.9729	-0.3138	1.3973
	h	-2.1733	-2.0531	-1.2554
	d	-2.2974	-0.7554	1.4304
IT	***y***	-2-1720	-0.5518	2.3052
	k	-2.4669	-0.5135	0.6318
	l	-3.1509	-1.2592	0.3692
	h	-0.5641	-1.4814	2.0789
	d	-0.5985	-2.4603	2.1287
NL	***y***	-1.8167	-2.4855	2.2671
	k	-2.7912	-2.4371	0.1985
	l	-1.2728	-0.2763	1.7524
	h	-2.2529	-0.1643	1.9099
	d	-1.3819	-0,1586	1.0583
PT	***y***	-0.7924	-2.1028	1.8841
	k	0.5611	-2.0484	0.1611
	l	-1.3539	-1.1791	0.7371

	h	-1.8500	-2.3604	1.7143
	d	-1.0314	-1.0858	1.2001
SP	***y***	-1.5694	-2.1594	1.5243
	k	-1.9370	-2.1556	1.6316
	l	-2.4506	-1.6025	0.4907
	h	-1.7033	-1.4070	1.7045
	d	-2.2347	-0.3025	1.8015

Notes: The ADF statistic is a test for the null hypothesis of a unit root.
τ_τ, τ_μ and τ denote the ADF statistics with drift and trend, and with and without drift respectively.
* and ** denote significance at the 1% and 5% levels respectively. Critical values based on MacKinnon (1996)
AT, BE, FI, FR, GE, GR, IE, IT, NL, PT and SP stand for Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Spain respectively.

We also compute the Kwiatkowski *et al.* (1992) (KPSS) tests, where the null is a stationary process against the alternative of a unit root. As argued by Cheung and Chinn (1997), the ADF and KPSS tests can be viewed as complementary, rather than in competition with one another; therefore, we can use the KPSS tests to confirm the results obtained by the ADF tests. As can be seen in Table 2, the results fail to reject the null hypothesis of stationarity in first-difference but strongly reject it in levels.

Table 2. KPSS tests for stationarity

Panel A: I (2) versus I (1) (Variables in first differences)			
Country	**Variable**	τ_τ	τ_μ
AT	Δ***y***	0.0812	0.3165
	Δ***k***	0.0675	0.0304
	Δ***l***	0.1068	0.3145
	Δ***h***	0.1023	0.1011
	Δ***d***	0.1129	0.2232
BE	Δ***y***	0.1118	0.3379
	Δ***k***	0.0580	0.3120
	Δ***l***	0.0943	0.3108
	Δ***h***	0.0938	0.0936
	Δ***d***	0.1073	0.2062
FI	Δ***y***	0.0679	0.3146
	Δ***k***	0.1125	0.3560
	Δ***l***	0.0596	0.0611
	Δ***h***	0.0820	0.0892
	Δ***d***	0.1033	0.1060
FR	Δ***y***	0.0679	0.3126
	Δ***k***	0.1239	0.2678
	Δ***l***	0.0784	0.0779
	Δ***h***	0.0934	0.0936
	Δ***d***	0.1032	0.1938
GE	Δ***y***	0.1118	0.3324
	Δ***k***	0.1075	0.3144
	Δ***l***	0.1110	0.2663
	Δ***h***	0.1042	0.3154
	Δ***d***	0.1121	0.3270
GR	Δ***y***	0.1065	0.3143
	Δ***k***	0.1107	0.3385
	Δ***l***	0.1065	0.1740
	Δ***h***	0.0636	0.3166
	Δ***d***	0.0496	0.3061
IE	Δ***y***	0.1114	1.1288
	Δ***k***	0.0697	0.1748
	Δ***l***	0.1017	0.2174
	Δ***h***	0.0598	0.3140

	Δ*d*	0.1082	0.1076
IT	**Δ*y***	0.0826	0.3291
	Δ*k*	0.0864	0.3267
	Δ*l*	0.0751	0.1335
	Δ*h*	0.1052	0.2715
	Δ*d*	0.0891	0.3154
NL	**Δ*y***	0.0972	0.2974
	Δ*k*	0.0912	0.3146
	Δ*l*	0.1015	0.1524
	Δ*h*	0.0648	0.2608
	Δ*d*	0.0992	0.2619
PT	**Δ*y***	0.0648	0.3184
	Δ*k*	0.1039	0.2679
	Δ*l*	0.1017	0.1912
	Δ*h*	0.0853	0.2618
	Δ*d*	0-1044	0.2150
SP	**Δ*y***	0.1175	0.2670
	Δ*k*	0.0639	0.2528
	Δ*l*	0.0878	0.1125
	Δ*h*	0.1150	0.2207
	Δ*d*	0.0806	0.0790

Table 2 (Continued)

Panel B: I (1) versus I (0) (Variables in levels)			
Country	**Variable**	τ_τ	τ_μ
AT	***y***	0.2249*	0.8641*
	k	0.2487*	0.8682*
	l	0.2198*	0.8092*
	h	0.2470*	0.8737*
	d	0.2261*	0.8394*
BE	***y***	0.2171*	0.8634*
	k	0.2335*	0.8706*
	l	0.2238*	0.8244*
	h	0.2368*	0.8749*
	d	0.2634*	0.7943*
FI	***y***	0.2199*	0.8604*
	k	0.2568*	0.8670*
	l	0.2776*	0.5317**
	h	0.2950*	0.8720*
	d	0.2386*	0.7864*
FR	***y***	0.2349*	0.8648*
	k	0.2419*	0.8593*
	l	0.2195*	0.9126*
	h	0.1995**	0.8604*
	d	0.1532**	0.8377*
GE	***y***	0.2349*	0.8648*
	k	0.2419*	0.8593*
	l	0.2195*	0.9126*
	H	0.1763**	0.8788*
	D	0.2226*	0.8645*
GR	***Y***	0.1885**	0.8998*
	K	0.2449*	0.8242*
	L	0.2038**	0.7367*
	H	0.2352*	0.8741*
	D	0.1988**	0.8221*
IE	***Y***	0.1786**	0.8617*
	K	0.1889*	0.8693*
	L	0.2182*	0.7515*
	H	0.2235*	0.8038*
	D	0.1988**	0.8926*
IT	***Y***	0.2442*	0.8301*
	K	0.2604*	0.8597*
	L	0.2762*	0.7464*
	H	0.2135**	0.8789*
	D	0.2440*	0.8250*
NL	***Y***	0.2164*	0.8650*
	K	0.2243*	0.8592
	L	0.1628**	0.8434*
	H	0.1509**	0.8612*
	D	0.1561**	0.9626*

PT	***Y***	0.2331*	0.8541*
	K	0.1533**	0.8666*
	L	0.2141**	0.7869*
	H	0.2337*	0.8722*
	D	0.1048*	0.8334*
SP	***Y***	0.1694**	0.8610*
	K	0.1643**	0.8772*
	L	0.1983**	0.7596*
	H	0.1926**	0.8762*
	D	0.2987*	0.8994*

Notes: The KPSS statistic is a test for the null hypothesis of stationarity.
τ_τ and τ_μ denote the KPSS statistics with drift and trend, and with drift respectively.
* and ** denote significance at the 1% and 5% levels respectively. Asymptotic critical values based on Kwiatkowski *et al.* (1992. Table 1)
AT, BE, FI, FR, GE, GR, IE, IT, NL, PT and SP stand for Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Spain respectively.

The single order of integration of the variables encourages the application of the ARDL bounds testing approach to examine the long-run relationship between the variables.

4.3. Empirical results from the ARDL bounds test

The estimation proceeds in stages. In the first stage, we specify the optimal lag length for the model (in this stage, we impose the same number of lags on all variables as in Pesaran, Shin and Smith*,* 2001). The ARDL representation does not require symmetry of lag lengths; each variable may have a different number of lag terms**.** As mentioned above, we use the AIC and SBC information criteria to guide our choice of the lag length. For the test of serial correlation in the residual, we use the maximum likelihood statistics for the first and fourth autocorrelation, denoted as $\chi^2_{SC}(1)$ and $\chi^2_{SC}(4)$ respectively. Due to constraints of space, these results are not shown here but they are available from the authors upon request.

Next we test for the existence of a long-run relation between the output and its components, as suggested by equation (1). Table 3 gives the values of the *F*- and *t*-statistics for the case of unrestricted intercepts and no trends (case III in Pesaran, Shin and Smith, 2001)[11]. These statistics are compared with the critical value bounds provided in Tables CI

[11] We also consider two additional scenarios for the deterministics: unrestricted intercepts, restricted trends; and unrestricted intercepts, unrestricted trends (cases IV and V in Pesaran, Shin and Smith, 2001). These additional results are

and CII of Pesaran, Shin and Smith (2001) and depend on whether an intercept and/or trend is included in the estimations, suggesting the existence of a single long-term relationship in which the production level would be the dependent variable and the stock of physical capital, the labour employed, the human capital and the stock of public debt the independent variables[12].

Table 3. *F*- and *t*-statistics for testing the existence of the long-run model

Country	**Bound testing to cointegration**		
	ARDL(p,q_1,q_2,q_3,q_4,q_5)	***F*-statistic**	***t*-statistic**
AT	(4, 3, 3, 4, 4)	6.8148*	-5.2908*
BE	(1, 2, 4, 4, 0)	5.0451**	-3.7093**
FI	(1, 4, 3, 1, 2)	5.0352**	-3.8220**
FR	(1, 0, 2, 4, 3)	4.1633**	-3.8685**
GE	(2, 2, 1, 0, 2)	6.0071*	-4.7023*
GR	(1, 3, 0, 0, 0)	4.5088**	-3.6953**
IE	(1, 2, 1, 0, 0)	4.6117**	-3.7436**
IT	(3, 2, 0, 4, 1)	5.3960*	-3.6283**
NL	(1, 4, 3, 4, 4)	6.7727*	-4.2859*
PT	(1, 3, 3, 0, 2)	4.3225**	-3.8598**
SP	(1, 3, 2, 0, 3)	4.3497**	-4.0635**

Notes: p,q_1,q_2,q_3,q_4 and q_5 denote respectively the optimal lag length for Δy_{t-i}, Δk_{t-i}, Δl_{t-i}, Δh_{t-i} and Δd_{t-i} in the UECM model (4) without deterministic trend.
* and ** indicate that the calculated *F*- and *t*-statistics are above the upper critical bound at 1% and 5% respectively.
AT, BE, FI, FR, GE, GR, IE, IT, NL, PT and SP stand for Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Spain respectively.

not shown here due to space constraints, but they are available from the authors upon request. Nevertheless, our estimation results indicate that the intercepts are always statistically significant, whereas the trends are not.

12 These results were confirmed using Johansen's (1991, 1995) approach in order to test for cointegration between *y*, *k*, *l*, *h* and *d*. In all cases, the trace tests indicate the existence of only one cointegrating equation at least at the 0.05 level, which can be normalized with *y* as the dependent variable. These additional results are not shown here due to space constraints but are available from the authors upon request.

The estimated long-run relationships between the variables are reported in Table 4.

Table 4. Long-run analysis

Country	Estimation results
AT	y_t= -0.0041 + 0.2964k_t + 0.3278l_t + 0.0855h_t - 0.1288d_t (-3.0331) (6.6280) (6.1756) (2.8922) (-4.3352)
BE	y_t= -0.0982 + 0.3963k_t+ 0.4515l_t+ 0.4210h_t- 0.0621d_t (-3.2144) (6.0705) (7.7879) (2.9783) (-5.5117)
FI	y_t= -0.0632 + 0.4261k_t + 0.4112l_t+ 0.5375h_t - 049021d_t (-3.5612) (5.6646) (7.2917) (4.13723) (-5.1371)
FR	y_t= -0.0504 + 0.4288$_t$ + 0.4277l_t+ 0.5068h_t – 0.5439d_t (-3.6212) (5.8255) (3.8349) (3.9981) (-5.8665)
GE	y_t= -0.0633 +0.4970k_t + 0.5204l_t+ 0.5843h_t - 0.0397d_t (-3.0207) (5.5325) (2.9449) (2.9769) (-2.9149)
GR	y_t= -0.1547 +0.2445k_t + 0.3115l_t+ 0.3457h_t - 0.0787d_t (-3.0207) (5.4884) (3.4825) (2.9321) (-3.1347)
IE	y_t= 0.3738 +0.2324k_t + 0.3945l_t+ 0.1311h_t - 0.0492d_t (2.9965) (6.1718) (3.5311) (3.1237) (-7.7831)
IT	y_t= 0.2315 +0.3117k_t + 0.4720l_t+ 0.1422h_t - 0.0831d_t (-3.1429) (5. 8428) (6.3747) (3.7232) (-6.7227)
NL	y_t= 0.0222 +0.4435k_t + 0.3576l_t+ 0.3571h_t - 0.0966d_t (3.0545) (6.2867) (6.3197) (4.1977) (-7.3175)
PT	y_t= 0.2740 +0.3297k_t + 0.3732$_t$+ 0.2054h_t - 0.3536d_t (3.0336) (4.2039) (2.9423) (2.9473) (-6.3360)
SP	y_t= -0.0615 +0.4891k_t + 0.3241$_t$+ 0.3527h_t - 0.3356d_t (-3.0515) (7.3996) (4.0399) (3.3946) (-4.8721)

Notes: In the ordinary brackets below the parameter estimates, the corresponding *t*-statistics are shown. AT, BE, FI, FR, GE, GR, IE, IT, NL, PT and SP stand for Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Spain respectively.

In order to examine the short-term dynamics of the model, we estimate an error-correction model associated with the above long-run relationship. These results are reported in Table 5, which shows that the short-run analysis seems to pass diagnostic tests such as normality of error term, second-order residual autocorrelation and heteroskedasticity (χ^2_N, χ^2_{SC} and χ^2_H respectively), rendering a satisfactory overall regression fit, as measured by the R^2 value (ranging from 0.6250 for France to 0.8947 for Germany).

Table 5. Short-run analysis

Country		Adjusted R^2	DW Test	χ^2_N	χ^2_{SC}	χ^2_H
AT	$\Delta y_t = 0.3357\Delta y_{t-1} + 0.2273\Delta y_{t-2} + 3.4635\Delta k_t + 1.6406\Delta k_{t-1} +$ (4.9587) (3.9848) (7.1120) (3.2281) $+ 0.5122\Delta l_t + 1.8360\Delta h_{t-1} - 0.1050\Delta d_t + 0.1169\Delta d_{t-1}$ (4.2105) (3.8970) (-3.5605) (3.5604) $+ 0.0771\Delta d_{t-3} - 0.5184 ECM_{t-1}$ (3.0602) (-7.5397)	0.8052	2.1035	1.3631 [0.5058]	0.4403 [0.8024]	6.7833 [0.7457]
BE	$\Delta y_t = 2.9234\Delta k_t + 1.9681\Delta k_{t.1} + 0.5450\Delta l_t + 2.1247.\Delta h_{t-1}$ (6.4798) (4.4537) (3.3954) (3.7914) $+ 1.8736\Delta h_{t-2} + 1.2525\Delta h_{t-3} - 0.0186\Delta d_t - 0.3011 ECM_{t-1}$ (3.5399) (3.5437) (-4.3458) (-4.3245)	0.6991	2.1682	0.7188 [0.6980]	1.6363 [0.4412]	8.7743 [0.5536]
FI	$\Delta y_t = 3.9141\Delta k_t + 4.2948\Delta k_{t-1} + 2.0681\Delta k_{t-2} + 0.7669\Delta l_t$ (5.9736) (6.5524) (3.8293) (5.3695) $+ 0.1491\Delta l_{t-2} + 1.2080\Delta h_t + 0.0589\Delta d_{t-1} - 0.5431 ECM_{t-1}$ (3.2632) (3.9132) (4.9503) (-5.0585)	0.8947	2.1812	1.8337 [0.3998]	0.6935 [0.7070]	8.3739 [0.3978]
FR	$\Delta y_t = 0.5483\Delta k_t + 2.7066\Delta l_t + 1.3583\Delta l_{t-2} + 2.7571\Delta h_{t-1}$ (4.1446) (6.7447) (3.2368) (3.4479) $- 0.0540\Delta d_{t-1} - 0.1594 ECM_{t-1}$ (-3.2524) (-4.7831)	0.6250	2.0703	1.0284 [0.5980]	2.7751 [0.2497]	11.6117 [0.1514]
GE	$\Delta y_t = 0.1245\Delta y_{t-1} + 4.5310\Delta k_t + 2.9485\Delta k_{t-1} + 0.6069\Delta l_t$ (3.4013) (6.2536) (-4.8966) (5.0089) $+ 0.3283\Delta h_t + 0.0888\Delta d_{t-1} - 0.5431 ECM_{t-1}$ (3.5278) (3.3255) (-5.7911)	0.8654	2.0727	1.7700 [0.4127]	2.0859 [0.3524]	8.8985 [0.3509]
GR	$\Delta y_t = 4.1491\Delta k_t + 2.2586\Delta k_{t-1} + 0.3111\Delta l_t - 0.0195\Delta d_{t-1}$ (5.6965) (4.4863) (3.2133) (-3.7315) $- 0.1898 ECM_{t-1}$ (-5.3528)	0.8233	2.0170	1.6641 [0.4352]	1.7768 [0.4113]	2.7153 [0.7438]
IE	$\Delta y_t = 4.1491\Delta k_{t-1} + 0.5946\Delta l_t + 3.6624\Delta h_{t-1}$ (4.2518) (5.2966) (3.3309) $- 0.0770\Delta d_t - 0.1750 ECM_{t-1}$ (-3.9022) (-6.8543)	0.6679	1.9876	0.4433` [0.8012]	2.6952 [0.2599]	6.6772 [0.2458]
IT	$\Delta y_t = 0.2820\Delta y_{t-1} + 0.1810\Delta y_{t-1} + 5.3075\Delta k_t + 3.4087\Delta k_{t-1}$ (3.9118) (3.5678) (7.7289) (6.6994) $+ 0.1468\Delta l_t + 0.8079\Delta h_{t-3} - 0.0770\Delta d_t - 0.2619 ECM_{t-1}$ (3.9912) (3.6477) (-3.5299) (-8.1758)	0,8933	1.9866	0.9128 [0.6335]	5.5305 [0.0630]	13.3690 [0.0998]
NL	$\Delta y_t = 3.3069\Delta k_t + 2.1191\Delta k_{t-1} + 0.8953\Delta k_{t-2} + 0.0971\Delta l_{t-1}$ (6.3711) (5.6906) (3.1207) (3.9035) $+ 0.1468\Delta h_{t-2} + 1.7061\Delta h_{t-3} - 0.1082\Delta d_{t-1} + 0.0615\Delta d_{t-2}$ (4.5438) (4.3054) (-5.2418) (3.1050) $+ 0.0152\Delta d_{t-3} - 0.3592 ECM_{t-1}$ (3.4608) (-7.9430)	0.8861	2.2133	2.6149 [0.2706]	4.0878 [0.1295]	11.7712 [0.54]
PT	$\Delta y_t = 1.9415\Delta k_t + 1.5631\Delta k_{t-1} + 0.8682\Delta k_{t-2} + 0.4788\Delta l_t$ (5.9323) (3.5441) (3.6882) (3.5463) $+ 0.4276\Delta l_{t-2} + 0.2646\Delta h_t + 0.0634\Delta d_{t-1} - 0.1293 ECM_{t-1}$ (3.7681) (3.6499) (3.1867) (-6.3868)	0.7258	2.1636	1.3451 [0.5104]	2,3736 [0.3052]	4.8974 [0.7685]
SP	$\Delta y_t = 3.1383\Delta k_t + 1.1341\Delta k_{t-1} + 0.6438\Delta k_{t-2} + 0.1711\Delta l_t$ (7.0649) (5.2017) (3.0516) (3.3203) $+ 0.2222\Delta l_{t-1} + 0.9868\Delta h_t + 0.0302\Delta d_t + 0.0366\Delta d_{t-2}$ (3.4312) (3.7048) (3.4194) (3.0740) $- 0.1757 ECM_{t-1}$ (-6.0164)	0.8213	2.1052	2.9858 [0.2247]	2.3263 [0.3125]	10.2919 [0.3274]

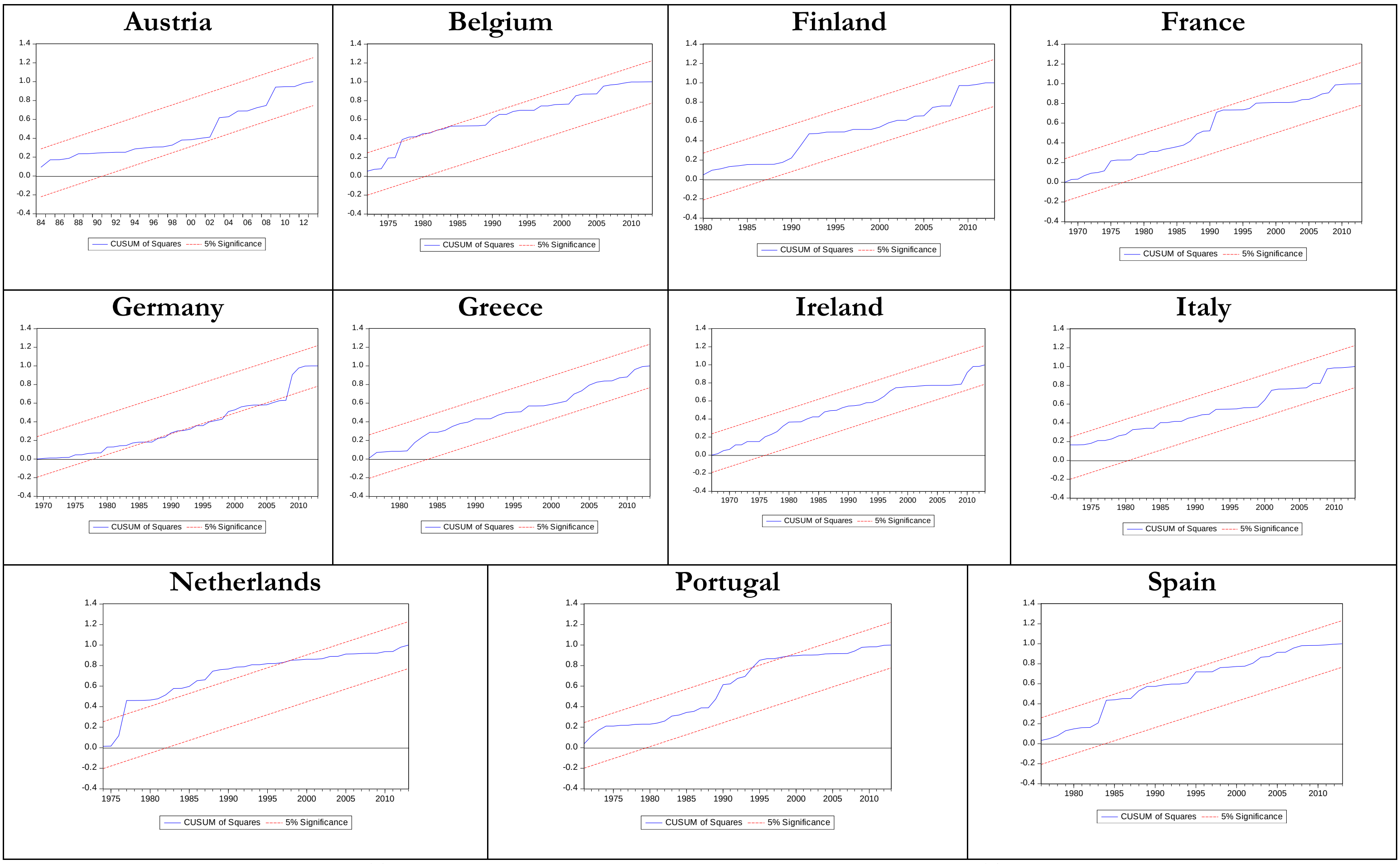
Notes: AT, BE, FI, FR, GE, GR, IE, IT, NL, PT and SP stand for Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Spain respectively.
In the ordinary brackets below the parameter estimates, the corresponding *t*-statistics are shown.
χ^2_N, χ^2_{SC} and χ^2_H are the Jarque-Bera test for normality, the Breusch-Godfrey LM test for second-order serial correlation and the Breusch-Pagan-Godfrey test for heteroskedasticity. In the square brackets, the associated probability values are given.

We examine the stability of long-run coefficients using the cumulative sum of recursive residuals (CUSUM) and the cumulative sum of squares of recursive residuals (CUSUMSQ) tests (Figures 1 and 2). These tests are applied recursively to the residuals of the error-correction model shown in Table 5. Since the test statistics remain within their critical values (at a marginal significance level of 5%), we are able to confirm the stability of the estimated equations.

Figure 1. Plot of cumulative sum of recursive residuals

Note: The straight lines represent the critical bounds at a 5% significance level. Belgium: 1977-1982; Spain: 1983-1987.

Figure 2. Plot of cumulative sum of squares of recursive residuals



Note: The straight lines represent the critical bounds at a 5% significance level. The Netherlands: 1977-1996; Portugal: 1994-1998.

Some very interesting results can be drawn from the empirical evidence presented in Tables 4 and 5. First, the long term effect of debt on economic performance (Table 4) seems to support the "conventional" view (Elmendorf and Mankiw, 1999), since it registers a negative value in all EMU countries. According to this approach, if the decrease in public savings is not fully compensated by an increase in private savings, total investment will drop, implying a negative effect on GDP. However, the magnitude of the negative impact differs significantly across countries, implying that our conclusions need to be qualified. While it is sizeable in the case of France (-0.5439), Spain (-0.3356), Portugal (-0.3356) and

Austria (-0.1288), in the rest of countries, although negative, the magnitude is very small with values close to zero. Ireland (-0.0492), Finland (-0.0490) and Germany (-0.0397) are the countries with the lowest negative impact. Therefore, our results suggest that, even though debt has a negative impact on output in all EMU countries, with the exception of France, Spain, Portugal and Austria, its magnitude is negligible.

Table 5 shows that the short-term impact of debt on economic performance differs clearly across countries. With respect to EMU peripheral countries, in spite of its important negative impact in Portugal and Spain in the long run, its effect in the short run is positive (just one period lagged in the case of Portugal). However, in Greece, Ireland and Italy an increase in public debt has a negative effect on GDP, not only in the long run, but in the short run as well. Among EMU central countries, it is noticeable that in Germany and Finland the effect of public debt on GDP is positive in the short run (one period lagged), despite the negative (though very small) effect in the long run. Finally, in the case of Austria, Belgium, France and the Netherlands our results suggest that public debt has a negative impact on economic activity in both the short and the long run.

All in all, it is interesting that in our empirical analysis we did not find a positive long-run relationship between public debt and output in any country, although the short-run link was positive in four EMU countries. Interestingly, in two peripheral countries, Spain and Portugal, while debt exerts an important negative effect on the long-run, its impact is positive in the short-run. These results are highly relevant since these countries have been hit especially hard by both the economic and sovereign debt crises. And, amid the crisis, they received rescue packages (in the Spanish case, to save its banking sector) which were conditional on highly controversial fiscal austerity measures and the implementation of structural reforms to improve competitiveness, whose positive effects are nevertheless typically related to the long run. Few macroeconomic policy debates have generated as much controversy as the austerity argument and, as Europe stagnates, the debate appears to be far from over. In this respect, according to some authors [see Cottarelli and Jaramillo (2013), Delong and Summers (2012), and Perotti (2012)], a reduction in government spending could lower the debt burden and increase the long-term growth perspectives, but it may well have negative effects on demand and production over the period of adjustment. In fact, our results do not favour the same austerity argument in all euro area countries; they indicate that, in the short term, expansionary fiscal policies may have a positive effect on output in some countries such as Spain and Portugal, regardless of its negative impact in the long run. Then, although our findings support the view that the unprecedented sovereign debt levels reached in euro area countries might have adverse consequences for their economies in the long run, they also suggest that the pace of adjustment may differ across them. In particular, within the peripheral countries, policy measures should bear in mind that while the short-run impact of debt on economic performance is negative in Greece, Ireland and Italy, it is positive in Spain and Portugal. So, in these two countries, our empirical evidence suggests that the pace of fiscal adjustment may be slower than in Greece, Ireland and Italy.

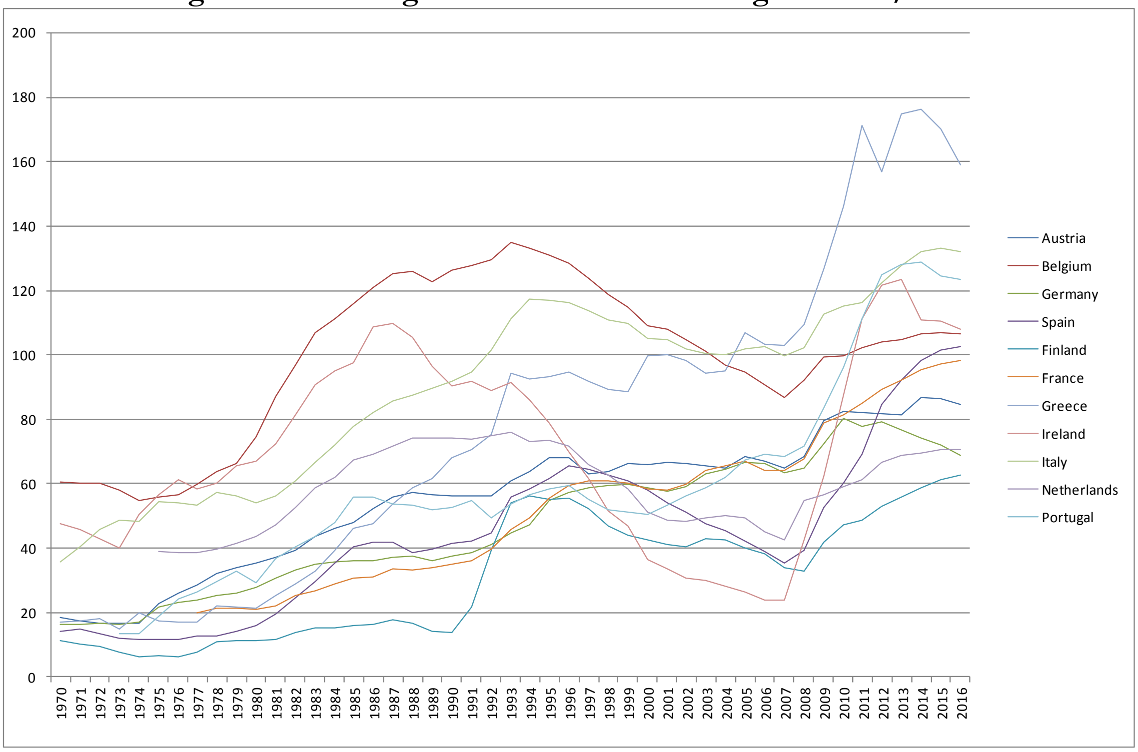
Finally, note that the estimated coefficients for the error correction terms (ECM_{t-1}) represent the speed of adjustment needed to restore equilibrium in the dynamic model following a disturbance. As can be seen, they show how slowly or quickly a variable returns to equilibrium and (as is the case here) they should be negative and significant. Banerjee *et al.* (1998) stated that a highly significant error correction term provides further support for the existence of a stable long-run relationship. The estimated coefficients of the ECM_{t-1} (see Table 5) rank from -0.71 and -0.54 for Austria, Germany and Finland (suggesting a relatively quick speed of adjustment back to the long-run equilibrium) to -0.12 and -0.07

for Portugal and Spain (implying relatively slow reactions to deviations from equilibrium), corroborating the above results.

4.4. Robustness analysis

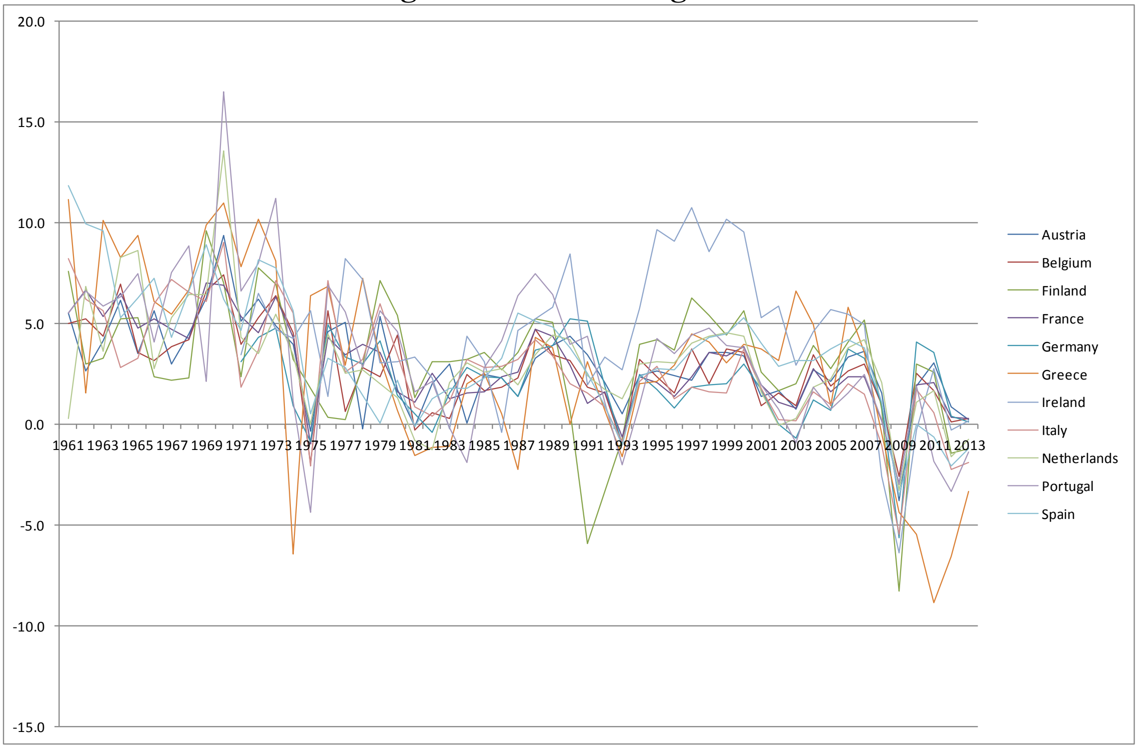
The fact that we have explored the impact of public debt on output during a time period that covers five decades (1960-2012) and extends beyond the economic and sovereign debt crisis (see Figures 3 and 4) may have distorted the results, in view of the sudden and significant rise in public debt levels of European countries following government interventions in response to the global financial crisis. These measures included not only fiscal stimulus programmes and bank bailouts, but also social safety nets that work as economic automatic stabilizers by responding to the increase in the unemployment rate.

Figure 3. General government consolidated gross debt/GDP



Source: AMECO (European Commission)

Figure 4. GDP rate of growth



Source: World Development Indicators (World Bank).

Therefore, we also analysed the time-varying impact of public debt on short-term economic performance by re-estimating the short-run model for two sub-periods, 1975-1992 and 1993-2007 (defined according to the Euro Area Business Cycle Dating Committee)[13]. We assessed whether the impact of government debt on output differed between the two periods. The results (not shown here due to space constraints but available from the authors upon request) qualify the previous ones.

In the case of central countries, public debt has a positive impact on output during the second sub-period (1993-2007) in the Netherlands, and during both sub-periods (1975-1992 and 1993-2007) in France and Germany. In peripheral countries, we also find a positive impact of debt on output during the second sub-period (1993-2007) in the case of

[13] Center for Economic Policy Research (2014)

Greece, Ireland and Italy, and between 1975 and 1992 in the case of Spain. So, according to these new results, debt might also have a positive short-term impact on economic performance in Greece, Ireland and Italy, provided that the economy is not in recession. Nonetheless, if periods of recession are included in the estimation, as Table 5 shows, the short-run effect of debt on output is not positive but negative in those three countries. However, in the Spanish case, the new results reinforce the ones obtained for the whole period, thus highlighting the positive short-term effect of debt on the country's economic performance.

5. Concluding remarks

Despite the severe sovereign debt crisis in the EMU, few papers have examined the relationship between debt and growth for euro area countries. The limited body of literature available lends support to the presence of a common debt threshold across euro-area countries and does not distinguish between short- and long-run effects. So, to our knowledge, no strong case has yet been made for analysing the incidence of debt accumulation on economic growth taking into account the particular characteristics of each euro area economy and examining whether the effects differ depending on the time horizon, even though this potential heterogeneity has been stressed by the literature.

This paper aims to fill this gap in the empirical literature. Unlike previous studies in the euro area, we do not make use of panel techniques but study cross-country differences in the debt-growth nexus both across EMU countries and across time horizons using time series analyses. To this end, our empirical examination of 11 euro area countries (both central and peripheral) during the 1960-2012 period is based on the estimation of a standard growth model including a debt variable as an additional instrument for each country, by means of the ARDL testing approach to cointegration.

As in every empirical analysis, the results must be regarded with caution since they are based on a set of countries over a certain period and a given econometric methodology. Nonetheless, they are in concordance with the conventional view that the positive effect of debt on output is more likely to be felt in the short rather than in the long run. In particular, our empirical evidence suggests a negative effect of public debt on output in the long run. Our findings thus support previous literature indicating that high public debt tends to hamper growth by increasing uncertainty over future taxation, crowding out private investment, and weakening a country's resilience to shocks (see, e. g., Krugman, 1988). However, they admit the possibility that high public debt may have a positive effect in the short run by raising the economy's productive capacity and improving efficiency depending on the characteristics of the country and the final allocation of public debt. Specifically, this short-run positive effect is found in Finland, Germany, Portugal and Spain, suggesting that in a context of low rates of economic growth, the path of fiscal consolidation may differ across the different euro area countries.

This issue is particularly relevant to policymakers because of its implications for the effectiveness of a common fiscal policy, in view of the pronounced differences in the responsiveness of output in the long and short run and across countries. These findings seem to corroborate the idea that there is no "one size fits all" definition of fiscal space but that, conversely, debt limits and fiscal space may be country-specific and depend on each country's track record of adjustment (see, e. g., Ostry *et al.*, 2010).

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Appendix 1: Definition of the explanatory variables and data sources

Variable	Description	Source
Level of Output (Y_t)	Gross domestic product at 2010 market prices	Annual Macroeconomic Database-European Commission (AMECO)
Capital Stock (K_t)	Net capital stock at 2010 market prices	AMECO
Accumulated public debt (D_t)	General government consolidated gross debt at 2010 market prices	AMECO and International Monetary Fund
Labour input (L_t)	Civilian employment	AMECO
Human capital (H_t)	Life expectancy at birth, total (years)	World Development Indicators, World Bank

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