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Short and Long-run Behaviour of Long-term Sovereign Bond Yields

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Abstract

This study assesses the short and long-run behaviour of long-term sovereign bond yields in OECD countries, for the period 1973-2008. We employ a dynamic panel approach to reflect financial and economic integration, and to increase the performance and accuracy of the tests. Given the existence of cross-country dependence regarding sovereign yields and its determinants, we resort to simulation and bootstrap methods for the analysis. Results based on the Common Correlated Effect estimator of Pesaran (2006) and on Panel Error Correction Models to sort out short- and long-run fiscal developments show that in addition to common movements in sovereign yields, investors also consider country differences arising from specific factors (inflation, budgetary and current account imbalances, real effective exchange rates, and liquidity).

JEL-Code: C23, E43, E62, G15, H62.

Keywords: long-term yields, EU, financial integration, panel cointegration, bootstrap.

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Non-technical summary

The idea that government debt accumulation has implications for long-term government bond interest rates is a common feature in a number of – otherwise diverse – theoretical models. One could expect that increases in the debt-to-GDP ratio or in the government deficit ratios may imply an increase in the long-term interest rate, since it may impinge negatively on the credit risk of the sovereign debt liabilities. Indeed, market participants may perceive an additional risk stemming from the implied loosening of fiscal stance under such conditions

From a policymaking point of view the relationship between government debt and deficit, and long-term interest rates is rendered timely in the context of central bank independence when pressures for macroeconomic activism are exercised on fiscal authorities, notably to face severe economic downturns and financial disruption. In the euro area and the EU the effects of fiscal policy stance on long-term interest rates have an additional dimension. Less prudent fiscal policies are not considered to be aligned with the fiscal limits set by the Maastricht treaty. Moreover, it is often argued that large and unsustainable deficits can endanger the coherence of national macroeconomic policies and may jeopardize the price-stability oriented monetary policy.

We assess the short and long-run determinants of real long-term government bond yields for a set of OECD countries, employing a dynamic panel approach for the period 1973-2008, to test for the existence of cointegration between real long-term interest rates and its potential determinants. Furthermore, we also resort to simulation and bootstrap methods to compute the critical values and to take into account the cross-country dependences regarding this segment of the capital markets. Afterwards, we estimate a complete panel error-correction model in order to also uncover the short-run parameters and the speed of convergence to the long-run relationship, taking advantage of non-stationary panel data econometric techniques.

The panel framework allows using information contained in the cross-section dimension and to increase the performance and accuracy of the tests. In addition, cross-country dependence can mirror common changes in the behaviour of fiscal authorities, for instance in the run-up to European and Monetary Union, the Stability and Growth Pact framework and peer pressure. Using the information contained in the cross-section dimension allows reflecting capital markets views, due notably to financial markets integration and liberalisation, or increased business cycle synchronization. From an economic point of view, it is also relevant to find such cross-section dependence, both for the financial series and for the macroeconomic and fiscal variables. In fact, this provides evidence of significant capital market integration at the OECD level, which sovereign government debt issuers cannot discard lightly.

The results of our analysis also show that in addition to common movements in sovereign yields, and credit and liquidity risk, investors are also aware of such country specific fundamentals as inflation, budgetary and current account imbalances, and real effective exchange rates. A better (more positive) government budget balance reduces (as expected) the real long-term interest rate in almost all countries. Moreover, the developments in current account balances also carry relevant long-run information for real interest rates. Indeed, the deterioration of the current account balance would signal a widening gap between savings and investment and long-term interest rates may be pushed upwards.

Moreover, our results illustrate that over the longer run real long-term interest rates and their potential determinants move together in this sample of OECD countries. Therefore, identifying the determinants of real long-term interest rates, over long periods as captured by the cointegration analysis, offers additional valuable information notably for financing choices decisions by the sovereign issuers and government investment decisions. Interestingly, some long-run determinants of real long-term interest rates, which were uncovered in the panel cointegration estimation, such as liquidity, are also relevant from a short-run perspective.

1. Introduction

The idea that government debt accumulation has implications for long-term government bond interest rates is a common feature in a number of – otherwise diverse – theoretical models. The long-run relationship between fiscal variables and long-term interest rates also constitutes an important part of policymakers' conventional wisdom. One could expect that increases in the debt-to-GDP ratio or in the government deficit ratios may imply an increase in the long-term interest rate, since it may impinge negatively on the credit risk and on the quality of the outstanding sovereign debt liabilities. Indeed, market participants may perceive an additional risk stemming from the implied loosening of fiscal stance under such conditions (see Alesina et al., 1992, and Ardagna et al., 2004). However, and as mentioned by Elmendorf and Mankiw (1999), difficulties arise when assessing the fiscal effects on long-term interest rates, since interest rates are likely to be linked to fiscal policy expectations, which is not an easy concept to measure.

Apart from default or creditworthiness, liquidity risk is also relevant for sovereign bond holders. Indeed it is logical to assume that sovereign debt investors look at both credit and liquidity risk, although liquidity seems to play a bigger role in times of market unrest (see, for instance, Beber et al., 2009).

Moreover, several other explanations can be at the root of the long-run developments of long-term yields, in addition to fiscal fundamentals: external variables and imbalances, liquidity issues, inflation rate developments, growth developments, and possible substitution or demonstration effects from the equity segment of the capital markets.

From a policymaking point of view the relationship between government debt and deficit, and long-term interest rates is rendered timely in the context of central bank independence when pressures for macroeconomic activism are exercised on fiscal authorities, notably to face severe economic downturns and financial disruption. In the euro area and the EU the effects of fiscal policy stance on long-term interest rates have an additional dimension. Less prudent fiscal policies are not considered to be aligned with the fiscal limits set by the Maastricht treaty. Moreover, it is often argued that large and unsustainable deficits can endanger the coherence of national macroeconomic policies and may jeopardize the price-stability oriented monetary policy.

In this study we assess the short and long-run determinants of real long-term government bond yields for a set of OECD countries, employing a dynamic panel approach for the period 1973-2008, to test for the existence of cointegration between real long-term interest rates and its potential determinants. Furthermore, we also resort to simulation and bootstrap methods to compute the critical values and to take into account the cross-country dependences regarding this segment of the capital markets. Specifically, we take advantage of non-stationary panel data econometric techniques and the new Common Correlated Effect (CCE) estimator (Pesaran, 2006, that allows common factors in the cross equation covariances to be removed).

Another important issue is how to model the reduced form relationship in the presence of possible non-stationarity in the panel. Indeed, a cursory reading of the formal literature on determinants of real long-term government bond yields in stochastic general equilibrium suggests that given the panel data employed, there could also be relevant short-run effects, which may vary across countries. Thus, in order to address this issue we employ the Pooled Mean Group approach of Pesaran, Shin and Smith (1999) to sort out the long-run versus short-run effects of the EU member states respective fiscal policies. The advantage of such approach is that it addresses the issue of unit-roots in the panel data and also allows for short run versus long run analyses of long-term sovereign bond yields in the same specification. Individual countries may well be on the same long-run path albeit with different short-run cyclical effects.

The panel framework allows using information contained in the cross-section dimension and to increase the performance and accuracy of the tests. In addition, cross-country dependence can mirror common changes in the behaviour of fiscal authorities, for instance in the run-up to European and Monetary Union (EMU), the Stability and Growth Pact (SGP) framework and peer pressure. Using the information contained in the cross-section dimension allows better reflecting capital markets views, due notably to financial markets integration and liberalisation, or increased business cycle synchronization. The existence of possible cross-section dependence, naturally relevant from an economic perspective, has been essentially unaccounted for in the applied related literature. However, one indeed expects capital markets' variables to be rather interlinked, while co-movements and cross-country spillovers are also expected at the macro level. Therefore, we also contribute to the literature in this respect.

Naturally, it is also important to i) grasp to what extent fiscal and macro variables move sovereign yields; and ii) to assess whether country differences arising from specific factors (government debt, current account balance, inflation), on top of common movements, may also be paramount regarding heterogeneous behaviour on sovereign yields. For instance, inflation and exchange rate developments can illustrate the behaviour of the monetary authorities towards price stability. In addition, in the context of financial crisis with overall risk aversion and uncertainty rising and increasing sovereign debt issuance, good fiscal performances also becomes more relevant, from the perspective of financial markets.

The remainder of the paper is organized as follows. Section two reviews the related literature. Section three presents the methodology. Section four conducts the empirical analysis and discusses the results. Section five concludes the paper.

2. Related literature

The participants in the capital markets may perceive additional risks stemming from the loosening of fiscal policies, which would then be reflected in higher bond yields demanded from sovereign issuers. Such increased risks usually also have an adverse impact on the sovereign debt ratings. For instance, Afonso et al. (2007, 2009) show that fiscal developments are among the relevant determinants of a country's credit rating, together with macroeconomic and government effectiveness variables.

On the other hand, capital markets may also value the increased liquidity associated to the existence of additional outstanding sovereign debt for a given country, and a decrease in the long-term yields cannot be discarded as well, given that default risk has been perceived in the past as rather subdued in the EU context (see Codogno et al., 2003, Bernoth et al., 2004, and Afonso and Strauch, 2007).

Certainly, the relationship between fiscal variables, such as government debt and budget deficits, and long-term interest rates and its several possible determinants remains largely an empirical question. Studies done in the 1980s, essentially for the US, in the context of crowding-out discussions were inspired by this debate (see, for instance, Evans, 1985, Wachtel, and Young, 1987, and Rose and Hakes, 1995). Indeed, abundant literature exists on the Ricardian versus non-Ricardian nature of fiscal policy (see, for instance, Afonso, 2008).

The related existing evidence does not seem to be clear cut in favour or against the relationship between government debt, deficit and long-term interest rates relationship. Some more recent literature tries to assess the empirical evidence regarding notably the fiscal determinants of long-term interest rates, notably the relevance of future fiscal variables. For instance, Canzoneri, Cumby and Diba (2002), who evaluate for the US the effect of CBO budget surplus projections on interest rates spreads, conclude that higher projected surpluses imply lower spreads of long-term rates over short-term rates. Engen and Hubbard (2004) regress the current real 10-year treasury rate on CBO 5-year ahead federal debt and deficit projections, and report that increases in the expected federal debt-to-GDP ratio increase the current real 10-year Treasury yield.

Again for the US, Laubach (2009) regresses expected future interest rates on projections published by the CBO and the OMB for the deficit-to-GDP ratio and the debt-to-GDP ratio 5 years ahead. According to the results, a one percentage point increase in the projected deficit-to-GDP ratio is estimated to raise long-term interest rates by roughly 25 basis points. In addition, in related research Thomas and Wu (2009) also used fiscal projections for the US.

For instance, in the context of a no-arbitrage affine term structure model for the US, Dai and Philippon (2005) also report that although the response of sovereign yields to fiscal shocks is mitigated in the shorter side of the yield curve, the response is amplified for the case of the 10-year bonds.

For the EMU countries (except Luxembourg), Faini (2006) argues that an expansionary fiscal policy in one EMU member will have a twofold effect, first on its spreads, and second on the overall level of interest rates for the currency union as a whole. Bernoth, von Hagen, and Schuknecht (2004) report that EU countries' sovereign bonds interest differentials, vis-à-vis Germany or the US, contain risk premia which increase with government debt, deficit, and debt-service, and also depend positively on liquidity, i.e. the issuer's relative bond market size.

In the European Union context, Heppke-Falk and Hüfner (2004) report that monthly deficit forecasts from financial market participants fiscal projections for France, Germany and Italy, over the period 1994-2004, have no significant effect on interest rate swap spreads of 10-year Treasury bonds. Afonso and Strauch (2007) in the context of an event-study of fiscal policy announcements in 2002, show that such fiscal events had small effects on daily swap spreads, mostly around five basis points or less. Using high frequency daily data, from January 1999 to April 2008, Manganelli and Wolswijk (2009) report that for the EMU members government bond spreads react more to short-term interest rate increases when the sovereign credit risk increases and that liquidity also plays a role.

On the other hand, Afonso (2009), using a panel of semi-annual vintages of growth and fiscal forecasts of the European Commission, shows that 10-year government bond yields increase with better growth forecasts, and with decreases in budget balance-to-GDP ratios, signalling that sovereigns may need to pay more to finance anticipated higher budget deficits in the market.[1]

Table 1 offers a summary of some of the findings in the abovementioned related literature, within different methodological frameworks. Interestingly, from the studies surveyed, the concern regarding the assessment of possible cross-section dependences and its technical, empirical, and economic implications for the analysis seems to be essentially absent.

Table 1 – Some existing empirical evidence regarding fiscal determinants of long-term interest rates

Reference	Data frequency	Data sample	Tests performed	Main results
Orr, Edey, and Kennedy (1995)	Quarterly	17 OECD countries (1981:Q1-1994:Q2)	Regression of real interest rates on long-term determinants	Monetary and fiscal variables have a significant influence on the trend of long-term real interest rates
Canzoneri, Cumby and Diba (2002)	Semi-annual	US (1984-2002)	Regression of interest rates spreads on CBO budget surplus projections	Higher projected surpluses imply lower spreads of long-term rates over short-term rates.
Engen and Hubbard (2004)	Annual	US (1976-2003)	Regression of current real 10-year treasury rate on CBO 5-year ahead federal debt or deficit projections	Increases in the expected federal debt-to-GDP ratio increase the current real 10-year Treasury yield.
Heppke-Falk and Hüfner (2004)	Monthly	France, Germany, Italy (Jan:1994-Jul:2004)	SUR estimation	No significant impact of expected deficits on swap spreads over the whole sample.
Faini (2006)	Annual	EMU, except Luxembourg (1979-2002)	3SLS.	An expansionary fiscal policy in one EMU member will have an effect on its spreads, and on the overall level of interest rates for the currency union.
Laubach (2009)	Quarterly	US (1976:Q1-2006:Q2)	OLS. Regress expected future interest rates on CBO and OMB projections for the deficit-to-GDP ratio and the debt-to-GDP ratio 5 years ahead.	1 percentage point increase in the projected deficit ratio (debt ratio) raises long-term interest rates by roughly 25 (3 to 4) basis points.

3. Methodology

In the subsequent empirical analysis, an initial baseline specification for the real long-term government bond yield, *r*, can be written as

[1] Such results are in line with the Gale and Orszag (2003) assessment of the existence of statistically significant effects from anticipated budget deficits on long-term interest rates.

$$r_{it} = (i_{it} - \pi_{it}) = \alpha_i + \gamma_i X_{it} + u_{it} . \quad (1)$$

where i is the long-term nominal government bond yield, π is the inflation rate, and X includes a set of additional explanatory variables. The index i (i=1,…, N) denotes the country, the index t (t=1,…, T) indicates the period, α_i stands for the individual effects to be estimated for each country i, and u_{it} the disturbances.

An error-correction form for the real long-term interest rates, which move towards their long-run level with a speed of adjustment δ, is given by

$$\Delta(i_{it} - \pi_{it}) = \sum_{j=1}^{k} \beta_j \, \Delta(i_{it-j} - \pi_{it-j}) + \sum_{j=0}^{k} \theta_j \, \Delta X_{it-j} + \lambda_i \left[(i_{it-1} - \pi_{it-1}) - \alpha_i - \gamma_i \, X_{it-1} \right] + v_{it} , \quad (2)$$

where v_{it} are the disturbances.

Specification (1) illustrates a long-run relationship for the long-term real government bond yield. Among the several long-run factors influencing the long-term real interest rate that are included in X, we consider such determinants as: the government balance-to-GDP ratio, the debt-to-GDP ratio, the current account balance ratio, inflation surprises, the real effective exchange rate, and a liquidity measure.

As mentioned above, financial markets want to differentiate among sovereign debt issuers due to the existence of different country-specific credit risk and of a non-zero probability of sovereign default. Therefore, such variables as the government balance and the debt-to-GDP ratios could convey relevant information regarding a country credit risk and help in explaining cross-country financial risk premia. On the other hand, we do not want to expand too much the possible set of variables since we are aiming at a parsimonious empirical specification, while for the purposes of the subsequent error correction analysis it is also preferable not to have too may variables.

In addition, such fiscal indicators also allow financial markets to assess the fiscal future developments in sovereign borrowers and its perceived credit risk, the country's long-run solvency, and repayment likelihood. Therefore, relevant information regarding a country's debt burden and whether its public finance behaviour is sustainable, or if the risk for a build-up of government debt arises.[2] In other words, they help in gauging whether a country can make the interest payments on the outstanding stock of

[2] Afonso and Rault (2010) report that over the period 1970-2006 some EU countries may have been threading unsustainable public finances' paths.

government debt, without being necessarily forced into additional borrowing in the market and embarking in an unpleasant debt arithmetic trap.

Regarding inflation developments, inflation variability is also relevant in order for market participants to assess whether an environment of low inflation is in place, notably via the occurrence of inflation surprises. One can hypothesise that since with high inflation a government tends to unilaterally and partially inflate away from its fiscal indebtedness, the need for a higher nominal and real long-term bond yield cannot be discarded. Moreover, expected inflation is also seen as an indicator of macroeconomic stability, and higher inflation implies higher sovereign risk. Deviations from past inflation can be assumed from the actual inflation rate, or taken as an average of past observations.

In addition, the external imbalance of a country, for instance as proxied by the current account balance-to-GDP ratio, can convey the existence of a gap between saving and investment and provide expectations regarding a future depreciation of the domestic currency. Under those circumstances the risk premia demanded by the markets on sovereign debt may also increase. Moreover, external imbalances tend to be linked to fiscal imbalances from a long-term perspective, notably when private saving does not increase sufficiently to offset the effects of increased budget deficits, and then they may also impinge via such channel on long-term bond yields.[3] In addition, real effective exchange rate developments are linked to a country's foreign competitiveness while being also linked to current account balance positions.

Sovereign debt yields also tend to be related to the depth or liquidity of the respective outstanding bond market. Indeed, liquidity risk is usually inversely related to the size of the respective market. Therefore, it seems also useful to consider a measure of liquidity as a possible determinant of long-term government bond yields. Our liquidity measure, liquidity debt share, is given by the share of outstanding government debt in country i, in year t, in the overall outstanding government debt of the full set of countries in our sample:

$$LIQ_{it} = Debt_{it} / \sum_{i=1}^{N} Debt_{it} \qquad (3)$$

where the index i=1, …, N indicates the country.

[3] Afonso and Rault (2008) uncover significant effects between budget balances and current account balances for several OECD countries.

Naturally, one has to be aware that full liberalisation and integration of capital and bond markets was not in place for the entire time sample under analysis. Indeed, capital markets were gradually liberalised in the 1970s and 1980s. For instance, this was a mandatory requirement for EU countries at the start of stage two of EMU, in 1994. Another caveat is the fact that some home bias can arise among investors, for instance, some institutional investors may face constraints leading to portfolio investments in the home country. In a stepwise approach we then i) assess cross-country dependences; ii) test for panel unit roots; iii) estimate the panel cointegration relationships and iv) assess the respective magnitudes of cointegration.

Afterwards, and once we have estimated the long-run relationships between real long-term interest rates and their potential determinants via the computation of the common correlated effect CCE and CCE-MG (Mean Group) estimators (Pesaran, 2006), we also a estimate complete panel error-correction (PECM) models given by equation (2) with the Pooled Mean Group approach of Pesaran, Shin and Smith (1999). This framework allows us to assess the adjustment mechanism to a deviation from the long-run equilibrium relationship along with the short-run dynamics. Note that the CCE-MG estimator yields consistent estimates even in the presence of common factors and is the most efficient (Kapetanios and Pesaran, 2007) and robust to alternative hypotheses of non-stationarity of variables (Coakley et al., 2006).

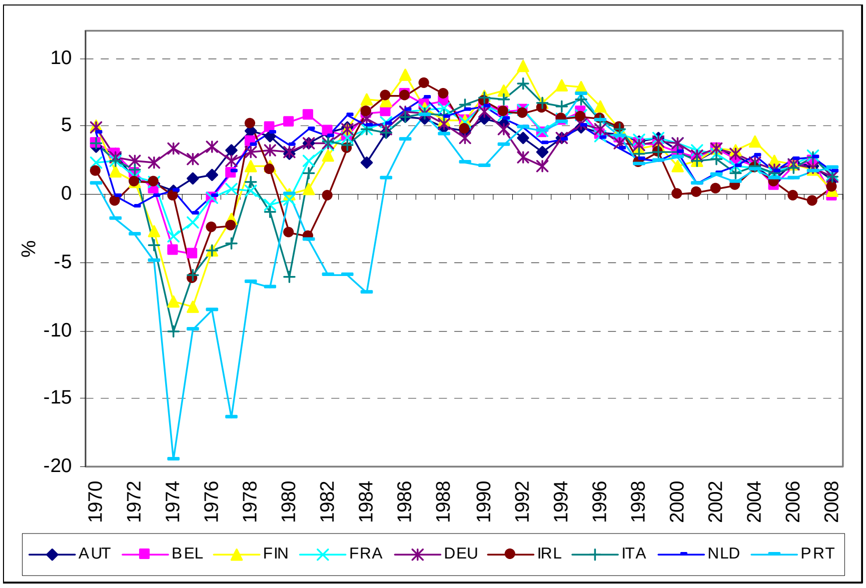
4. Empirical analysis

4.1. Data

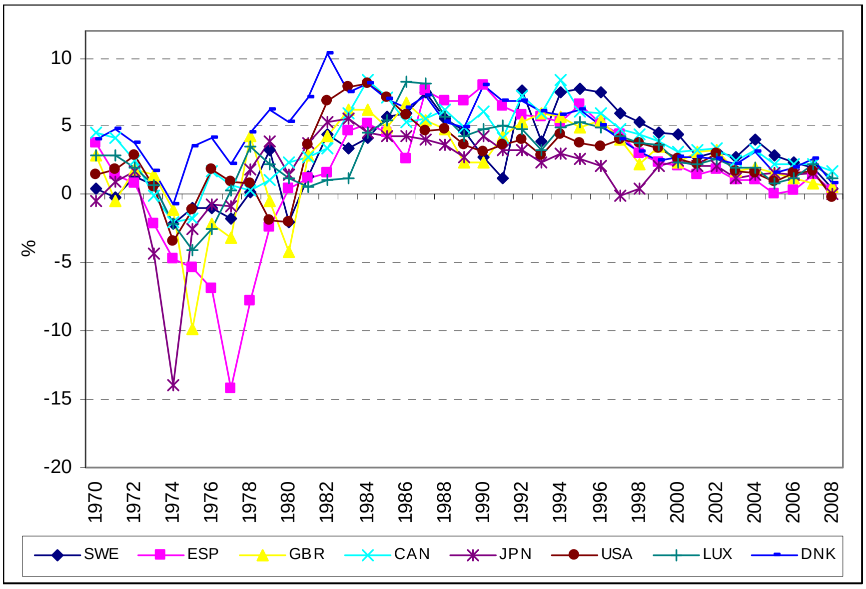
In our analysis we consider, for the period 1973-2008, the following set of 17 OECD countries: Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Luxembourg, Netherlands, Portugal, Sweden, Spain, UK, Canada, Japan, and U.S. Figure 1 illustrates the development of the long-term real interest rates for those countries.

Figure 1 – Long-term real interest rates

1a



1b



Source: IMF, International Financial Statistics, and authors' calculations.

From a simple visual inspection we can observe an upward movement in real long-term interest rates until the beginning of the 1980s, followed by a subsequent downward trend until the end of the time sample. Real long-term interest rates have been essentially positive apart from the period of the seventies and early eighties, when high inflation rates were also prevalent, particularly in such countries as Finland, Italy, Japan, Portugal, Spain, and the UK.

Regarding the liquidity measure that we computed following (3), Table 1 shows that the U.S. and Japan accounted in 2008 for more than half of the outstanding stock of sovereign debt in the set of OECD countries considered in our country sample.

We build inflation surprises (π^e) taking the difference between actual inflation and a 2-year moving average of past inflation (see the Appendix for data sources).

Table 1 – Shares of outstanding government debt in the total outstanding debt of the country sample

	1970	1980	1990	2000	2008
Austria	0.35	0.90	0.94	0.73	0.88
Belgium	2.02	2.90	2.59	1.44	1.54
Canada		3.84	4.45	3.42	3.27
Denmark		0.85	0.86	0.48	0.39
Finland	0.16	0.19	0.20	0.31	0.31
France	7.41	4.48	4.45	4.38	6.63
Germany	4.74	8.97	7.37	6.53	8.17
Ireland	0.26	0.46	0.45	0.21	0.40
Italy	5.10	8.18	10.90	6.90	8.31
Japan	3.06	18.28	21.19	36.70	28.83
Luxembourg	0.04	0.02	0.01	0.01	0.03
Netherlands	2.88	2.56	2.30	1.19	1.73
Portugal		0.30	0.42	0.33	0.55
Spain	0.73	1.16	2.26	1.98	2.16
Sweden	1.18	1.62	1.02	0.76	0.62
UK	12.13	8.93	3.42	3.49	4.71
US	59.93	36.36	37.16	31.15	31.45
	100.00	100.00	100.00	100.00	100.00

Source: European Commission AMECO database and authors' computations.

4.2. Cross-section dependence

In recent years it has become more widely recognized that the advantages of panel unit root tests within the macro-panel setting include the use of data for which the spans of individual time series data are insufficient for the study of many hypotheses of interest. The adoption of such new panel data methods is preferred to the usual time series techniques to circumvent the well known problems associated with the low power of traditional unit root tests. Therefore the body of literature on panel unit root and panel cointegration testing has grown considerably in the past ten years and now distinguishes between: first-generation tests (Maddala and Wu, 1999, Levin et al., 2002, and Im et al., 2003) developed on the assumption of the cross-sectional independence of panel units (except for common time effects), which is often unrealistic in many empirical settings; and second-generation tests (Bai and Ng, 2004, Smith et al., 2004, Moon and Perron, 2004, Choi, 2006, and Pesaran, 2007) allowing for a variety of dependence across the different units. These tests differ according to the way they eliminate the factors of structural dependence and the way they aggregate the individual information.[4]

Therefore, the first question to deal with is the possible presence of cross-section dependence in the data. Indeed, as put in evidence for instance, by O'Connell (1998) in the case of PPP testing, or by Banerjee et al. (2005), panel unit root tests of the first generation can lead to spurious results (because of size distortions) if there exists significant degrees of error cross-section dependence and this is ignored. Consequently, the implementation of second-generation panel unit root tests is desirable only when it has been established that the panel is effectively subject to a significant degree of error cross-section dependence. In the cases where cross-section dependence is not sufficiently high, loss of power might result if second-generation panel unit root tests that allow for cross-section dependence are used. Therefore, before an appropriate choice of a panel unit root test is made it is crucial to provide some evidence on the degree of residual cross-section dependence.

One way of testing for the presence of cross-section dependence in the data is to carry out the test of Pesaran (2004) and to compute the Cross section Dependence (CD) statistic. The test of Pesaran (2004) is based on a simple average of all pair-wise correlation coefficients of the OLS residuals (e_{it}) obtained from standard augmented Dickey-Fuller (1979) regressions for each individual in the panel. Denoting by $\hat{\rho}_{ij}$ the sample estimate of the pair-wise correlation coefficient for the residuals for countries i and j calculated over T periods, we get:

$$\hat{\rho}_{ij} = \hat{\rho}_{ji} = \left[\sum_{t=1}^{T} e_{it} e_{jt} \right] / \left[(\sum_{t=1}^{T} e_{it}^2)^{1/2} (\sum_{t=1}^{T} e_{jt}^2)^{1/2} \right]. \tag{4}$$

The test statistic proposed by Pesaran (2004), which does not depend on any particular spatial weight matrix when the cross-sectional dimension (N) is large, is given by

[4] Note that a specific form of cross-sectional dependence that has become popular is the factor structure approach. This has been used extensively in empirical work (see, for instance, Barro and Sala-i-Martin, 1992) and it has been analysed in theoretical treatments at even greater length. Therefore, in our study we use the notions of error cross-sectional dependence and factor structure dependence interchangeably.

$$CD = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{t=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{ij} \right), \quad (5)$$

and under its null hypothesis of cross-sectional independence it has asymptotically a standard normal distribution. The results reported in Table 2 provide evidence in favour of the existence of cross-sectional dependence in the data since for all series the CD statistics are always highly significant whatever the number of lags (from 1 to 4) included in the ADF regressions. In other words, one rejects the null hypothesis of cross-section independence

Table 2 – Cross-section correlations of the errors in the ADF(p) regressions of real long-term interest rates and potential determinants (1973-2008; N = 17)#

	Real Long-Term Interest Rate (R)				Government Balance Ratio (GBR)			
Test Statistic	p=1	p=2	p=3	p=4	p=1	p=2	p=3	p=4
CD	12.21	12.02	11.85	11.54	23.12	22.58	21.45	21.36
P value	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
	Inflation Surprises (Π^e)				Current Account Balance Ratio (CA)			
Test Statistic	p=1	p=2	p=3	p=4	p=1	p=2	p=3	p=4
CD	20.24	17.99	17.78	17.10	32.79	30.47	31.56	32.15
P value	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
	Liquidity Debt Share (LIQ)				Real Effective Exchange Rate (TCR)			
Test Statistic	p=1	p=2	p=3	p=4	p=1	p=2	p=3	p=4
CD	22.32	22.15	21.85	20.76	19.25	18.45	18.35	17.41
P value	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
	Debt Ratio (DR)							
Test Statistic	p=1	p=2	p=3	p=4				
CD	27.12	26.58	25.12	25.16				
P value	(0.00)	(0.00)	(0.00)	(0.00)				

Note: Under the null of cross-sectional independence the CD statistic is distributed as a two-tailed standard normal. # Results based on the test of Pesaran (2004).
The variable Inflation Surprises is calculated for each country as the difference between actual inflation and a moving average of two periods.

4.3. Panel unit root testing

Having put in evidence the presence of cross section dependence in real long-term interest rates, we now turn to the determination of the degree of integration of the series (real long-term interest rate, government balance ratio, current account balance, inflation surprises, real effective exchange rate, liquidity debt share, debt ratio) in our panel of 17 countries, using two second-generation panel unit root tests.

The first 2nd generation unit root test that we use is the test by Pesaran (2007) who suggests a simple way of getting rid of cross-sectional dependence that does not require the estimation of factor loading. His method is based on augmenting the usual ADF regression with the lagged cross-sectional mean and its first difference to capture the cross-sectional dependence that arises through a single-factor model. The resulting individual ADF test statistics (CADF) or the rejection probabilities can then be used to develop modified versions of the t-bar test proposed by Im et al. (2003), such as the Cross-sectionally augmented IPS ($CIPS = N^{-1}\sum_{i=1}^{N} CADF_i$), or a truncated version of the CIPS statistic (CIPS*) where the individual CADF statistics are suitably truncated to avoid undue influences of extreme outcomes that could arise when T is small (between 10 and 20), or the inverse normal test (or the Z test) suggested by Choi (2001) that combine the p-values of the individual tests (CZ). Critical values reported in Pesaran (2007) are provided through Monte Carlo simulations for a specific specification of the deterministic component and depend both on the cross-sectional and time series dimensions. The null hypothesis of all tests is the unit root.

The second set of unit root tests of the 2nd generation are the bootstrap tests of Smith et al. (2004), which use a sieve sampling scheme to account for both the time series and cross-sectional dependencies of the data through bootstrap blocks. The specific tests that we consider are denoted $\bar{t}$, $\overline{LM}$, $\overline{\max}$, and $\overline{\min}$. $\bar{t}$ is the bootstrap version of the well known panel unit root test of Im et al. (2003), $\overline{LM} = N^{-1}\sum_{i=1}^{N} LM_i$ is a mean of the individual Lagrange Multiplier (LM_i) test statistics, originally introduced by Solo (1984), $\overline{\max}$ is the test of Leybourne (1995), and $\overline{\min} = N^{-1}\sum_{i=1}^{N} min_i$ is a (more powerful) variant of the individual Lagrange Multiplier (LM_i), with $\min_i = \min(LM_{fi}, LM_{ri})$, where LM_{fi} and LM_{ri} are based on forward and backward regressions (see Smith et al., 2004 for further details). We use bootstrap blocks of m=20.[5] All four tests are constructed with a unit root under the null hypothesis and heterogeneous autoregressive roots under the alternative, which indicates that a rejection should be taken as evidence in favour of stationarity for at least one country.

The results of the second generation panel unit root tests proposed by Pesaran (2007) are reported in Table 3 and provide support of the existence of a unit root in all series under consideration. This conclusion, which is robust to the number of lags introduced in the ADF regressions (from p=1 to 4), should be considered as safe given

[5] The results are not very sensitive to the size of the bootstrap blocks.

the large and significant degree of cross-section dependence in all series documented in Table 2.

Table 3 – Panel unit root tests of Pesaran (2007) for real long-term interest rates and potential determinants (1973-2008; N = 17)

	Real Long-Term Interest Rate (R)				Government Balance Ratio (GBR)			
Test Statistics	p=1	p=2	p=3	p=4	p=1	p=2	p=3	p=4
CIPS	-1.92	-188	-1.84	-1.96	-2.10	-1.96	-1.72	-1.68
CIPS*	-1.58	-1.52	-1.46	-1.62	-2.09	-1.95	-1.71	-1.68
	Inflation Surprises (Π^e)				Current Account Balance Ratio (CA)			
Test Statistic	p=1	p=2	p=3	p=4	p=1	p=2	p=3	p=4
CIPS	-2.09	-2.07	-1.98	-1.91	-2.20*	-1.90	-1.43	-1.28
CIPS*	-2.08	-2.06	-1.98	-1.91	-2.19*	-1.89	-1.43	-1.28
	Liquidity Debt Share (LIQ)				Real Effective Exchange Rate (TCR)			
Test Statistic	p=1	p=2	p=3	p=4	p=1	p=2	p=3	p=4
CIPS	-1.95	-1.92	-1.90	-2.01	-2.03	-1.99	-1.97	-1.94
CIPS*	-1.93	-1.91	-1.89	-2.01	-2.02	-1.98	-1.97	-1.94
	Debt Ratio (DR)							
Test Statistic	p=1	p=2	p=3	p=4				
CIPS	-1.75	-1.68	-1.82	-1.78				
CIPS*	-1.74	-1.65	-1.82	-1.77				

Notes: 1) A constant is included in the estimations.
2) Rejection of the null hypothesis indicates stationarity at least in one country.
3) Critical values are respectively of -2.40 at 1%, -2.22 at 5%, and -2.14 at 10%.
* denotes rejection of the null at the 10 % significance level.
CIPS – Cross-section augmented Im-Pesaran-Shin test. CIPS* – truncated CIPS test.

Similar results in Table 4, suggest that for all the series the unit root null cannot be rejected at any conventional significance level by the four bootstrap tests of Smith et al (2004).[6] Therefore, we conclude that real long-term interest rates and their potential determinants (government balance ratio, current account balance ratio, inflation surprisess, real effective exchange rate, liquidity debt share, and government debt ratio) are non-stationary and integrated of order one at the five percent level of significance in our country panel. [7]

[6] The order of the sieve is allowed to increase with the number of time series observations at the rate $T^{1/3}$ while the lag length of the individual unit root test regressions are determined using the Campbell and Perron (1991) procedure. Each test regression is fitted with a constant term only.
[7] The lag order in the individual ADF type regressions is selected for each series using the AIC model selection criterion. Another crucial issue is the selection of the order of the deterministic component. In particular, since the cross-sectional dimension is rather large here, it may seem restrictive not to allow at least some of the units to be trending, suggesting that the model should be fitted with both a constant and trend. However, since the trending turned out not to be very pronounced, we have considered that a constant is enough in our analysis. Actually, the results of the bootstrap tests of Smith et al. (2004) are not very sensitive to the inclusion of a trend in addition to a constant in the estimated equation (see Statistic b in Table 4). We have of course also checked using the tests by Pesaran (2007) and the bootstrap tests of Smith et al. (2004) that the first difference of the series are stationary, hence confirming that the series expressed in level are integrated of order one.

Table 4 – Panel unit root tests of Smith et al. (2004) for real long-term interest rates and potential determinants (1973-2008)*

	Real Long-Term Interest Rate (R)				Government Balance Ratio (GBR)			
Test	Statistic (a)	Bootstrap P-value*	Statistic (b)	Bootstrap P-value*	Statistic (a)	Bootstrap P-value*	Statistic (b)	Bootstrap P-value*
$\bar{t}$	-1.528	0.325	-2.108	0.555	-1.590	0.275	-2.836	0.424
$\overline{LM}$	2.309	0.187	5.040	0.554	2.478	0.128	5.759	0.391
$\overline{\max}$	-0.812	0.264	-1.526	0.807	-1.308	0.142	-1.419	0.128
$\overline{\min}$	1.301	0.357	3.061	0.824	1.659	0.165	4.188	0.168
	Inflation Surprises (Π^e)				Current Account Balance Ratio (CA)			
Test	Statistic (a)	Bootstrap P-value*	Statistic (b)	Bootstrap P-value*	Statistic (a)	Bootstrap P-value*	Statistic (b)	Bootstrap P-value*
$\bar{t}$	-1.888	0.101	-1.893	0.754	-1.454	0.536	-2.218	0.330
$\overline{LM}$	4.375	0.122	4.640	0.742	3.551	0.288	5.609	0.354
$\overline{\max}$	-0.825	0.686	-1.560	0.701	-1.366	0.065	-1.955	0.178
$\overline{\min}$	1.652	0.633	3.366	0.731	3.200	0.057	4.568	0.218
	Liquidity Debt Share (LIQ)				Real Effective Exchange Rate (TCR)			
Test	Statistic (a)	Bootstrap P-value*	Statistic (b)	Bootstrap P-value*	Statistic (a)	Bootstrap P-value*	Statistic (b)	Bootstrap P-value*
$\bar{t}$	-1.965	0.133	-2.207	0.333	-1.815	0.131	-2.448	0.157
$\overline{LM}$	2.986	0.105	5.676	0.312	4.272	0.194	6.193	0.176
$\overline{\max}$	-1.325	0.114	-1.538	0.712	-1.293	0.134	-1.772	0.189
$\overline{\min}$	2.297	0.333	3.183	0.724	2.443	0.125	4.275	0.182
	Debt Ratio (DR)							
Test	Statistic (a)	Bootstrap P-value*	Statistic (b)	Bootstrap P-value*				
$\bar{t}$	-1.570	0.413	-2.216	0.419				
$\overline{LM}$	2.649	0.659	4.949	0.510				
$\overline{\max}$	-1.541	0.087	-1.720	0.652				
$\overline{\min}$	2.584	0.105	3.217	0.780				

Notes: (a) Model includes a constant. (b) Model includes both a constant and a time trend.
* Test based on Smith et al. (2004). Rejection of the null hypothesis indicates stationarity at least in one country. All tests are based on 5,000 bootstrap replications to compute the p-values.
Null hypothesis: unit root (heterogeneous roots under the alternative).

4.3. Panel cointegration

Given that all the series under investigation are integrated of order one, we now proceed with the two following steps. First, we perform 2nd generation panel data cointegration tests (that allow for cross-sectional dependence among countries) to test for the existence of cointegration between real long-term interest rates and its potential determinants. Second, if a cointegrating relationship exists for all countries, we estimate for each country the cross-section augmented cointegrating regression

$$r_{it} = (i_{it} - \pi_{it}) = \alpha_i + \gamma_i X_{it} + \mu_1 \bar{r}_t + \mu_2 \overline{X}_t + u_{it}, \quad i = 1,...,N; \quad t = 1,...,T \quad (6)$$

by the CCE estimation procedure proposed by Pesaran (2006) that allows for cross-section dependencies that potentially arise from multiple unobserved common factors. The cointegrating regression is augmented with the cross-section averages of the dependent variable and the observed regressors as proxies for the unobserved factors. Accordingly, $\bar{r}_t$ and $\bar{X}_t$ denote respectively the cross-section averages of r_i and X_i in year t. Note that the coefficients of the cross–sectional means (CSMs) do not need to have any economic meaning as their inclusion simply aims to improve the estimates of the coefficients of interest. Therefore, this procedure enables us to estimate the individual coefficients γ_i in a panel framework.[8]

In addition, we also compute the CCE-MG estimators of Pesaran (2006). For instance, for the γ parameter and its standard error for N cross-sectional units, they are easily obtained as follows: $\hat{\gamma}_{CCE-MG} = \frac{\sum_{i=1}^{N} \hat{\gamma}_{i-CCE}}{N}$, and $SE(\hat{\gamma}_{CCE-MG}) = \frac{\sum_{i=1}^{N} \sigma(\hat{\gamma}_{i-CCE})}{\sqrt{N}}$, where $\hat{\gamma}_{i-CCE}$ and $\sigma(\hat{\gamma}_{i-CCE})$ denote respectively the estimated individual country time-series coefficients and their standard deviations.

We now use the bootstrap panel cointegration test proposed by Westerlund and Edgerton (2007). This test relies on the popular Lagrange multiplier test of McCoskey and Kao (1998), and makes it possible to accommodate correlation both within and between the individual cross-sectional units. In addition, this bootstrap test is based on the sieve-sampling scheme, and has the advantage of significantly reducing the distortions of the asymptotic test. Another appealing advantage is that the joint null hypothesis is that all countries in the panel are cointegrated. Therefore, in case of non-rejection of the null, we can assume that there is cointegration between real long-term interest rates and the potential determinants contained in *X*. In what follows we consider the following sets of variables included in *X*, which cover the main relevant economic determinants:

[8] Note that in order to estimate the long-run coefficients we have also implemented the Pooled Mean Group (PMG) estimators (see Pesaran and Smith (1995), Pesaran, Shin and Smith (1999)), which allowed us to identify significant differences in country behaviour. However, we only report the results of the Common Correlated Effects (CCE) estimators developed by Pesaran (2006), since they allow taking unobservable factors into account, which would not be the case of the PMG estimators.

i) $X_1 = (\Pi^e, CA, DR)$,

ii) $X_2 = (\Pi^e, CA, GBR)$,

iii) $X_3 = (\Pi^e, CA, DR, GBR, TCR)$,

iv) $X_4 = (\Pi^e, CA, DR, LIQ)$.

The panel cointegration results from the asymptotic tests shown in Table 5, including a constant term, indicate the absence of a cointegrating relationship between real long-term interest rates and the different sets of potential determinants for our country panel. However, this result is based on conventional asymptotic critical values, calculated on the assumption of cross-sectional independence of countries, an assumption that is not true here given the significant cross-sectional correlation among the series documented previously (in Table 2).

Table 5 – Panel cointegration between real long-term interest rates and different sets of potential determinants (1973-2008; N = 17), model with a constant term

	LM-stat	Asymptotic p-value	Bootstrap p-value #
$X_1 = (\Pi^e, CA, DR)$	7.430	0.000	0.840
$X_2 = (\Pi^e, CA, GBR)$	7.385	0.000	0.782
$X_3 = (\Pi^e, CA, DR, GBR, TCR)$	14.168	0.000	0.783
$X_4 = (\Pi^e, CA, DR, LIQ)$	9.125	0.000	0.751

Notes: the bootstrap is based on 2000 replications.
a - The null hypothesis of the tests is cointegration of Real Long-Term Interest Rates and potential determinants series.
\# Test based on Westerlund and Edgerton (2007).

Therefore, given the existence of some cross-section dependence among individuals, we used bootstrap critical values.[9] In this case the conclusions of the tests are now more compelling, and retaining a 10% level of significance, we conclude that there is a long-run relationship between real long-term interest rates and most of the different sets of potential determinants for our panel of OECD countries. This implies in particular that over the longer run real long-term interest rates and their determinants move together in our OECD sample. In addition, Table 5 implies that strictly relying upon asymptotic critical values (i.e. neglecting cross-sectional dependence) may lead to wrong (opposite) conclusions about the macroeconomic and fiscal long-run links between real long-term interest rates and their potential determinants.

[9] As pointed out by a referee "provided that the bootstrap method is appropriate for the problem and implemented correctly, then the bootstrap critical values will be appropriate also in the absence of cross-sectional correlation: they would just be closer to the asymptotic ones."

4.5. The magnitudes of the cointegration relationship

We then estimate equation (6) for the four different sets of variables included in *X* to assess the magnitude of the individual γ_i coefficient in the cointegrating relationship with the CCE estimation procedure developed by Pesaran (2006), which addresses cross-sectional dependency. The estimated equations are

$$r_{it} = \alpha_i + \gamma_{1i}\Pi^e{}_{it} + \gamma_{2i}CA_{it} + \gamma_{3i}DR_{it} + u_{it}\,, \tag{6a}$$

$$r_{it} = \alpha_i + \gamma_{1i}\Pi^e{}_{it} + \gamma_{2i}CA_{it} + \gamma_{3i}GBR_{it} + u_{it}\,, \tag{6b}$$

$$r_{it} = \alpha_i + \gamma_{1i}\Pi^e{}_{it} + \gamma_{2i}CA_{it} + \gamma_{3i}DR_{it} + \gamma_{4i}GBR_{it} + \gamma_{5i}TCR_{it} + u_{it}\,, \tag{6c}$$

$$r_{it} = \alpha_i + \gamma_{1i}\Pi^e{}_{it} + \gamma_{2i}CA_{it} + \gamma_{3i}DR_{it} + \gamma_{4i}LIQ_{it} + u_{it}\,, \tag{6d}$$

with $i = 1,...,N,\ t = 1,...,T$, and the respective estimation results are reported in Table 6.

Table 6a – Individual country CCE estimates for 17 OECD countries (1973-2008) between real long-term interest rates and the X_1= (Π^e, CA, DR) determinants

Country	Π^e		CA		DR		Constant	
	γ_1	t-Stat	γ_2	t-Stat	γ_3	t-Stat	α	t-Stat
Austria	-0.681	-8.012	-0.149	-3.725	-0.011	-3.611	3.962	2.907
Belgium	-0.850	-14.167	-0.045	-2.682	-0.024	-2.600	-0.281	-2.257
Canada	-0.885	-9.725	-0.133	-2.509	-0.032	-2.000	-0.978	-2.736
Denmark	-0.645	-4.778	-0.040	-2.727	0.028	2.000	-1.497	-3.749
Finland	-0.515	-4.769	-0.006	-0.120	-0.045	-1.957	2.222	3.467
France	-0.766	-12.355	-0.176	-3.520	0.049	5.444	-2.234	-3.659
Germany	-0.875	-4.581	0.262	2.148	-0.005	-2.167	-4.331	-4.746
Ireland	-0.726	-8.643	-0.022	-2.759	0.023	2.300	1.573	2.637
Italy	-0.690	-16.429	0.183	2.346	0.121	4.321	-5.388	-3.456
Japan	-0.944	-17.164	-0.109	-2.652	-0.047	-2.611	-3.396	-2.741
Luxembourg	-1.045	-40.192	0.003	0.029	-0.038	-1.267	-3.200	-2.839
Netherlands	-0.779	-10.819	-0.134	-1.523	0.017	3.850	-0.079	-2.034
Portugal	-0.803	-13.164	0.114	3.563	0.001	0.000	-0.995	-3.531
Spain	-0.875	-13.258	0.058	2.784	-0.007	-1.400	5.189	6.163
Sweden	-0.932	-22.732	-0.045	-0.517	-0.140	-5.385	4.017	3.088
UK	-0.806	-14.140	0.051	2.125	0.081	3.375	0.267	2.286
US	-0.407	-2.928	0.010	2.192	0.020	2.176	3.051	5.202

Note the coefficients of the variables $\bar{r}_t$ *and* $\bar{X}_{1t}$ of equation (6a) have not been reported in the table.

Table 6b – Individual country CCE estimates for 17 OECD countries (1973-2008) between real long-term interest rates and the X_2= (Π^e, CA, GBR) determinants

Country	Π^e		CA		GBR		Constant	
	γ_1	t-Stat	γ_2	t-Stat	γ_3	t-Stat	α	t-Stat
Austria	-0.668	-9.408	-0.147	-3.675	0.027	2.692	2.397	3.405
Belgium	-0.796	-12.438	-0.142	-2.407	0.015	2.313	1.109	3.391
Canada	-0.821	-8.642	-0.151	-3.283	0.076	2.854	0.680	2.932
Denmark	-0.471	-3.680	-0.008	-2.178	-0.357	-1.812	-0.728	-2.874
Finland	-0.399	-5.182	-0.039	-2.345	-0.200	-2.597	-0.231	-2.486
France	-0.892	-12.389	-0.045	-2.776	-0.100	-2.099	-1.499	-1.669
Germany	-0.992	-6.161	0.137	3.593	0.128	2.422	0.461	2.645
Ireland	-0.699	-8.034	-0.022	-0.564	0.058	2.289	1.909	3.094
Italy	-0.650	-13.000	0.142	2.958	-0.462	-4.200	0.021	2.017
Japan	-0.912	-22.244	-0.081	-1.306	0.114	3.081	0.178	3.231
Luxembourg	-1.021	-30.029	-0.047	-0.758	0.007	0.092	-2.851	-2.589
Netherlands	-0.762	-12.915	-0.028	-0.431	-0.179	-2.210	-0.322	-2.643
Portugal	-0.909	-16.527	0.048	1.455	0.199	4.854	0.234	2.600
Spain	-0.982	-9.627	-0.146	-2.168	0.175	2.869	0.421	2.636
Sweden	-0.990	-14.559	-0.397	-3.970	0.278	3.159	-0.989	-1.169
UK	-0.767	-11.984	-0.004	-0.148	0.068	2.194	0.221	2.795
US	-0.341	-3.217	-0.132	-2.859	-0.118	-2.532	0.868	2.018

Note the coefficients of the variables $\bar{r}_t$ and $\overline{X}_{1t}$ of equation (6b) have not been reported in the table.

Table 6c – Individual country CCE estimates for 17 OECD countries (1973-2008) between real long-term interest rates and the X_3= (Π^e, CA, DR, GBR, TCR) determinants

Country	Π^e		CA		DR		GBR		TCR		Constant	
	γ_1	t-Stat	γ_2	t-Stat	γ_3	t-Stat	γ_4	t-Stat	γ_5	t-Stat	α	t-Stat
Austria	-0.776	-10.778	-0.191	-4.244	-0.015	-2.750	0.065	2.121	-0.030	-2.714	-7.861	-3.551
Belgium	-1.076	-13.450	0.050	2.042	-0.039	-4.333	0.021	1.583	-0.069	-4.600	8.764	4.382
Canada	-0.851	-5.420	-0.038	-2.369	-0.005	-0.167	-0.058	-2.487	0.063	2.969	-2.404	-4.409
Denmark	-0.706	-4.903	-0.087	-2.526	0.020	1.909	-0.166	-2.107	-0.022	-2.917	12.695	5.577
Finland	-0.423	-5.494	-0.031	-0.633	-0.039	-2.786	-0.224	-3.733	0.006	0.375	-8.697	-2.077
France	-0.745	-10.643	-0.252	-3.360	0.056	3.294	0.011	0.116	-0.047	-2.136	-9.387	-6.380
Germany	-1.185	-5.925	0.127	2.221	-0.027	-0.675	0.147	2.909	-0.051	-2.244	-9.621	-7.912
Ireland	-0.780	-5.612	-0.057	-2.727	-0.002	-0.105	0.075	2.415	-0.065	-2.140	-3.507	-4.634
Italy	-0.733	-16.289	0.178	2.507	0.110	3.929	-0.227	-2.610	-0.115	-2.018	10.527	6.426
Japan	-0.917	-17.635	-0.139	-2.106	-0.010	-0.455	0.072	2.323	-0.015	-1.000	-9.007	-5.795
Luxembourg	-1.049	-38.852	0.059	1.656	-0.046	-2.769	-0.038	-0.413	0.014	2.560	-10.050	-2.012
Netherlands	-0.827	-15.315	0.136	2.333	-0.021	-2.050	-0.044	-0.647	-0.062	-6.889	0.454	2.054
Portugal	-0.927	-22.071	0.006	0.109	-0.012	-2.200	0.165	3.667	-0.010	-1.000	1.843	2.376
Spain	-0.961	-21.356	-0.114	-2.107	-0.009	-2.250	0.070	2.258	-0.021	-2.750	1.798	2.271
Sweden	-0.884	-24.556	-0.158	-2.179	-0.045	-1.818	0.104	2.852	0.026	2.625	3.535	3.399
UK	-0.832	-21.333	0.009	0.409	0.101	4.391	0.106	4.417	-0.062	-3.444	-2.073	-2.726
US	-0.435	-5.000	-0.066	-2.245	-0.041	-2.864	-0.150	-2.705	0.034	3.919	21.341	4.109

Note the coefficients of the variables $\bar{r}_t$ and $\overline{X}_{1t}$ of equation (6c) have not been reported in the table.

Table 6d – Individual country CCE estimates for 17 OECD countries (1973-2008) between real long-term interest rates and the X_4= (Π^e, CA, DR, LIQ) determinants

	Π^e		CA		DR		LIQS		Constant	
Country	γ_1	t-Stat	γ_2	t-Stat	γ_3	t-Stat	γ_4	t-Stat	α	t-Stat
Austria	-0.647	-7.890	-0.091	-2.459	0.008	2.400	-0.439	-3.663	-28.650	-2.304
Belgium	-1.063	-18.982	0.145	2.843	-0.031	-2.214	-0.900	-3.180	24.210	2.907
Canada	-0.918	-9.180	-0.139	-2.725	-0.009	-2.600	-5.752	-5.321	-10.417	-0.576
Denmark	-0.707	-4.842	0.090	2.250	0.053	3.359	-0.218	-2.172	41.473	4.401
Finland	-0.536	-6.700	-0.026	-2.667	-0.025	-2.087	-0.263	-3.697	-5.988	-2.528
France	-0.759	-13.316	-0.215	-4.778	0.052	3.250	-1.751	-2.447	-47.188	-4.047
Germany	-0.899	-5.197	0.191	1.201	0.032	2.744	-0.271	-2.362	-14.353	-5.467
Ireland	-0.729	-8.679	0.022	0.786	0.010	2.500	0.420	3.719	-12.185	-3.954
Italy	-0.755	-15.729	0.185	2.569	0.222	5.286	-12.00	-3.087	-36.057	-6.857
Japan	-0.914	-12.694	-0.056	-2.824	-0.042	-3.680	-0.310	-2.314	-26.735	-7.756
Luxembourg	-1.039	-30.559	0.050	0.538	-0.059	-2.458	1.052	3.930	-1.959	-2.053
Netherlands	-0.826	-14.000	-0.126	-2.636	0.015	2.556	-0.159	-1.924	-58.738	-4.485
Portugal	-0.802	-12.935	0.086	2.867	-0.001	-2.100	-1.460	-2.209	16.218	2.074
Spain	-0.950	-20.652	-0.088	-1.846	0.002	2.000	-0.100	-2.632	22.840	3.579
Sweden	-0.846	-15.107	0.006	1.067	-0.149	-5.960	0.059	2.458	34.877	3.864
UK	-0.820	-14.643	0.052	2.080	0.101	2.629	-12.78	-6.423	8.191	5.334
US	-0.522	-3.896	0.030	2.698	-0.024	-3.500	5.151	5.575	-12.222	-3.207

Note the coefficients of the variables $\bar{r}_t$ *and* $\bar{X}_{1t}$ of equation (6d) have not been reported in the table.

From Table 6a we can observe that real long-term interest rates are statistically and positively affected by changes in the debt-to-GDP ratio for seven out of 17 countries. Regarding inflation surprises they have a negative and statistically significant effect on real long-term interest rates in all countries. In addition, the effect of the external imbalances is statistically significant and negative (positive) for nine (five) countries. In other words, the deterioration of the current account balance would signal mostly a widening gap between savings and investment and long-term interest rates may be pushed upwards.

The results of an alternative specification are reported in Table 6b where the debt-ratio is replaced by the government budget balance ratio. In this case, a better (more positive) government budget balance reduces the real long-term interest rate in six countries

In Table 6c, the CCE estimations include simultaneously the two fiscal determinants of real long-term interest rates, the government budget balance ratio and the debt-to-GDP ratio, together with current account balances and the real effective exchange rate. According to the results, improvements in the government budget balance reduce the real long-term interest rate in seven countries (five in a statistically significant way). The real effective exchange rate has a statistically significant negative effect in ten countries, with a depreciation reducing real long-term interest rates in the cointegration relationship.

Concerning the relevance of the liquidity of the outstanding government debt, defined in (3), as a determinant of long-term government bond yields, the related results are reported in Table 6d, considering the debt-to-GDP ratio as a determinant as well The effect of an increased country specific sovereign liquidity in the government debt market contributes to reduce long-term interest rates in 13 cases. In addition, we can observe that inflation surprises still have a statistically significant negative effect on real long-term interest rates in all countries, while higher debt ratios also imply higher long-term interest rates for nine countries, and current account deteriorations push up real interest rates in ten cases.

Finally, the results from the common correlated effects mean group (CCE-MG) method are reported in Table 7. We can see that the estimated long-run relationships for the real long-term interest rates confirm the statistical relevance of inflation, current account balances, budgetary balances, government debt and of the liquidity proxy.

Table 7 – Results for common correlated effects mean group (CCE-MG) estimations, 17 OECD countries (1973-2008)

	(6a) X_1= (Π^e, CA, DR)	(6b) X_2= (Π^e, CA, GBR)	(6c) X_3= (Π^e, CA, DR, GBR, TCR)	(6d) X_4= (Π^e, CA, DR, LIQ)
Constant	-0.123	0.110	-0.091	-6.216
	(-4.15)	(3.96)	(-5.26)	(-5.28)
Π^e	-0.777	-0.761	-0.829	-0.807
	(-20.19)	(-15.10)	(-17.25)	(-21.72)
CA	-0.010	-0.062	-0.030	-0.008
	(-3.96)	(-5.25)	(-4.32)	(-4.28)
DR	-0.060		-0.137	-0.009
	(-3.27)		(-6.25)	(-3.36)
GBR		-0.015	-0.041	
		(-3.34)	(-2.48)	
TCR			-0.026	
			(-3.98)	
LIQS				-1.749
				(-5.35)

Note: t-statistics are in parentheses.

4.6. Estimation of a panel ECM representation

In the previous sub-section we have estimated the long-run relationships between real long-term interest rates and their potential determinants for our panel of 17 OECD countries, using the common correlated effects mean group (CCE-MG) estimates (see Table 7). Having established the long-run structure of the underlying data and given that there exists a long-run relationship for all countries in our four panel sets, we turn to the estimation of the complete panel error-correction model (PECM) described by equation (7):

$$\Delta(i_{it}-\pi_{it})=\sum_{j=1}^{p}\beta_j\,\Delta(i_{it-j}-\pi_{it-j})+\sum_{j=0}^{p}\theta_j\,\Delta X_{it-j}+\lambda_i\left[(i_{it-1}-\pi_{it-1})-\alpha-\gamma\,X_{it-1}\right]+\varepsilon_{it}\,.(7)$$

We use the Pooled Mean Group (PMG) approach of Pesaran, Shin and Smith (1999), with long-run parameters obtained with CCE techniques, in order to obtain the estimates of the loading factors λ_i (weights or error correction parameters, or speed of adjustment to the equilibrium values), as well as of the short-run parameters β_j and θ_j for each country of our panel. Consequently, the loading factors and short-run coefficients are allowed to differ across countries.[10]

The lag length structure p is chosen using the Schwarz (SC) and Hannan-Quinn (HQ) selection criteria, and by carrying out a standard likelihood ratio testing-down type procedure to examine the lag significance from a long-lag structure (started with $p=4$) to a more parsimonious one. Afterwards, in order to improve the statistical specification of the model, we implemented systematically Wald tests of exclusion of lagged variables from the short-run dynamic (they are not reported here) to eliminate insignificant short-run estimates at the 5% level. We tested the residuals from each PECM model for the absence of heteroscedasticity, autocorrelation, ARCH effect, and we can report that they are not subject to misspecification. The results of the PECM estimations based on (7) for the different sets of potential determinants previously considered are reported in Table 8, only for significant short-run estimates at the 5% level.

[10] Note that before considering equation (7), we first used a Wald statistic to test for common parameters across countries (i.e $\lambda_i=\lambda$, and $\gamma_i=\gamma$, for i=1,...,N) with the CCE techniques of Pesaran, (2006), that allow common factors in the cross-equation covariances to be removed. We found that only the null hypothesis $\gamma_i=\gamma$, for i=1,…,N was not rejected by data, whereas the speeds of adjustment λ_i vary considerably across countries (results are available upon request).

Table 8a – Panel Error-Correction estimations for r_{it}, X_1= (Π^e, CA, DR), 1973-2008

	Δr_{it-1}	Δr_{it-2}	$\Delta \Pi^e_{it}$	$\Delta \Pi^e_{it-1}$	ΔCA_{it}	ΔCA_{it-1}	ΔCA_{it-2}	ΔDR_{it}	ΔDR_{it-1}	Loading factor λ_i
Austria		-0.20 (-2.21)	-0.69 (-7.79)	0.24 (2.91)					0.10 (2.08)	-0.06 (-6.31)
Belgium			-0.69 (-9.34)		-0.05 (-2.25)			0.06 (2.05)		-0.14 (-3.54)
Canada			-0.82 (-11.3)		-0.03 (-2.78)			0.04 (2.64)		-0.10 (-3.36)
Denmark			-0.62 (-6.31)		-0.34 (-1.99)					-0.08 (-2.91)
Finland			-0.51 (-4.35)							-0.10 (-3.12)
France			-0.97 (-4.12)	-0.41 (-2.62)	0.73 (3.94)	0.79 (5.11)		0.04 (4.42)	-0.35 (-3.48)	-0.61 (-4.13)
Germany			-0.97 (-11.7)	0.43 (3.20)	-0.05 (-2.45)	-0.38 (-2.79)		0.06 (2.37)	0.39 (3.37)	-0.14 (-2.95)
Ireland			-0.48 (-4.85)		-0.09 (-2.79)			0.01 (2.58)	0.08 (2.15)	-0.29 (-3.34)
Italy			-0.84 (-22.4)		-0.04 (-3.29)			0.05 (3.31)	-0.14 (-3.37)	-0.14 (-4.83)
Japan			-0.76 (-11.2)		-0.09 (-3.34)	0.27 (2.78)		0.01 (3.19)		-0.29 (-5.36)
Luxembourg			-1.07 (-11.4)		-0.05 (-2.96)			0.06 (2.90)		-0.15 (-3.95)
Netherlands			-0.85 (-16.8)		-0.03 (-2.05)		-0.39 (-3.22)	0.04 (2.17)		-0.09 (-2.38)
Portugal			-0.69 (-9.02)		-0.06 (-2.72)			0.008 (2.56)		-0.20 (-3.37)
Spain			-0.84 (-22.3)		-0.03 (-2.56)			0.0045 (2.70)		-0.11 (-3.14)
Sweden			-0.90 (-9.05)		-0.06 (-2.16)			0.0074 (2.68)		-0.18 (-2.35)
UK			-0.83 (-13.1)		-0.02 (-1.99)					-0.06 (-2.84)
US			-0.50 (-5.39)		-0.16 (-3.28)			0.02 (2.74)		-0.49 (-4.47)

	intercept	Π^e_{it-1}	CA_{it-1}	DR_{it-1}
CCE-MG	-0.123 (-4.15)	-0.777 (-20.19)	-0.010 (-3.96)	-0.060 (-3.27)

Notes: The estimations are obtained from the Pooled Mean Group approach with long-run parameters estimated with CCE techniques. The coefficients of the variables $\bar{r}_t$ *and* $\overline{X}_{1t}$ of equation (6a) have not been reported in the table. t-statistics are in brackets. *r* – real long-term interest rate; *CA* – current account balance; π^e – inflation surprises; *DR* – debt ratio.

Table 8b – Panel Error-Correction estimations for r_{it}, X_2= (Π^e, CA, GBR), 1973-2008

	$\Delta\ r_{it-1}$	$\Delta\ r_{it-2}$	$\Delta\Pi^e_{it}$	$\Delta\Pi^e_{it-1}$	ΔCA_{it}	ΔCA_{it-1}	ΔCA_{it-2}	ΔGBR_{it}	ΔGBR_{it-1}	Loading factor λ_i
Austria			-0.53		-0.07			-0.09		-0.21
			(-5.81)		(-2.29)			(-2.43)		(-2.53)
Belgium			-0.68		-0.04			-0.05		-0.12
			(-9.31)		(-2.63)			(-2.88)		(-2.99)
Canada			-0.81		-0.03			-0.04		-0.09
			(-10.4)		(-2.11)			(-2.17)		(-2.27)
Denmark			-0.67		-0.04	-0.50		-0.06	0.34	-0.14
			(-7.18)		(-2.00)	(-2.74)		(-2.13)	(2.84)	(-2.22)
Finland			-0.49							-0.07
			(-3.99)							(-2.86)
France			-0.76		-0.02	-0.26		-0.03		-0.07
			(-10.7)		(-2.02)	(-3.72)		(-2.13)		(-2.17)
Germany	0.24		-0.76	0.33	-0.05			-0.06		-0.15
	(2.03)		(-12.1)	(2.57)	(-2.46)			(-2.82)		(-2.94)
Ireland			-0.04		-0.11			-0.15		-0.35
			(-1.99)		(-2.95)			(-3.26)		(-3.65)
Italy			-0.83	0.13	-0.05			-0.07	0.20	-0.16
			(-22.7)	(3.04)	(-3.35)			(-3.96)	(2.24)	(-4.45)
Japan	-0.47	-0.16	-0.57	-0.23	-0.08	0.21		-0.11		-0.26
	(-3.6)	(-2.4)	(-8.56)	(-2.07)	(-3.78)	(2.25)		(-4.82)		(-5.46)
Luxembourg			-1.04							-0.17
			(-2.02)							(-3.74)
Netherlands			-0.83		-0.04	-0.08	-0.39	-0.06		-0.14
			(-16.6)		(-2.49)	(-2.56)	(-3.38)	(-2.88)		(-3.01)
Portugal			-0.64		-0.07			-0.10		-0.23
			(-8.09)		(-3.16)			(-3.65)		(-3.94)
Spain			-0.81		-0.02			-0.03		-0.09
			(-19.7)		(-2.18)			(-2.40)		(-2.46)
Sweden			-0.77							-0.13
			(-7.69)							(-2.70)
UK			-0.83							-0.05
			(-13.2)							(-2.76)
US			-0.72		-0.32	0.39		-0.42		
			(-8.25)		(-4.94)	(2.61)		(-8.62)		

	intercept	Π^e_{it-1}	CA_{it-1}	GBR_{it-1}
CCE-MG	0.110	-0.760	-0.062	-0.015
	(3.96)	(-15.1)	(-5.25)	(-3.34)

Notes: The estimations are obtained from the Pooled Mean Group approach with long-run parameters estimated with CCE techniques. The coefficients of the variables $\bar{r}_t$ *and* $\overline{X}_{1t}$ of equation (6b) have not been reported in the table. t-statistics are in brackets. *r* – real long-term interest rate; *CA* – current account balance; π^e – inflation surprises; *GBR* – budget balance ratio.

Table 8c – Panel Error-Correction estimations for r_{it}, X_3= (Π^e, CA, DR, GBR, TCR), 1973-2008

	Δr_{it-1}	$\Delta\Pi^e_{it}$	$\Delta\Pi^e_{it-1}$	ΔCA_{it}	ΔCA_{it-1}	ΔDR_{it}	ΔDR_{it-1}	ΔGBR_{it}	ΔGBR_{it-1}	ΔTCR_{it}	ΔTCR_{it-1}	Loading factor λ_i
	0.14	-0.54	0.24	-0.07				-0.09	0.27	0.11		-0.26
Austria	(2.65)	(-6.74)	(2.23)	(-2.3)				(-2.36)	(3.47)	(2.31)		(-2.97)
		-0.67		-0.03				-0.04				-0.12
Belgium		(-8.41)		(-2.16)				(-2.49)				(-2.66)
		-0.86		-0.02				-0.04				-0.10
Canada		(-11.9)		(-2.33)				(-2.25)				(-2.96)
		-0.70		-0.03				-0.05	0.33			-0.13
Denmark		(-7.78)		(-2.02)				(-2.07)	(2.85)			(-2.57)
		-0.50										-0.08
Finland		(-4.11)										(-2.04)
				-0.27				-0.37				-0.97
France				(-3.52)				(-4.02)				(-6.25)
	0.25	-0.84	0.35	-0.03		-0.10		-0.05				-0.13
Germany	(2.21)	(-12.6)	(2.84)	(-2.17)		(-2.01)		(-2.51)				(-2.93)
		-0.56		-0.05				-0.06				-0.18
Ireland		(-5.58)		(-2.13)				(-2.24)				(-2.68)
		-0.83	0.18	-0.03				-0.04	0.17	0.03	-0.17	-0.12
Italy		(-25.6)	(4.66)	(-2.47)				(-2.57)	(2.10)	(2.71)	(-3.77)	(-3.54)
	-0.31	-0.72	-0.25	-0.07	0.21	-0.12		-0.10				-0.28
Japan	(-2.78)	(-11.1)	(-2.24)	(-2.84)	(2.28)	(-3.21)		(-3.18)				(-5.02)
		-0.94		-0.04				-0.05				-0.15
Luxembourg		(-10.0)		(-2.21)				(-2.26)				(-2.92)
	-0.10	-0.82				0.06	-0.16					-0.06
Netherlands	(-2.09)	(-17.3)				(1.99)	(-2.07)					(-2.35)
		-0.68		-0.05				-0.07				-0.20
Portugal		(-8.65)		(-2.45)				(-2.94)				(-3.57)
		-0.82						-0.03				-0.08
Spain		(-20.2)						(-3.02)				(-3.30)
		-0.81		-0.05		-0.16		-0.07		0.05		-0.20
Sweden		(-9.82)		(-2.41)		(-2.81)		(-2.54)		(2.22)		(-3.28)
		-0.91		-0.02				-0.02				-0.07
UK		(-13.4)		(-2.02)				(-2.15)				(-2.52)
	-0.36	-0.48	-0.37	-0.10				-0.14				-0.38
US	(-2.80)	(-5.77)	(-2.70)	(-2.59)				(-2.94)				(-3.79)

	intercept	Π^e_{it-1}	CA_{it-1}	DR_{it-1}	GBR_{it-1}	TCR_{it-1}
CCE-MG	-0091	-0.892	-0.03	-0.137	-0.041	-0.026
	(-5.26)	(-17.2)	(-4.32)	(-6.25)	(-2.48)	(-3.98)

Notes: The estimations are obtained from the Pooled Mean Group approach with long-run parameters estimated with CCE techniques. The coefficients of the variables $\bar{r}_t$ *and* $\overline{X}_{1t}$ of equation (6c) have not been reported in the table. t-statistics are in brackets. *r* – real long-term interest rate; *CA* – current account balance; π^e – inflation surprises; *DR* – debt ratio; *GBR* – budget balance ratio; *TCR* – real effective exchange rate.

Table 8d – Panel Error-Correction estimations for r_{it}, X_4= (Π^e, CA, DR, LIQ), 1973-2008

	Δr_{it-1}	$\Delta\Pi^e_{it}$	$\Delta\Pi^e_{it-1}$	ΔCA_{it}	ΔCA_{it-1}	ΔDR_{it}	ΔDR_{it-1}	$\Delta LIQS_{it}$	$\Delta LIQS_{it-1}$	Loading factor λ_i
		-0.621		-0.052		0.0094				-0.185
Austria		(-7.52)		(-2.71)		(2.81)				(-2.97)
		-0.768								-0.095
Belgium		(-7.51)								(-3.15)
		-0.823		-0.0223	-0.146	0.0040		7.671	-12.95	-0.0792
Canada		(-11.0)		(-2.08)	(-2.19)	(2.19)		(2.44)	(-2.55)	(-2.54)
		-0.659	0.185	-0.448						-0.0060
Denmark		(-6.84)	(2.87)	(-2.54)						(-1.87)
		-0.511								-0.0566
Finland		(-4.34)								(-2.17)
		-0.349	-0.616	-0.281		0.050	-0.256	1.55		-0.951
France		(-2.90)	(-3.32)	(-3.67)		(4.37)	(-2.05)	(2.64)		(-3.2)
	0.348	-0.980	0.340	-0.060	-0.415	0.010	0.175	0.332		-0.213
Germany	(3.13)	(-12.8)	(2.49)	(-2.45)	(-3.27)	(2.67)	(2.83)	(2.27)		(-3.63)
		-0.735								-0.073
Ireland		(-7.68)								(-2.38)
	-0.144	-0.842		-0.041		0.007		0.230		-0.147
Italy	(-3.43)	(-22.5)		(-2.78)		(3.34)		(2.37)		(-4.91)
		-0.767		-0.081	0.268	0.014		0.448		-0.288
Japan		(-11.3)		(-2.91)	(2.65)	(3.04)		(2.65)		(-5.15)
		-1.102		-0.047	-0.262	0.0085		0.261		-0.167
Luxembourg		(-11.6)		(-2.58)	(-2.10)	(2.94)		(2.10)		(-3.84)
	-0.107	-0.817						0.070		-0.045
Netherlands	(-2.16)	(-18.6)						(2.95)		(-2.91)
		-0.702		-0.059		0.010		0.327	-6.553	-0.210
Portugal		(-9.26)		(-2.41)		(2.55)		(2.19)	(-7.22)	(-3.41)
		-0.872								-0.0033
Spain		(-21.1)								(-2.34)
		-0.853				-0.215				-0.0150
Sweden		(-9.34)				(-2.80)				(-2.05)
		-0.857		-0.021		0.0038				-0.075
UK		(-13.5)		(-2.91)		(2.06)				(-3.28)
		-0.404		-0.170	0.242	0.030		0.944		-0.606
US		(-4.64)		(-2.82)	(2.33)	(2.89)		(2.21)		(-4.14)
	intercept	Π^e_{it-1}	CA_{it-1}	DR_{it-1}	$LIQS_{it-1}$					
CCE-MG	-6.21	- 0.807	-0.008	-0.009	-1.749					
	(-5.28)	(-21.7)	(-4.28)	(-3.36)	(-5.35)					

Notes: The estimations are obtained from the Pooled Mean Group approach with long-run parameters estimated with CCE techniques. The coefficients of the variables $\bar{r}_t$ *and* $\bar{X}_{1t}$ of equation (6d) have not been reported in the table. T-statistics are in brackets. *r* – real long-term interest rate; *CA* – current account balance; π^e – inflation surprises; *DR* – debt ratio; *LIQ* – liquidity proxy based on the debt share.

The government debt ratio shows up in the short-run estimated coefficients with a positive sign (Table 8a), implying an upward pressure on the real interest rate, and also with a negative sign at one lag for some cases, which could eventually be related to gains from liquidity. Again, improvements in the government budget balance (Table 8b) contribute to a reduction in the real sovereign yields. The real effective exchange rate depreciation movements push up the real interest rate, which may reflect higher exchange risk linked to external imbalances (Table 8c), while better current account positions also reduce real interest rates. Interestingly, our liquidity proxy (Table 8d) shows up as statistically significant in the error-correction estimations, implying that liquidity is also relevant from a short-term perspective, as already seen before in the long-run estimation results. Naturally, we can not discard different perceptions towards liquidity if financial conditions are extremely disruptive as for instance, in a financial crisis. Finally, the magnitude of the speed of adjustment across the several error-correction specifications is rather similar.

5. Conclusion

In this paper we assessed the short and long-run behaviour of long-term government bond yields for a set of 17 OECD countries, for the period 1973-2008. We employed a dynamic panel approach, in order to reflect financial and economic integration, and to increase the performance and accuracy of the tests. We also used simulation and bootstrap methods to compute the critical values, and to take into account the cross-country dependences in the sovereign bond segment of the capital markets. Indeed, as put in evidence in the literature, panel unit root tests of the first generation can lead to spurious results if one ignores the existence of significant degrees of positive error cross-section dependence (reflecting, for instance, financial markets integration and liberalisation, or increased business cycle synchronization).

Indeed, we were able to reject the null hypothesis of cross-section independence for the real long-term interest rates (and for its determinants), since the positive error cross-section dependence cannot be ignored. Therefore, after having established, with the use of 2^{nd} generation panel unit root tests, that all the series in the sample panel are non-stationary and integrated of order one, we have undertaken afterwards an adequate bootstrap panel cointegration analysis.

Specifically, first we estimated for each country the long-run relationship between real long-term interest rates and its potential determinants using the Common Correlated Effects (CCE and CCE-MG) procedures of Pesaran (2006), that allow for cross-section dependencies (potentially arising from multiple unobserved common factors). Then, we estimated the panel error correction models for the real long-term interest rates, where short- and long-run effects are estimated jointly from a general autoregressive distributed-lag (ARDL) model using the Pooled Mean Group approach of Pesaran, Shin and Smith (1999), and where short-run effects were allowed to vary across countries. This approach takes advantage of the increased power of panel techniques.

From an economic point of view, it is relevant to find indeed such cross-section dependence, both for the financial series and for the macroeconomic and fiscal variables. In fact, this provides evidence of significant capital markets integration at the OECD level, stemming, for instance, from common fiscal behaviour, notably in the European Union, financial markets' integration and liberalisation, and business cycle synchronization, which sovereign government debt issuers cannot discard lightly.

The results of our analysis also show that in addition to common movements in sovereign yields, investors are also aware of such country specific factors as inflation, budgetary and current account imbalances, and real effective exchange rates. A better (more positive) government budget balance reduces, as expected, the real long-term interest rate in almost all countries. Additionally, the developments in current account balances also carry relevant long-run information for real interest rates. Indeed, the deterioration of the current account balance would signal a widening gap between savings and investment and long-term interest rates may be pushed upwards.

Moreover, our results illustrate that over the longer run real long-term interest rates and their potential determinants move together in this sample of OECD countries. Therefore, identifying the determinants of real long-term interest rates, over long periods as captured by the cointegration analysis, offers additional valuable information notably for financing choices decisions by the sovereign issuers and government investment decisions. Interestingly, some long-run determinants of real long-term interest rates, which were uncovered in the panel cointegration estimation, such as liquidity, are also relevant from a short-run perspective.

From a policymaking point of view the relationship between fiscal and external imbalances and long-term interest rates is timely in the context of economic and financial uncertainty, when pressures for macroeconomic activism are exercised on fiscal authorities, increasing short- and medium-term fiscal risks. Furthermore, it is often argued that large and unsustainable fiscal deficits can endanger the coherence of national macroeconomic policies and may jeopardize the price-stability objectives of a monetary union.

Regarding future work, we could envisage evaluating the data at different frequencies depending on the time-span and variable availability. For instance, mixed frequencies can, to some extent be recovered for the relevant financial and macro variables, and the issue has not been dealt in the existing literature at length.

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Appendix – Data sources

Long-term interest rate – Government Bond Yield, refers to one or more series representing yields to maturity of government bonds or other bonds that would indicate longer term rates, code IFS 61.Z.F. Source: International Financial Statistics, IMF.

Inflation – Consumer Price Index, code IFS 64.XZF. Source: International Financial Statistics, IMF.

Government debt ratio – Debt-to-GDP ratio, code 1.0.319.0.UDGGL. Source: European Commission AMECO database.

Government balance ratio – Budget balance-to-GDP ratio, code 1.0.319.0.UBLGE. Source: European Commission AMECO database.

GDP – GDP at market prices, code 1.0.0.0.UVGD. Source: European Commission AMECO database.

Current account balance – Current Balance as a Percentage of GDP, code CBGDPR. Source: Balance of Payments, OECD Economic Outlook.

Real effective exchange rate – chain-linked index with base period 2005, code CCRETT01.IXOB. Source: OECD Main Economic Indicators.