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Bond Markets and Economic Growth

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Bond Markets and Economic Growth

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Abstract

This paper examines the relationship between the development of the aggregate bond markets and real GDP in 13 highly developed economies. The recent interest in the ties between the real and the financial sector has usually been on the banking sector and the stock markets, rather ignoring the bond markets as a third essential source of external finance. We fill this gap by providing empirical evidence for causality patterns supporting the supply-leading approach in the USA, UK, Switzerland, Germany, Austria, the Netherlands and Spain over the 1950 to 2000 period. In the cases of Japan, Finland and Italy we find evidence of interdependence between bond market capitalization growth and real output growth. Granger causality test and co-integration approach are employed to support this conjecture.

JEL Classification: E44, O16, O40.

Keywords: bond markets, financial markets, economic development, Granger causality.

1. Introduction

The continuous attention, which attracts the nexus between the financial and the real sector, reflects the importance of this topic for economic development.[1] However, the growing literature on the subject has largely ignored the bond markets despite their role as an essential source of external finance (Wachtel, 2001; DeBondt, 2002). The few previous research efforts on bond markets concentrated only on certain regions and/or involved rather short time horizons (DeBondt, 2002), dealt with financial crisis situations rather than the whole business cycle (Herring and Chatusripitak, 2000; Batten and Kim, 2000), or linked GDP growth to the term structure of interest rates in order to forecast recessions (Harvey, 1989, 1991; Gamber, 1996; Gerlach and Smets, 1997; Ahrens, 2002).

We close this gap and provide first empirical, multi-country time series evidence on the link between bond markets and real GDP growth. Specifically, we model the causal relationship from the bond market capitalization to GDP growth for the USA, Japan, Switzerland and most European Economic Area (EEA) countries over the last fifty years (data-permitting).

The paper is organized as follows: section two presents descriptive statistics of the bond markets and the rationale for this paper, section three reviews the most recent studies which partially fit the line of our research and provides a tabular overview of the empirical literature on the finance-growth nexus, section four outlines the methodology used, section five presents the empirical findings and section six discusses the results and sketches some ideas for further research. Section seven concludes. A description of the data set is provided in the Data Appendix.

2. Rationale

In comparison with bank loans and trade credits, which are the most important sources of external finance in the European Union (ECB, 2001:42 ff.), stocks and bonds play a less important role as financing instruments. Even in the USA, whose system is commonly classified as securities-oriented, external financial transactions of non-financial corporations tend to be dominated by credits and loans (ECB, 2001:44). Notwithstanding this assertion, a closer analysis of the financial markets shows that the aggregate bond market capitalization is as important as stock market capitalization (both measured by the amounts outstanding) and amounts in total to about one-third of the entire financial assets in the USA (32.8 %), and in the EU (EU15: 33.1%, Euro-12: 35.1%).

[1] To name but two large research networks: Research teams of the World Bank (http://www.worldbank.org/research/projects/finstructure) specialise on financial structures issues and CEPR (http://www.cepr.org/research/networks/fertn) on the legal and political framework of finance and their implications for real growth.

In ten out of the nineteen countries in our sample, the volume of debt securities (government and corporate bonds) is higher than the stock amounts outstanding measured as a percentage of total financial assets (TFA): Japan, Norway, Germany, Italy, Belgium, Denmark, Austria, Luxembourg, Ireland and Portugal. Luxembourg, which serves as a hub for international bond issues, has the highest bonds to TFA ratio (68.5%), whereas in Switzerland this ratio amounts to only 14%. In six bond-oriented EEA members the debt securities sector makes up the largest share of TFA and exceeds both the credit and the stock markets: Norway, Italy, Belgium, Denmark, Luxembourg and Ireland.

Table 1: Bond amounts outstanding and relative size of major financial sectors in 1999

Country	*Bond amounts outstanding in mn $*	*Financial assets in % of GDP*				*Financial sectors in % of TFA*		
		Bonds	*Credit*	*Stocks*	*TFA*	*Bonds*	*Credit*	*Stocks*
EU15+4	31,599,500	149	194	145	488	30.4	39.8	29.8
USA	15,442,700	166	162	179	507	32.8	31.9	35.3
JPN	6,570,300	136	331	92	558	24.3	59.2	16.5
SUI	185,700	76	185	285	546	14.0	33.9	52.2
NOR	94,500	64	58	43	165	38.7	35.2	26.1
EU15	9,306,300	116	121	113	350	33.1	34.6	32.2
Euro-12	7,179,000	114	120	91	326	35.1	36.9	28.0
GER	2,160,500	108	147	72	328	33.1	44.9	22.0
UK	1,504,900	104	127	198	430	24.3	29.5	46.1
ITA	1,451,300	130	96	65	291	44.7	32.8	22.5
FRA	1,381,300	102	101	111	314	32.5	32.2	35.3
NL	676,500	180	141	185	506	35.6	27.8	36.6
E	393,600	70	116	77	263	26.7	44.1	29.2
BEL	368,400	157	140	78	375	41.8	37.3	20.9
SWE	333,200	144	114	161	420	34.3	27.3	38.4
DNK	289,200	176	57	64	297	59.2	19.3	21.5
AUT	228,000	115	125	17	257	44.8	48.7	6.5
GRE	162,426	99	94	170	363	27.2	26.0	46.8
FIN	118,100	97	58	286	441	21.9	13.1	64.9
IRL	94,900	108	107	78	293	36.8	36.5	26.7
LUX	79,366	627	92	197	916	68.5	10.0	21.5
POR	77,800	72	122	63	258	28.0	47.4	24.5

Notes: Calculations are based on real 1999 data. Countries are ranked in the descending order of the size of the bond market in absolute terms. TFA: Total financial assets = bond amounts outstanding + stock amounts outstanding + domestic credit provided by banks and MFIs.
Sources: IFS, BIS, FIBV, OECD, national sources.

Despite the growing corporate bond issuance in Europe, especially after the introduction of the euro as a tangible currency in the European Union (Fink and Fenz, 2002; Neaime and Paschakis, 2002), bond financing has not attracted the deserved attention by the empirical literature. Denomination in euro facilitates companies to issue corporate bonds on other than their domestic markets. The outstanding amount of bonds has been steadily increasing over time and since 1998 at a relatively faster pace than before. Net issuance rose sharply in the first two quarters of 1998. After some temporarily decline it reached its peak in 1999 (more than 360 billions of US dollars) and remained at relatively high levels.

Considering the floating regime since its inception in 1973 until 2000, the average growth rates of the bond markets in the countries covered in our study are very impressive. Unfortunately, the time series for Norway, Denmark, Greece, Ireland and Luxembourg only start in 1990, what disqualifies them for further econometric analysis. However, we can provide average annual growth rates for all the EU15 and the USA, Japan, Switzerland and Norway from 1981-1990 and 1991-2000 (Table 2). The growth rates are calculated as $\left(\sqrt[k]{\frac{x_{t+k}}{x_t}}-1\right)*100$, where k is the number of years, and x_t denotes the bonds amount outstanding in the starting year. Netherlands' bond markets display the same growth rate in both periods, whereas the bond markets in the other countries grew at a different pace during the sub-periods.[2]

Table 2: Bond Market Development:
Average annual bond market growth in percent between 1981-1990 and 1991-2000

Average growth rate p.a.	*POR*	*FIN*	*Nl*	*FRA*	*USA*	*SUI*	*AUT*	*E*	*SWE*	*JPN*	*ITA*	*BEL*	*GER*	*UK*	*LUX*	*NOR*	*GRE*	*DNK*	*IRL*
1981-90	50.72	13.00	11.75	9.52	9.07	9.03	7.27	7.04	5.71	6.37	5.83	4.78	3.85	-1.36	Time series start in 1990.				
1991-00	5.21	4.86	11.52	5.97	4.97	1.00	7.60	6.80	5.06	5.02	3.83	2.97	9.58	11.33	8.82	7.70	6.75	2.89	10.43

Note: Values are given in percent. Countries are ranked in the descending order of the first period's growth rates.
Source: own calculations.

3. Literature Review

3.1. Possible causal relationships between the financial and real sector

Before turning to the overview of the vast literature on the finance-growth nexus, it is useful to discuss briefly the possible hypotheses describing the causality patterns between economic activity and finance.

Five possible hypotheses with respect to the causal ties between the bond and the real sectors can be derived: (1) supply-leading; (2) demand-leading; (3) interdependence; (4) no causal relation; and (5) negative causality from finance to growth. These five hypotheses can be combined in a number of syntheses, basically with respect to evolutionary sequencing or different needs of industries and/or enterprises. The supply-leading hypothesis maintains that accumulation of financial assets triggers economic growth. The modern branch of the supply-leading finance literature unarguably begins with the works of Mc Kinnon (1973) and Shaw (1973). Since then, the major consensus is that finance positively influences real economic performance. The demand-leading hypothesis assumes that real growth drives the emergence and establishment of financial centers. This hypothesis regards financial development as endogenously determined by the real economy or its needs - a view that is consistent with the Coase (1956) theorem and the new institutional economics, which argue that institutions adjust to market imperfections in a way that maximizes individual utilities.

[2] The reason for the astonishing growth rate of average 51% p.a. during the first sub-period in Portugal is probably attributable to regulation changes in course of the EU accession as bond markets start to develop rapidly after the EU accession. From 1981 until 1987 the real bond market capitalization measured in 1995 USD grew from 894.41 to 2,348.05 millions, i.e. at an annual growth rate of 15% on average, and within only a year, from 1987 to 1988, the value reached 45,538.17 millions.

The view that there is no causal link between the financial sector and real economic development is expressed most prominently by Lucas (1988). This assertion holds only in a neo-classical world of zero transaction costs (Graff, 2000) and perfect information. In such a world, the Modigliani-Miller theorem (1958) holds and institutions, particularly financial institutions, do not matter. The irrelevance hypothesis has come under attack from various sides. Economists increasingly deny the existence of frictionless markets, primarily due to informational and agency problems, transaction costs, etc. Generally, the importance of institutions is now more acknowledged than in the past and a large number of empirical studies present strong evidence for the relevance of finance for real growth. Most of them use as impulse variables one or more financial variables, such as various money aggregates, usually relative to (real) GDP, credits to non-financial institutions, and bank assets as a measure of the banking sector. Market capitalization and sometimes stock market turnover are drawn to describe the capital markets. Researchers have only recently attempted to include internal financing (notably Rivaud, Dunset, Dubocage, Salais, 2001, and to a lesser extent Claessens and Laeven, 2002). Control variables and conditioning sets of empirical studies cover a wide range of macroeconomic, institutional and educational indicators. In the studies reviewed here, the dependent variable is mainly real per capita GDP growth. In some cases investment-related variables are used also. Studies operating on the industry-level additionally use value added as a dependent variable.

3.2. Empirical studies on bank intermediation, stock markets and growth

While cross-country studies normally test the a priori assumption that finance influences the real economy (Levine, Renelt, 1992; King, Levine, 1992; Atje, Jovanovic, 1993; Sala-i-Martin, 1997; Berthélemy, Varoudakis, 1997; Graff, 2000; Rivaud-Danset, Dubocage, Salais, 2001; Leahy, Schich, Wehinger, Pelgrin, Thorgeirsson, 2001; Claessens, Laeven, 2002, Hahn, 2002), time-series analyses address the question of causality mostly via Granger causality tests (Arestis and Demetriades, 1997; Neusser and Kugler, 1998; Darrat, 1999; Shan, Morris and Sun, 2001).

There is weak empirical evidence in favor of the supply-leading pattern in the early stages of financial and general economic development, as well as of demand-leading causality in the very long-run. The supply-leading pattern can be traced back to Joseph Schumpeter (1911), who asserted that a well-functioning banking system promotes technological innovation by selectively funding the most innovative and creative enterprises having the best prospects of success. Most studies using time-series analysis focus on the banking sector and the stock markets in order to evaluate the extent to which bank intermediation and efficient stock markets positively influence real growth. The evidence provided is heterogeneous.

The most prominent scholar in cross country analysis is Ross Levine. Levine (1997) summarizes the existing theories into the “finance-growth nexus” framework and assesses the quantitative importance of the financial system to economic growth, though he also restricts himself to banking intermediation and stock markets only. He advocates the functional approach of the role of financial systems, which focuses upon the relationships between growth and the quality of the functions provided by the financial system (these include monitoring managers, facilitating trading risk, allocating capital, mobilizing savings, etc.). Third factors, such as a country's legal system and political institutions, are further believed to play a role for the financial and economic development.[3] The statistical evidence he presents is based on cross-country comparisons and individual country studies for 35 countries from 1960 to 1989. Four measures of the degree of smooth functioning of the financial system relative to real per capita GDP in 1985 are used: the ratio of liquid liabilities to GDP; the ratio of bank credit to bank credit plus central bank domestic assets; the ratio of credit allocated to private enterprises to total domestic credit (excluding credit to banks); and credit to private enterprises relative to GDP. A positive, statistically significant correlation is found between real per capita GDP and the extent to which loans are directed to the private sector. In a more recent empirical paper Levine and Zervos (1998) examine the role of stock markets and banks in spurring economic growth. Using data for 47 countries over the 1976 to 1993 period they find that both stock market liquidity and banking development together positively predict real growth, capital accumulation and productivity improvements.

A selection of cross-section studies with contrasting findings is provided in table 3. Broader overviews can be found in Graff (2000) and Blum et al. (2002). King and Levine (1993), Leahy et al. (2001) and Claessens and Laeven (2002) find a positive impact on growth, to which a finance-friendly legal framework can also contribute. Harris (1997) and Hahn (2002) find minor causal links. Rivaud-Danset et al. (2001) find no correlation between financial variables and industry performance, and Berthelemy and Varoudakis (1997) find a negative relationship between financial depth and real growth. The studies that employ stock market capitalization as explanatory variable indicate a positive influence on economic growth. Using data for up to 47 countries during the 1976-1993 period, Levine and Zervos (1998) find that when entered together in regressions, both stock market liquidity and banking development positively predict real growth, capital accumulation and productivity improvements.

[3] For extensive elaboration on this notion see Levine (1999), Law, Finance, and Economic Growth, *Journal of Financial Intermediation*, also Levine (1998), The Legal Environment, Bank, and Long-Run Economic Growth, *Journal of Money, Credit, and Banking, Vol. 30, No. 3*, pp.597-620, Levine, Loayza, Beck (2000), Financial intermediation and growth: Causality and causes, *Journal of Monetary Economics 46*, pp.31-77, and La Porta, Lopez-de-Silanes, Shleifer and Vishny (1998), Law and Finance, *Journal of Political Economy, Vol. 106, No. 6*, pp.1113-1115.

Table 3: Cross-country studies on the links between financial intermediation, stocks and real growth ordered chronologically by year of publication

Authors/Year of publication	*Country sample*	*Time span*	*Dependent variable*	*EFV I: Banking sector*	*EFV II: Stocks*	*Other explanatory variables*	*Control/ other variables*	*Method*	*Investigated links between*	*Major findings*	*Additional results*
King, Levine 1993	Up to 80 countries (up to 19 EU+; up to 3 ACC)	1960-1989	Output growth; capital stock growth; productivity growth	Liquid liabilities/GDP; Assets of commercial & central banks/GDP Private credit/GDP Credits to private enterprises/GDP; Credits to private, public enterprises & local governments			Initial income; School enrolment rate; Trade exposure; Government spending/GDP Inflation	Panel analysis: 5-year periods	Banking intermediation & economic growth	Positive relationships between financial intermediation and economic growth	Case studies and firm-level studies: liberalisation of the financial market leads to higher growth rates, seems to be related to the financial crisis that occurred in many countries 3-5 years after financial liberalisation
Berthélemy, Varoudakis 1997	85 countries (details not available)	1960-1990	Real income per capita growth	ln (M2/GDP)		Convergence; investment rate; inflation; education; trade exposure; etc.		Panel regression analysis: 5-year periods	Financial depth and economic growth	Negative relationship between financial depth and real growth	
Harris 1997	49 countries: 19 EU+; 2 ACC.	1980-1991	Real GDP per capita growth		Investment ratio*stock market turnover/ GDP		Investment rate; population growth	2SLS regression analysis	Financial sectors and economic growth	Very limited positive relationship between capital markets & growth only for high-income countries, contrasting to Atje, Jovanovic (1993)	
Levine, Zervos 1998	Up to 47 countries	1976-1993	Real GDP growth; Capital stock & productivity Savings	Bank credit/GDP (bank credit = loans by commercial banks and other deposit-taking banks)	Capitalisation of domestic listed companies/ GDP; Value traded/ GDP; Share returns volatility		Initial output; Enrolment rate; Number of revolutions & other social & political variables	Cross-country regression analysis	Links between financial sectors and economic growth	Strong positive relationship between both financial segments and economic growth; no negative effects caused by share volatility or international capital market integration	
Leahy, Schich, Wehinger, Pelgrin, Thorgeirsson 2001	16 EU+	1970-1997	Gross investment	Private credit from deposit banks/GDP; Liquid liabilities/GDP	Stock market capitalisation/GDP	Real long-term interest rate; output growth; inflation & its variation; public revenue & spending; trade exposure; legal indicators		Panel error correction approach building on an ARDL model	Financial development, financial system, innovation & economic growth	Financial development and finance-friendly legal framework enhance growth via innovation	All financial sectors have a positive impact on growth, but shares have the greatest
Rivaud-Danset, Dubocage, Salais	9 EU+	1990-1996	Mark-up Value added Return on			Own funds Leverage Financial debt	Variables controlling for company size	Cluster analysis	Companies' financing structures &	No correlation between financial variables & industry	Companies' financing structures depend on country

2001			investment			structure Liquid capital			performance	performance	characteristics
Claessens, Laeven 2002	Up to 51 countries	1980-1989	Real value added growth	Private credit/GDP		Cash flow	Legal indicators, esp. dealing with property rights; Country-specific & industry-specific characteristics; Stock market capitalisation	Panel regression (industry level)	Financial development & property rights, and growth of companies depending on external finance	Companies depending on external finance grow faster in economies with developed financial systems & high property protection	Asset allocation effect as important as finance effect. Poor (intellectual) property protection leads to less investment in intangible assets
Hahn 2002	23 countries: 19 EU+; 1 ACC.	1970-2000	Gross investment	Bank credit to the private sector/GDP; Liquid liabilities/GDP	Stock market capitalisation/GDP		Value traded & turnover (to control for forward-orientation of financial markets)	Panel error correction approach building on an ARDL model	Financial development, financial system & economic growth controlling for forward-looking price effects of financial markets	Minor causal links between financial development and economic growth, strong relationship due to forward-looking price effects	

Notes: Papers are ordered chronologically by year of publication. Slashes ("/") in table texts denote division, asterisk denotes multiplication; "EU+" refers to EU plus 5 OECD countries. *EFV* denotes explanatory financial variables. *ARDL*: an autoregressive distributed lag model method, which builds on time series methods, but a panel is analysed.

After having offered a concise selection of the existing cross country studies, we address the time-series analyses on the finance-growth nexus (see table 4). As mentioned before, previous time-series studies concentrate only on the banking sector and stock markets and neglect the debt markets, but they partly employ a similar testing methodology as utilized here. The focus on the banking and the stock market segments is attributable to the traditionally dominant market position of these financial sectors and the availability of data. The most commonly found pattern supports the supply-leading hypothesis, and in some cases bi-directional causality. The study of Hannson and Jonung (1997) is unique as it examines a single country (Sweden) over a very long time period, which makes it possible to test long-term causality patterns in the sense of Patrick (1966) or Gerschenkron (1962). The evidence based on cointegration analysis shows bi-directional causality between bank credit and real per capita GDP for most of the time from 1834 to 1991 and supply-leading causality for the 1890 to 1934 period. Al-Yousif (2002) applies a cross-country and a time-series approach (1970-1999) and finds bi-directional causality to be the dominant, yet not the only observable pattern. Rousseau and Wachtel (1998) find evidence for supply-leading finance in USA, Canada, UK, Sweden, and Norway for the 1871 to 1929 period. Arestis and Demetriades (1997) employ the classical Granger causality test together with system exogeneity analysis to investigate the links between the stock market capitalisation and per capita GDP in the USA and Germany based on quarterly data for a shorter period (1979-1991). In the case of Germany, they find evidence for supply-leading finance, and no causality in the USA. Kugler and Neusser (1998) use financial sector GDP of 13 OECD countries from 1960 to 1997 as an explanatory variable, and the GDP of the manufacturing sector as a dependent variable, and find a whole range of possible causality patterns varying across countries and time. The recent study by Shan, Morris and Sun (2001) covers 6 EU countries, three EU accession countries and China, and investigates the dynamic link between real per capita GDP growth and bank credits measured by the loans granted by commercial and other deposit-taking banks. They utilize the Granger causality test within a VAR setting with a conditioning set and find demand-leading, supply-leading, and bi-directional evidence.

Table 4: Time series studies on the links between banking intermediation, stocks and real growth

Authors/Year of publication	*Country sample*	*Time span*	*Proxy for economic growth*	*FV I: banking intermediation*	*FV II: securities*	*Method: Granger causality test together with*	*Investigated links and causality between*	*Major findings*
Hansson, Jonung 1997	Sweden	1834-1991	Real per capita GDP	Bank credit to non-financial sector per capita		Cointegration analysis (investment per capita as conditioning variable)	Co-evolution of banking intermediation & real income	Mostly unstable relationship between intermediation and output; supply-leading 1890-1934; positive influence of education on supply-leading pattern
Al-Yousif 2002	30 developing countries	1970-1999	Real per capita GDP growth	M1, M2 / GDP		Error correction model/ Panel data approach	Banking intermediation and economic growth	Strong evidence for bi-directional causality; limited evidence for other patterns
Rousseau, Wachtel 1998	USA, UK, Canada, Sweden, Norway	1870-1929	Real per capita GDP growth	Money base; various proxies for intermediation based on bank deposit and credit		Granger causality tests within VAR framework	Banking intermediation and economic growth	Supply-leading in early phase of economic development
Arestis, Demetriades 1997	USA & Germany	1979-1991 (quarterly)	Real per capita GDP	Germany: M2/GDP; USA: Domestic bank credit/GDP	Stock market capitalisation/ GDP; Stock market volatility (16-month standard deviation of share prices)	System exogeneity analysis	Financial sectors and economic growth	Cross-country analysis oversimplifies results; links between financial sectors and growth are different in GER & USA: financial-to-real sector causality in GER; no evidence for unidirectional causality in USA.
Kugler, Neusser 1998	13 EU+	1960-1997	GDP of manufacturing industry (MGDP); Total factor productivity of manufacturing industry (MTFP)		Financial sector GDP (FGDP)	Granger causality tests	Financial sector & growth	MGDP & FGDP are co-integrated in 7 countries; MTFP & FGDP are co-integrated more often; financial-to-real sector-causality only for USA, JPN and GER; inverse causality in some other countries; no evidence for causal ties in small countries.
Shan, Morris, Sun 2001	10 countries (6 EU+ China)	1960-1998 max, 1982-1998 min.	Real per capita GDP growth	Bank credit/GDP (Bank credit = loans by commercial banks and other deposit-taking banks)		VAR with conditioning set	Banking intermediation & economic growth	Bi-directional causality in five countries, demand-leading in three, no causality in two.
Darrat, 1999	Saudi Arabia, Turkey, UAE	1964-1993 (annual), UAE:1973-1993	Percentage changes of real GDP in 1990 prices	Currency ratio = currency/M1; Broad money velocity = M2/ nominal GDP		VECM (Johansen procedure)	Financial deepening (esp. banking development) & growth	Strong support for the supply-leading hypothesis in Turkey. UAE: short-run support. Saudi Arabia: no short-run relations, bi-directional causality over the long run.

While the vast majority of research is concerned with the impact of banks and stocks on growth, studies on the bond-growth nexus are scarce. The few studies on the link between debt securities and growth can be grouped in four categories:

- Theoretical articles,
- Empirical studies,
- Descriptive reports on bond market developments,
- And studies with a regional focus on Central and Eastern Europe (see Table 5).

3.3. Bond-market based forecasts of economic performance

Table 5: The link between debt securities and growth: Overview of the studies on debt securities

Category, Authors, Year of publication		*Focus*	*Market segments and Sample coverage*	*Major findings*
Theoretical articles	*Goldsmith 1969*	Financial structure-financial development-real development relationship.	Primarily intermediation in up to 20 countries: 1860-1963.	The ratios of financial assets to real wealth in different countries converge in the long run.
Empirical studies	*Harvey 1989*	Forecasting the US economic growth via the yield curve.	USA output growth: 1953-89.	Yield-curve based measures are able to explain > 30% of the variation in economic growth, while stock market measures account to about 5% only.
	Harvey 1991	Forecasting German economic crises via the yield curve.	German economic crises: 1969-91.	The IRFM outperforms the forecasts of economic growth of major research institutes and predicts correctly the economic turning points during the surveyed period.
	Gamber 1996	The information content of the yield curve.	USA data Three policy regimes within the 1955-1996 period.	Empirical evidence obtained from the Granger causality tests indicates that when the Fed sets policy to counter inflation or output growth movements, the slope of the yield curve loses its forecasting power with respect to the particular variable. Monetary policy is exogenous with respect to the yield curve.
	Ahrens 2002	The predicting power of the term structure as a lead indicator of recessions.	Eight OECD countries: 1970:01-1996:12.	A simple univariate model is found to be a filter which transforms accurately term spread changes into turning point predictions, and that the term structure is a good recession indicator.
	G. de Bondt, August 2002	Corporate debt securities in the Euro area.	Euro area corporate debt securities: 1999:01-2001:06.	Corporate debt issuance is positively related to M&A and industrial production. Substitution with other sources of finance is related to cost differentials. Timing and size of explanatory factors of corporate debt securities issuance differ across maturity. Corporate bond spreads lag short-term interest rates and lead real economic growth.
Descriptive bank reports on recent developments	*ECB 2001*	Theoretical analysis of corporate finance issues and external sources of corporate finance.	Bond and stock markets, venture capital market. Banking intermediation. Euro area in comparison with USA and Japan: 1997-2000.	The dominant liability in the USA and Europe in 1999 was equity debt, in Japan-loans. Debt securities as financing form was not as important in Euroland as in the US and Japan. External financing increased between 1997-1999. Strong increases in debt securities and venture capital after 1997. The value of equities increased most. According to flow data analysis this was mainly due to higher stock prices and in the Euro area financing in loans exceeded flows in shares.

	Merrill Lynch, 2002	Size and structure of the world bond market 2002	The world bond market size went up by 5% compared to last year. In spite of widening fiscal deficits, the share of central governments' bonds continued to fall to 30%, a 12-year low, and was chiefly an English-speaking and Scandinavian countries phenomenon. In Japan and Euroland government markets continued to grow. Agencies were the fastest growing market in 2001, reflecting the strength of the housing market. Eurobonds ranked second due to their increasing popularity. The rating quality of the corporate sector continued to fall. The emerging markets tradable debt universe increased by 8% compared to 2001.	
	Claes, De Ceuster, Polfliet 2002	Primary market activity in the Eurobond market.	Eurobond market (Bondware Databse): 1980-2000.	Geographical concentration of the Eurobond market with respect to the nationality of issuers. More than 11% issued by US-based entities. Introduction of the euro reinforced the market share of European issues. Majority of Eurobonds are bearer bonds.
Studies dealing with Central and Eastern European debt markets	*Mihaljek, Scatigna, Villar 2002*	Recent trends in debt markets development in the emerging economies.	Domestic debt markets in Asia, Latin America and Emerging Europe: 1994-2000.	Despite considerable growth primarily due to public sector deficits associated with fiscal adjustment and related banking and corporate sector reforms, domestic debt markets remain small compared to industrial countries. Public sector debt issues on average account for 2/3 of domestic debt market volume. Banks hold the largest proportion of bonds in domestic markets; institutional investors have become key holders of domestic debt in Latin America and Central Europe; corporate debt markets remained underdeveloped.
	Köke, Schröder 2002	Eastern European capital markets: historical development & legal settings	Czech Republic, Hungary, Poland: 1995-2000	Stock markets grew unevenly across countries, with developments in the Czech Republic being particularly disappointing. Bond markets dominated by public sector debt securities (exception: the Czech Republic, where corporate bonds are significant).. The importance of internal finance for firms' investment is higher than in EU countries (exception: Hungary). Bank credit remains the main source of external finance even for listed firms.. Most CEE exchanges are too small to survive.
	Haiss, Marin 2002	CEEC bond market developments	Central and Eastern European Countries (CEEC)	The bond market is still dominated by public issuers and bank-based intermediation. Similar to the EU, corporate bond issuance was driven by the telecom and utility sectors. Rising demand from pension funds and limited bank loans volume carry the potential for growth of the corporate bond sector.

Over the last ten years, several studies have shown that the changes in the slope of the yield curve (the difference between the long-term and the short-term yield) contain information about future real economic performance (Harvey, 1989, 1991; Bernanke, 1990; Estralla and Hardouvelis, 1991; Gamber, 1996; Ahrens, 2002).

Using a standard ordinary least squares (OLS) regression linking real growth as a response variables with the yield spread (impulse variable), Harvey (1989) shows empirically that yield-curve based measures are able to explain more than 30% of the variation in economic growth in the USA over the 1953-89 period, while stock market measures account to about 5% only. Moreover, yield-curve based forecasts are found to compare favorably with leading econometric models' forecasts, while forecasts from stock market models do not (Harvey, 1989, p. 38). Harvey (1991) uses the interest rate based forecasting model (IRFM) to forecast the German economic crises during the 1969-1991 period. He utilizes the annual GNP growth in 1980 DM, the money-market rate and the government bond yield from 1969 till 1991. The first component of the IRFM is the yield curve, and the second is a measure of the average propensity to hedge. The intercept is interpreted as the expected level of economic growth when the long-term rate equals the short-term rate, and the coefficient of the term structure is the average level of risk tolerance in the economy. He shows that the IRFM ex post outperforms the forecasts of economic growth of major research institutes and that it has correctly predicted the economic turning points in the period 1969-1991. Gamber (1996) uses the difference between 10-year and 3-month government bond yields over the 1955-1996 period to investigate with Granger causality test whether the information content of the yield curve is only due to its relationship to monetary policy or whether it is an independent indicator of future inflation and output growth. He finds that the slope of the yield curve loses its forecasting power when the Federal Reserve Bank sets policy measures to counter movements in either inflation or output growth. Therefore, it seems that monetary policy is exogenous with respect to the yield curve. Ahrens (2002) uses Markov-switching models to estimate the predicting power of the term structure as a lead indicator of recessions in eight OECD countries between January 1970 and December 1996. He finds that a simple univariate model is a filter which accurately transforms term spread changes into turning point predictions, and that the term structure is a good recession indicator.

Recently, G. De Bondt (2002) provided first evidence on the structure and dynamics of the corporate bond market in Euroland after the introduction of the euro. The time horizon of his study is rather short (1999:01-2001:06). The paper reviews three theoretical frameworks to model corporate bond issuance: the first one models simultaneously all corporate financial liabilities in a portfolio modeling framework, the second one models supply and demand for corporate bonds simultaneously, and the third one is to specify a supply function of corporate bonds. Due to data limitations the last approach is adopted. The financing needs and substitution between debt securities and other sources of corporate finance are considered as main explanatory factors for bond supply. The essential findings are: First, since the introduction of the single currency the euro area debt securities markets contributed to finance mergers and acquisitions (M&A), investment and working capital needs. The introduction of the euro triggered (at least partially) corporate restructuring, which in turn resulted in financing needs reflected by M&A. M&A are reflected in short-term bond issuance in the same month and with a one-quarter lag, and in long-term bond issuance after three quarters. A one-to-one relation between corporate debt securities issuance and industrial production is found. Second, the spread between long-term and short-term interest rates is found to be a relevant factor for the mix between long and short-term debt securities. The direct substitution effects are found to be particularly strong for short-term bond issuance. Third, regression results and Granger causality tests show that various macroeconomic factors explain corporate bond spread movements and that the latter have forecasting power for real output growth.

Our paper advances on earlier work and on the most recently published ECB working paper by De Bondt (2002) in several ways: Our time horizon encompasses 25 to 50 years with several business cycles. We look at the aggregate bond market instead of single segments and thus capture the aggregate effect of gross public and private investment financing. We use changes in stock variables instead of flow variables to account for valuation changes and implicitly for changes in the interest rate level, too. We employ the aggregate bond market capitalization instead of bond issuance or interest rate spreads to achieve comparability with earlier work on stock markets and real growth. Accordingly, our model consists of real GDP growth and real aggregate bond market capitalization. Theoretically, real capital stock growth might be used as an intermediate variable between financial markets and GDP growth. However, there is the usual problem to assemble reliable data on the real capital stock (taking into account depreciation rates and valuation changes over time).

4. Methodology

The purpose of the paper is to examine the hypothesized causal relationship between the bond sector development and economic growth in 13 developed economies over the 1950 to 2000 time period. The direction of the causality is of a particular interest as it may help partly explain the driving forces for the emergence and development of financial markets. Thereby it can convey policy implications how to enhance continuing financial growth, which eventually leads to capital accumulation. The direction of the causality cannot be tested via cross-country approaches, as they usually assume a priori that finance influences growth (e.g. Darrat, 1999).

The most popular procedure when testing causality between any two variables has become the Granger causality test (Granger, 1969). This is partly due to its simplicity, and due to the fact that this method saves degrees of freedom – this property is a great advantage when relatively small samples are utilized (n<50 observations). Granger (1969) defined causality in terms of predictability: a stationary time series, say *X*, is said to cause another stationary time series, *Y*, if the prediction error (measured by the variance) of current *X* declines by using past values of *Y* in addition to the past values of *X*. The question is how much of the current *Y* can be explained by its own past values and does an adding lagged value of *X* improve the explanation. Two-way causation, called "feedback", is frequently the case, that is *X* Granger-causes *Y* and *Y* Granger-causes *X*. Conversely, *Y* is not Granger-causal for *X*, when removing the past values of *Y* from the model does not change the optimal forecast for *X* at any forecast horizon. It is important to note that the statement *"X Granger-causes Y"* does not imply that *Y* is the effect or the result of *X*. Granger causality measures precedence and information content, but does not indicate by itself causality in the more common use of the term. Formally, the bivariate autoregressive model looks like this:

$$\begin{bmatrix} y_t \\ b_t \end{bmatrix} = \sum_{i=1}^{n} \begin{bmatrix} a_{1,i} & a_{2,i} \\ b_{1,i} & b_{2,i} \end{bmatrix} \begin{bmatrix} y_{t-i} \\ b_{t-i} \end{bmatrix} + \begin{bmatrix} u_{y_t} \\ u_{b_t} \end{bmatrix} \qquad (1)$$

where $b_{1,i} = 0, i = 1,2,..., n$, the residual series are *iid*, $u_t \sim IN[o, s^2]$, and the time series variables are stationary, $[y_t, b_t]' \sim I(0)$. Nonstationarity of the variables incorporated in a model leads to the spurious regression problem discussed by Granger and Newbold (1974), Phillips (1986) and Watson (1994). High adjusted R^2 then may only indicate correlated trends instead of truly captured economic relations. Simple correlation tests including four lead and lag years show, that the GDP and the bond market capitalization are highly correlated. The correlation coefficients decrease from 0.99 in the first to a maximum of 0.5 to 0.7 in the fourth year. As both time series exhibit sustained growth trends over time such correlations are not unexpected.

Granger (1986) illustrated that if the non-stationarity is purely due to unit roots, i.e. $x_t \sim I(1)$, then the time series can be made stationary by the linear transformation of differencing, i.e. $x_t - x_{t-1} = Dx_t \sim I(0)$. As usual in applied econometric work, this refers to the concept of weak stationarity (constant first moments) as opposed to strong stationarity. Lower-case letters denote natural logarithms of the non-stationary time series in levels. Defined this way, changes can be interpreted as growth rates in the level of real economic activity and bond market capitalization. The order of integration of the time series is determined via the Augmented Dickey and Fuller (ADF) and Phillips-Perron (PP) unit root tests. The ADF unit root test is most commonly used, but it requires homoscedastic and uncorrelated errors in the data generating process. The PP nonparametric test generalizes the ADF test and allows for less restrictive assumptions for the investigated time series (Darrat, 1999). Both tests are run to ensure unbiased inference about the order of integration of the variables of interest. The optimal lag order of the ADF regressions that guarantees uncorrelated residuals is selected via the Akaike's Information Criterion (AIC), which minimizes the final prediction error (FPE) of the regression.

Having determined the order of integration of the variables, equation (1) can be set up. The number of lags to be included in the system is selected by the AIC with an upper bound of 4 years (that is the assumed duration of one business cycle). A greater lag order is not practicable because of the low frequency of the data and would reduce degrees of freedom. The null hypothesis that b_t does not Granger-cause y_t is rejected if $\sum a_{2,i} \neq 0$. Whether the estimated coefficients are statistically significant is checked with the F-test. By analogy, testing the reverse null hypothesis that y_t does not Granger-cause b_t is equal to testing whether $\sum b_{1,i} \neq 0$.

However, these standard Granger-causality tests are valid only if the two variables are not co-integrated.[4] The economic interpretation of co-integrated systems is that the series may exhibit short-term dynamics but in the long run they converge to some state of equilibrium. As Engle and Granger

(1987) showed, in the presence of co-integration inferences from the above system will be biased as they overlook long-run information. In this case, the Granger's Representation Theorem postulates that it is better to set up an Error Correction Model (ECM). Engle and Granger (1987) proposed a single-equation based two-step approach to test for co-integration between two series. Alternatively, the system-based Johansen co-integration test can be used. Apparently, when investigating the interaction among three or more time series variables the problem of which test to use becomes redundant in favor of the second approach. Relying on Gonzalo and Lee (1998) and Haug (1996), the preference here is given to the Engle-Granger method. On the whole, Gonzalo and Lee (1998) find that the Engle-Granger two-step approach is more robust than the Johansen likelihood-ratio tests in most of the investigated cases. The main reason is that misspecifications of the variables' long-memory components are found to affect their correlation structure more than their variances.[5] Haug (1996) presents evidence for a trade-off between power and size distortions of the examined co-integration tests. Regarding size distortions, when the regressors are endogenous, the Engle-Granger's ADF test and the Johansen maximum eigenvalue statistic exhibit the least size distortions. When the regressors are exogenous, the ADF test performs best. Based on these studies, the EG two-step procedure in connection with the Granger test for causality is used to identify the long-run equilibrium relation between the variables under investigation, if any, and to determine the direction of the short-run causal ties.

To determine whether there exists any long-run equilibrium relationship, we follow Engle-Granger (1987) and fit the co-integrating regressions:

$$y_t = b_0 + b_1 b_t + e_{t,y} \text{ and} \tag{2a}$$

$$b_t = a_0 + a_1 y_t + e_{t,x}, \tag{2b}$$

where y_t is the logarithmic level of real economic activity measured by the real GDP and b_t is the size of the bond markets in the individual countries proxied by the outstanding amount of debt securities issued by central governments and the private and public sectors. To achieve a heteroscedasticity and autocorrelation consistent covariance matrix, Newey and West (1987) corrected standard errors are used instead of the conventional computed standard errors, as in the presence of heteroskedasticity the latter are no longer valid. If the residuals resulting from the above

[4] If the variables of interest have a *common* stochastic trend, linear combinations of them are possibly stationary, i.e. there is no unit root in the relation linking X_t and Y_t. The variables are said to be co-integrated. Formally: if $X_t \sim I(1)$ and $Y_t \sim I(1)$, then $aX_t + bY_t \sim I(0)$.

[5] Gonzalo and Lee (1998) show that the Johansen LR tests tend to find spurious co-integration with probability asymptotically approaching one in two cases: first, when variables with long-memory properties and trend behavior are involved that are not distinguishable from pure I(1) via standard unit root tests, and second, when the variables are non-stationary with a singular or near-singular error covariance matrix. Such a "size pitfall" is more serious than either a simple finite sample problem or a size distortion. When the Johansen procedure is used to analyze co-integrating relations among variables with deterministic linear trends, which are actually not co-integrated, the LR tests statistics diverge and the null is wrongly rejected. The EG test, on the other hand, is not sensitive to incorrect modeling of such components and produces an empirical size very close to the nominal size.

regressions $\{e_t\}$ are not stationary, the series are not co-integrated. The correct model to be specified is a standard VAR*(p)* in first differences with *(p)* being the lag order, as given by equation (1). Conversely, the variables are co-integrated in case of stationary residuals. The econometric model to be specified is a dynamic ECM. Essentially, it is a VAR augmented with an error-correction term. The EC-term is the residual series obtained from the associated co-integrating regressions and is the formal representation of the equilibrium. In addition to the independent lags, the error-correction term provides an extra channel for Granger-causality (Granger, 1986). In the transformed regressions, the Granger-causality restriction corresponds to the restriction that the EC-term does not enter the regression. Therefore, b_t Granger-causes y_t if $\sum a_{2,i} \neq 0$ *and/or* the coefficient of the EC term is statistically significant. It has been argued that the significant lags of b_t represent short-run Granger-causality, and the coefficient of the error-correction term corresponds to long-run Granger-causality (Darrat, 1999). The empirical analyses should therefore focus on the long-run relationships as the investigated relationship is inherently a long-term one. To end with, it has to be noted, that a statistically significant coefficient of the EC-term does not allow for further inferences on the direction of the long-run causality. This difficulty is due the fact that the EC term results from a regression on two non-stationary variables.

5. Empirical Results

The tests are in real terms in order to account for possible inflationary effects. The nominal data are deflated with the GDP deflator. All series are converted to USD using the 1995-dollar exchange rate to reflect real 1995 US dollars in order to achieve a better readability. The 1995 USD exchange rate is preferred to the current USD exchange rate in order to avoid the impact of exchange rate fluctuations on the time series variables. Otherwise the information on the size of the debt markets and the level of economic activity would be distorted. Subsequently, the series are transformed into natural logarithms to produce more homogeneous series. Due to insufficient number of observations, no tests could be performed on Denmark, Luxembourg, Greece, Ireland, Belgium and Norway. Thus, the sample comprises 13 countries.

As expected, the ADF and Phillips-Perron unit-root tests performed on the logs of the series in levels show that all series are non-stationary in levels and stationary in first differences.[6] Therefore we test for co-integration. The decision whether the time series variables are co-integrated or not, is taken according to the ADF test on the resulting residuals. The ADF test rejects the null of non-stationarity

[6] To save space, the complete output of the unit root tests on the time series in levels and first differences (test specifications include an intercept and a linear trend), the co-integrating regressions and the ADF test on the residuals from the co-integrating regressions (test specifications do not include any exogenous variables due to assumed fluctuations around the zero line) are not reproduced here, but are available from the corresponding author upon request.

in Japan, Italy, Finland and Portugal. The evidence is that in these countries real bond market capitalization and real GDP are co-integrated, i.e. there is common equilibrium convergence.

VAR models in first differences with a maximum lag order of four years are applied to seven countries. In the case of Great Britain, only the residuals from the first regression (with bond market capitalization being the dependent left hand-side variable) are found to be stationary, not the residuals from the reverse regression. Therefore, the evidence for co-integration is rather weak and not reliable enough to serve as a basis for an error-correction specification.

No intertemporal relationship could be established for France and Sweden (Table 4). According to the AIC (it selects a lag order of zero) and the F-test on the coefficients of the lagged terms when null restrictions are imposed, all the lags are not significantly different from zero, thus filtering out possible intertemporal interrelations and short-run causality patterns. The impossibility to specify a dynamic system may be interpreted as an absence of causality patterns between the bond sector growth and real output growth based on annual data. It cannot be excluded that lags with duration of less than one year may be significant.

Table 6: OLS test results for co-integration ordered by fit (R^2). Model selection

Model/ Country		*Sample size/ Observations*	*DV*	*ADF residuals test*	R^2	*DW*	*Heteroscedasticity*	*Model*
Long-run equilibrium convergence	*JPN*	1967-1999/ 33	Bond GDP	ADF (6) -3.573 ** ADF (6) -3.627 **	.97834	.227 .230	.0111***	VECM
	ITA	1963-2000/ 38	Bond GDP	ADF (12) -3.370 * ADF (12) -3.342 *	.93149	.281 .267	.0181***	VECM
	FIN	1972-2000/ 29	Bond GDP	ADF (12) -3.914 ** ADF (12) -6.371 ***	.87314	.329 .345	.4300	VECM
	POR	1977-2000/ 24	Bond GDP	DF -3.419 ** DF -3.437 **	.83382	1.321*** 1.122***	.2661	VECM
Short-run dynamics No equilibrium convergence	*AUT*	1960-2000/ 41	Bond GDP	DF -2.005 DF -2.009	.99049	.389 .389	.1869	ΔVAR(1)
	USA	1952-2000/ 49	Bond GDP	ADF (1) -1.471 ADF (1) -1.433	.90866	.045 .045	.7994	ΔVAR(1)
	SUI	1952-2000/ 49	Bond GDP	DF -1.544 DF -1.434	.80547	.181 .159	.1874	ΔVAR(1)
	E	1962-1999/ 38	Bond GDP	ADF (1) -3.114 ADF (1) -3.013	.73765	.409 .318	.2822	ΔVAR(1)
	UK	1976-2000/ 25	Bond GDP	ADF (10) -7.554 *** ADF (8) -2.970	.72745	.285 .221	.2339	ΔVAR(1)
	GER	1960-2000/ 41	Bond GDP	ADF (1) -2.493 ADF (1) -2.468	.97503	.339 .340	.8797	ΔVAR(2)
	NL	1951-2000/ 50	Bond GDP	ADF (1) -1.654 ADF (1) -1.521	.61109	.052 .039	.3281	ΔVAR(2)
Undetermined	*FRA*	1977-2000/ 24	Bond GDP	ADF (1) -1.059 ADF (1) -1.101	.95751	.582 .566	.0406 ***	
	SWE	1964-2000/ 37	Bond GDP	ADF (1) -2.609 ADF (1) -2.642	.68060	.182 .157	.000006 ***	

DV: denotes the dependent variable. DW: is the co-integrating regression Durbin-Watson test statistics.
Heteroskedasticity: the White's test for heteroskedasticity on the regression of squared residuals on fitted values (with cross terms). The reported values are *p*-values for rejection of the null hypothesis of no heteroscedasticity.
Relevant critical values for the Durbin-Watson statistic are tabulated by Engle & Yoo (1987): 1% (1.00), 5% (0.78), 10% (0.69 for 2 variables and sample size n=50. The ADF test regressions include an intercept and a trend. ADF critical values from MacKinnon (1991). "***" denotes significance at the 1% level, "**" at 5%, "*" at 10%. The included lags are reported in brackets.

5.1. Evidence for long-run equilibrium convergence and interdependence

In Japan, Italy, Finland, and Portugal real GDP and the bond market size follow a common stochastic trend in the long run, which is represented by the error-correction term (EC-term). As outlined in the previous section, a statistically significant coefficient of the EC-term can be interpreted as a long-run equilibrium adjustment and long-run causality.

Table 7: Long-run equilibrium convergence VECM models

Country	*DV*	*a*	*EC-term*	Δb_{t-1}	Δb_{t-2}	Δb_{t-3}	Δb_{t-4}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}	*const*	R^2
JPN	Δb_{t-1}	-.231 .153	$y = .423b + 8.863$		1.576 .0001	-1.141 .001	.323 .006	1.301 .00	-1.816 .014	1.456 .008	-.369 .013	.045 .024	.762
	Δy_{t-1}	.217 .003		.406 .00	-.757 .0002	.629 .0003	-.192 .0002		1.824 .00	-1.455 .00	.366 .00	-	.767
ITA	Δb_{t-1}	-.275 .042	$y = .535b + 6.351$		1.576 .00	-1.095 .00	.279 .0004	1.724 .00	-2.497 .0005	1.557 .008	-.334 .052	-	.784
	Δy_{t-1}	.047 .314		.179 .135	-.326 .089	.21 .131	-.049 .192		1.164 .0004	-.666 .016	.164 .061	.018 .006	.585
FIN	Δb_{t-1}	.156 .818	$y = .218b + 9.273$		1.220 .008	-.922 .019	.265 .098	-1.382 .606	1.675 .708	-.766 .821	.227 .829	.124 .020	.579
	Δy_{t-1}	.246 .00		-.017 .645	.045 .582	-.026 .686	-.003 .838		1.634 .00	-1.214 .00	.362 .0005	.024 .00	.951
FIN'	Δb_{t-1}	-	-	-	1.376 .0001	-1.134 .001	.348 .005	-	-	-	-	.091 .0002	.526
POR	Δb_{t-1}		$y = .118b + 10.266$	Due to near-singular square matrix (i.e. it has a determinant of one and cannot be inverted) the model could not be established.									
	Δy_{t-1}												

Notes: y_t and b_t denote the logarithms of the GDP level and bond market capitalization at time *t*, respectively; *D* is the first-difference operator. Growth rates are calculated as $\Delta y_t = \ln y_t - \ln y_{t-1}$. *P*-values are reported below the corresponding coefficients.
DV: denotes dependent variable, ***a*** is the adjustment coefficient; *EC-term*: denotes co-integrating equation.

JPN: GDP growth adjusts to the long-run equilibrium state defined by the EC-term (the *a* is statistically significant). In the short run, there is evidence for bi-directional causality as all the coefficients on the bond market changes in the regression on GDP growth are statistically significant, and the coefficients on real output changes in the reverse regression are statistically significant also. Although ranked seventh with regard to its bonds/GDP ratio, the empirical evidence for Japan is that real growth follows the developments in the bond markets.

ITA: Only bond market growth converges to the long-run equilibrium (*a* is statistically significant). The lagged GDP changes in the regression on the bond market growth are also significant, which suggests both long-run and short-run causal relationship from GDP to the bond market. Turning to GDP growth, only the second bond market lag is statistically different from zero. These findings can be interpreted as evidence in favor of the demand-leading hypothesis.

FIN: The significant adjustment coefficient *a* in the regression on real GDP growth indicates long-run causality, whereas the bond market changes are not found to influence real GDP growth in the short-run. With bond market growth being the dependent variable, neither the adjustment coefficient nor the coefficients on the lagged changes are statistically significant. Therefore the EC-term is dropped from the regression and the model is re-estimated. It can be observed that bond market growth depends on own lags only.

POR: The model could not be set up, possibly due to the small number of observations (n = 24). The series on bond market capitalization starts in 1977 and start growing after 1987 (EU accession impact).

5.2. Evidence for short-run dynamics. One-way causality patterns

The evidence from the short-run autoregressive models supports the supply-leading hypothesis.

Table 8: Interdependence between bond market capitalization and GDP. Short-run VAR estimates

Country	***DV***	Db_{t-1}	Db_{t-2}	Dy_{t-1}	Dy_{t-2}	***Const***	R^2	***SSR***
AUT	Db_t	.42*** (2.64)	-	.04 (.06)	-	.07*** (2.57)	.178	.216
	Dy_t	.1*** (2.65)	-	.123 (.81)	-	.02*** (2.58)	.181	.010
USA	Db_t	.777*** (7.78)	-	-.187 (-1.21)	-	.015** (2.12)	.579	.024
	Dy_t	.164* (1.73)	-	-.027 (-.18)	-	.026*** (3.73)	.063	.022
SUI	Db_t	.009 (.06)	-	-.236 (-.43)	-	.044* (1.99)	.004	.504
	Dy_t	.089*** (2.69)	-	.524*** (4.35)	-	.009* (1.82)	.362	.024
E	Db_t	.254 (1.51)	-	-.259 (-.15)	-	.056 (.7)	.067	2.318
	Dy_t	-.014 (-1.217)	-	.663*** (5.77)	-	.012** (2.3)	.527	.010
UK	Db_t	.447** (2.27)	-	-1.273 (-1.46)	-	.058** (2.15)	.264	.123
	Dy_t	.096*** (2.6)	-	.53*** (3.27)	-	.006 (1.24)	.470	.004
GER	Db_t	.007 (.05)	.048 (.31)	-.488 (-1.30)	.608 (1.62)	.073*** (3.10)	.101	.115
	Dy_t	-.033 (-.503)	.162*** (2.45)	.356*** (2.19)	-.068 (-.42)	.009 (.93)	.244	.022
NL	Db_t	.121 (.81)	.306** (2.06)	-.675 (-.62)	-.233 (-.22)	.063 (1.47)	.148	.734
	Dy_t	.019 (.95)	-.036* (-1.78)	.469*** (3.16)	-.281* (-1.97)	.025*** (4.27)	.228	.014
FRA	Db_t	.102 (.43)	-.011 (.08)	-.708 (-.48)	1.715 (1.18)°°	.058 (1.27)°°	.091	.079
	Dy_t	-.013 (-.33)	.055 (.21)	.395 (1.58)°°°	-.049 (-.20)	.014 (1.83)**	.163	.002
SWE	Db_t	.111 (.55)°	.106 (.53)°	-2.798 (-1.25)°°	1.424 (.62)°	.075 (1.00)°°	.102	1.198
	Dy_t	.0006 (.034)	-.011 (-.661)°	.516 (2.649)***	-.296 (-1.48)°°°	.018 (2.72)***	.230	.009

Notes: y_t and b_t denote the logarithms of the GDP level and bond market capitalization at time *t*, respectively; *D* is the first-difference operator. Growth rates are calculated as $\Delta y_t = \ln y_t - \ln y_{t-1}$. *DV* is the dependent-variable column, followed by the lagged differenced terms and the constant. Standard errors are reported in parentheses below the corresponding coefficients. Triple, double and single asterisk denote significance at p=.01 (or .02); .05 and .10, respectively. "°°°", "°°" and "°" denote significance at p = .20; .40 and .60 respectively.

AUT: The coefficients of the bond market growth in the VAR system are statistically significant. The bivariate Granger causality test supports the hypothesis that GDP growth is caused by bond market development.

USA: The regressions support the hypothesis of supply-leading finance. The high bonds/GDP ratio (USA is ranked fourth) could be the ultimate cause for this observation.

SUI: Bond market capitalization changes seem to be independent from own lags or lagged output growth. The low R-squared and the high sum of squared residuals implies that the regression does not explain much of the variance in the growth of the bond market. On the other hand, real bond market capitalization is found to Granger-cause real economic activity. This is supported by the autoregressive representation, where all coefficients are statistically significant.

E: More than 50% of the variation in real economic growth can be explained by the VAR (1) model, although the bond market growth has a minimal explanatory power for real output growth.

UK: The residuals from the co-integrating regression with real bond market capitalization as response variable are stationary, whereas the residuals from the reverse regression are not. Since the critical values are computed for a sample size of 20 observations, they may not be reliable when only 16 observations are available (the ADF test has been performed with 8 lags). Thus it is ambiguous whether the series are co-integrated. Therefore the Johansen test has been performed on the levels of the series including two time lags as selected by the AIC. Both test statistics fail to reject the null of no co-integration. Based on these findings, the UK real output and the bond market are not co-integrated. A VAR with one lagged annual term on the changes is estimated. There is strong evidence for short-run causality from bond market growth to real economic growth, obtained both from the VAR(1) and the Granger causality test.

GER: There is supply-leading causality from the bond market development to real growth. The reverse causality from the real sector development to the bond market development is weak, as the coefficient on the second lag of the output changes is significant at the 20% level only.

The NL: There is only weak evidence for Granger-causality from bond market to real output growth. According to the autoregressive representations, bond market capitalization changes can be explained by own two-year lags, whereas changes in real economic activity are strongly related to own lags and to the second lag of bond market growth.

FRA and SWE: The VAR models are not significant. Causal links between bond markets and GDP growth cannot be analyzed with annual data.

Table 9: Granger causality tests within the VAR systems

Country	*Austria*	*USA*	*SUI*	*E*	*UK*	*GER*	*NL*	*FRA*	*SWE*
Model specification *Included observations*	VAR(1) n = 39	VAR(1) n = 47	VAR(1) n = 34	VAR(1) n = 36	VAR(1) n = 27	VAR(2) n = 21	VAR(2) n = 47	VAR(2) n = 21	VAR(2) n = 34
Null hypothesis H_0	*Probability for null rejection*								
Δy_t *does not cause* Δb_t	.952	.226	.671	.880	.145	.197	.740	.497	.443
Δb_t *does not cause* Δy_t	.008	.084	.007	.222	.009	.044	.175	.918	.803
Note: Bivariate tests for Granger causality within a VAR setting. Δy_t and Δb_t denote the growth rates of the GDP and the bond market capitalization, respectively.									

6. Summary and discussion

The causal patterns detected by the bivariate Granger causality test are summarized in table 10. Regarding the error-correction models it has to be noted, that although the Granger causality test does not support any causal patterns, in the presence of co-integration, there is causality in at least one direction (Granger, 1969). This applies to Japan, Italy and Finland. Since the Granger causality test is known for its sensitivity to lag length specification, and therefore its results need be interpreted in the context of the coefficients of the vector-autoregressive models. We were able to find a bi-directional causality, or interdependence between the bond sector growth and real growth, in the cases of Japan, Finland and Italy. There is support for supply-leading causality from bond market capitalization change to real growth in USA, Great Britain, Germany, Austria, Switzerland, and to a weaker extent in the Netherlands and Spain. Finally, we find no support for the reverse case, i.e. demand-leading causality from real economic activity to the bond market.

Table 10: Causality patterns

$\Delta y_t \leftrightarrow \Delta b_t$	$\Delta b_t \rightarrow \Delta y_t$	$\Delta b_t \rightarrow \Delta y_t$	$\Delta y_t \rightarrow \Delta b_t$	*Insufficiently determined*
Interdependence	*Supply-leading approach*	*Weak causality*	*Demand-leading approach*	
Japan	Switzerland	Netherlands	No evidence found	Portugal
Finland	Austria	Spain		France
Italy	UK			Sweden
	Germany			
	USA			

The causality patterns differ from country to country due to the heterogeneity of market structures and the different degree of openness and international integration of the capital markets in our country sample. The data available for the individual countries differs substantially across time: the Netherlands has the longest time series (n = 50 observations), whereas for Portugal and France we only dispose of 24 observations (data start in 1977). For these reasons the time series analysis is a more suitable approach than panel data methods, as these would constrict the time sample and lead to generalized assertions, which normally do not allow for country-specific inference.

The deficiency of the for some countries relatively short time sample is that it cannot be split to test for structural breaks and linearity in the relationships found. The latter is desirable considering the Portuguese and Spanish financial market development from autarky to more integrated financial sector after the EU entry in 1986. Utilization of higher frequency data instead of the annual data used in this study would offer a possibility to reject our findings or to find even stronger causal ties. To test the obtained results for robustness, in a next step the same tests may be performed on the growth rates calculated as percentages instead on the first differences. Our research will also proceed by taking into account possible different financial sector influences on real economic activity. This will be done by splitting the aggregate bond market into corporate and government bonds. To make the picture complete, a further step could be to estimate the same relationship augmented by the influence of stock markets and bank credit on economic growth.

As the log levels of the variables under investigation are co-integrated, the examined relationships imply level effects and not permanent growth effects. For the second to be true, the growth rate of real output needs be co-integrated with the level of the bond market capitalization. This case necessitates that the growth rate of output has the same order of integration as the level of real bond market capitalization, which is ruled out by the finding that both variables are difference stationary.

The same research for the Central and Eastern European countries would convey very interesting insights for comparative and policy-making purposes. Due to unavailability of sufficiently long time series, this could not be fulfilled here, but this perspective provides also a positive and practical step for future research in international finance.

7. Conclusions

This study has presented first evidence on the relationship between bond markets and real economic growth in the EU15, USA, Japan, Switzerland and Norway. The employed econometric techniques take into account stationarity, co-integration, and causality features of the variables of interest.

When applied to aggregate domestic bond markets in most of the EU member states and USA, Japan, Switzerland and Norway, the techniques we use deliver well-specified relationships backed by an intuitive economic interpretation. The estimated models support the supply-leading hypothesis implying that real economic activity is significantly influenced by the development of the bonds market.

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Data Appendix

The USA, Japan, Switzerland, Norway and the 15 EU member countries are individually examined. The aggregate bond market capitalization is drawn as a measure for the size and volume of bond markets. The nominal data are deflated by the means of the GDP deflator. The surveys on the impact of the bond markets on GDP growth are conducted in real terms in order to filter out possible inflationary effects. GDP data are from the IFS of the IMF, and the OECD Historical Statistics (1960-1997, 1998 Paris) and National Accounts of OECD Countries, Volume 1 (Main Aggregates, Paris, diverse issues), and are drawn from the WIFO database (http://www.wifo.at/db/index.html). The nominal GDP time series are from the IFS (99B). The real GDP series until 1969 are converted by transforming the nominal GDP from the IFS with the GDP-deflator from the IFS (99BVP). The real GDP series from 1970 to 1999 are drawn from the OECD National Accounts (FIN"B1_GE_VOB). The OECD Historical Statistics index is used for four countries: Norway (till 1965), Austria (till 1963), Denmark (till 1965), and Portugal (till 1968).

The data on bond markets capitalization from 1990 to 2000 comes from the BIS. Data until 1989 is composed as follows:

Table A1: Sources of bond market capitalization data before 1990

Country	*Time horizon*	*Series*	*Source*
USA	1950-1989	All bond types	BIS (EDBAUS01: securities outstanding) *
	1990-2000		BIS
JPN	1963-1989	All bond types	IFS (row 26AB) *
	1990-2000		BIS
NOR	1974-1989	Financial, corporate & foreign bonds	IFS *
	1974-1989	Public sector	GFS *
	1990-2000	All bond types	BIS
SUI	1950-1959	Bonds	IFS (row 26) *
	1960-1989	Financial, corporate & foreign bonds	IFS *
	1960-1989	Public sector	GFS (total public debt by type of instrument) *
	1990-2000	All bond types	BIS
AUT	1953-1973	Bonds	IFS (row 26) *
	1974-1989	Financial & corporate sector	IFS *
	1974-1989	Public sector, foreign bonds	GFS (total foreign debt) *
	1990-2000	ALL BOND TYPES	BIS
BEL	1980-1989	All bond types	BIS (EDNABE01: securities outstanding, bonds, total)*
	1990-2000		
FIN	1972-1989	All bond types	BIS (EDNAFI91: securities outstanding, bonds, total)*
	1990-2000		
FRA	1977-1989	Financial & corporate sector	IFS *
	1977-1989	Public sector	GFS (total domestic debt) *
	1990-2000	All bond types	BIS
GER	1950-1989	Financial, corporate & foreign bonds	IFS
	1950-1989	Public sector	GFS (1950-1974: total government debt; 1975-1989: total public debt)
	1990-2000	All bond types	BIS
ITA	1963-1977	All bond types	IFS *
	1978-1989	Financial, corporate & foreign bonds	IFS *
		Public sector	GFS (domestic government bonds & bills) *
	1990-2000	All bond types	BIS
NL	1951-1964	All bond types	GFS (domestic government securities) *
	1965-1989	Financial, corporate & foreign bonds	IFS *
	1965-1989	Public sector	GFS *
	1990-2000	All bond types	BIS
POR	1977-1989	All bond types	IFS *
	1990-2000		BIS
E	1963-1989	Financial, corporate & foreign bonds	IFS *
	1963-1989	Public sector	GFS (domestic public debt) *
	1990-2000	All bond types	BIS
SWE	1964-1969	All bond types	IFS *
	1970	All bond types	GFS (domestic public debt) *
		Financial, corporate & foreign bonds	IFS (row 46AB) *
	1971-1989	Public sector	GFS (domestic public debt) *
	1990-2000	All bond types	BIS
UK	1976-1989	All bond types	National sources (outstanding amounts of government bonds) *
	1990-2000		BIS

Note: Asterisk denotes series that partly consist of original data in value terms. Missing data have been calculated with the help of index time series which were made available mostly by the same reporting institution.