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New evidence on assessing the level of competition in the European Union banking sector: A panel data approach

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ABSTRACT

The goal of this paper is to empirically assess the level of banking competition in the European Union (EU) across three economic blocks (i.e. EU-27, EMU-17 and the remaining EU countries). Furthermore, the paper assesses the impact of the on-going financial crisis (2008–2011) on the competition pattern of the banking sector in the European Monetary Union (EMU) as a whole, where little attention has been paid by the relevant literature. The analysis employs the Panzar and Rosse ((1987). *Journal of Industrial Economics*, 35, 443) methodology and draws upon a panel dataset of EU banks, spanning the period 1996–2011. The empirical findings are robust, providing updated evidence in favour of a monopolistic competition pattern across all EU economic blocks examined. The level of competition in the EMU countries triggered by bank consolidations seems to have undergone a small, albeit a significant decline, after the adoption of the euro currency and the on-going financial crisis.

1. Introduction

There is no doubt that the banking sector constitutes one of the most important sectors of the EU economy, since it represents over 50% of total EU activity in terms of gross income (EC, 2007). It is stated that in 2004, retail banking activity in the EU generated gross income of 250–275 billion euros, equivalent to approximately 2% of the total EU GDP. The sector is also critical for the competitiveness, economic growth and prosperity of the EU since it has significant spillovers on all other economic activities. However, a number of idiosyncratic characteristics, such as market fragmentation, price rigidity and customer immobility, suggest that competition in the EU retail banking market may be hindered (EC, 2007). Therefore, the investigation of the level of competition in the EU banking sector is a rather crucial issue, with important economic and managerial implications.

Many empirical studies have attempted to examine the competitive conditions in the banking sector and its specific market structure (i.e., oligopoly, monopolistic competition, monopoly, perfect competition). The majority of these studies consent that banks operate in a monopolistic competitive environment. The theory of the monopolistic competition suggests that firms compete by offering differentiated products (Chamberlin, 1933). In this context, a firm operates as a price-taker since there are many producers in the relevant market and none of them is able to set his own price (Chamberlin, 1933). Consequently, each firm has limited, if any, control over the final price of its offerings. By contrast, consumers perceive that there are non-price differences among the competitors' goods, while there are few barriers to entry and exit the market, at least in the long-run. However, producers have to some extent control over the market price.

It is worth emphasising that many markets are dominated by monopolistic competition characteristics (i.e., advertising, hotel and restaurants, insurance). In terms of the banking sector, the knowledge on the level of monopolistic competition is a crucial and important issue not only from a policy making perspective, but also from a managerial standpoint. This is justified inter alia by two reasons. First, the presence of low entry barriers in tandem with a low level of Significant Market Power (SMP) in the sector might affect managerial decisions towards their engagement into strategic alliances and mergers and acquisitions. Second, the

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presence of differentiated products may induce bank managers to expand their sources of earnings through diversification of assets and liabilities, as well as by reducing the operational cost and/or increasing non-interest revenues (Andries & Capraru, 2014).

From a methodological perspective, over the last decades, two non-structural models of competitive behaviour have been developed within the emerging New Empirical Industrial Organization (NEIO) framework. These models measure competition and focus on the detailed competitive conduct of firms without using explicit information on the structure of the market (Bresnahan, 1982; Panzar & Rosse, 1987). Both models measure competitive conditions by estimating deviations from competitive pricing and can be formally derived from profit maximising equilibrium conditions, which are their main advantage over structural measures (Bikker, Shaffer, & Spierdijk, 2012).

In the empirical banking literature, the widely used Panzar–Rosse model builds a competition indicator, the so-called *H*-statistic, which provides a quantitative assessment of the competitive nature of a market. The *H*-statistic is calculated by means of reduced-form revenue equations and measures the elasticity of total revenues with respect to changes in factor input prices (Panzar & Rosse, 1987). This methodology, based on four steps (Fig. 1), makes use of bank level data. It examines the extent to which a change in factor input prices is reflected on (equilibrium) revenues earned by a specific bank. Under perfect competition, an increase in input prices leads to proportional increases of both marginal costs and total revenues. Under monopolistic conditions, an increase in input prices will increase marginal costs and will reduce equilibrium output, thus, consequently, total revenues. A value below zero denotes a collusive (joint monopoly) competition; a value bellow one denotes the presence of monopolistic competition; and a value equal to one characterizes perfect competition. Furthermore, Shaffer (1982) shows that *H* is negative for a conjectural variations' oligopolistic market or for a short-run competitive market; it is equal to one for a natural monopoly in a contestable market; or, it is equal to zero for a firm that maximises sales subject to a breakeven constraint.

The advantage of this methodology is that it uses bank-level data and allows for bank-specific differences in production. However, the methodology does not allow the study of explicit differences across different banks, e.g. large versus small or foreign versus domestic banking institutions, since the *H*-index cannot be interpreted as an ordinal statistic (Bikker et al., 2012).There is a striking dichotomy between the reduced form of the price/revenue relationship, as estimated in the empirical literature. Some researchers estimate a price or a revenue function that does not include total banking assets as a control variable (Bikker, Spierdijk, & Finnie, 2006, 2012; Polemis, 2014). Others, estimate a price/revenue function in which the dependent variable is either the gross interest revenues or the total banking revenues divided by total assets (Bikker & Haaf, 2002; Claessens & Laeven, 2004; Mamatzakis, Staikouras, & Koutsomanoli-Fillipaki, 2005; Yildirim and Phillipatos, 2007Yildirim & Phillipatos, 2007). It is noteworthy that Bikker et al. (2006, 2012) show that both the price and the scaled revenue equations lead to a biased estimate of the *H*-index. The misspecification is due to the use of the bank revenues divided by total assets as a dependent variable instead of the unscaled bank revenues. This finding has important consequences, given that the *H*-indices cannot be reliably used as a measure of the degree of competition; moreover, various conditions can cause a reverse of the sign of values, regardless the degree of competition (Bikker et al., 2012). In order to overcome these problems and strengthen our findings, this paper makes use of both scaled and unscaled price and revenue equations as a robustness check to assess the degree of competition in the European Union (EU) banking sector.

The contribution of this paper is four-fold. First, it goes beyond the current literature, in a sense that it attempts to assess the level of banking competition across three economic blocks (i.e., EU-27, EMU-17 and the remaining EU countries). Next, our research focuses on the competitive conditions prevailing in the EMU by

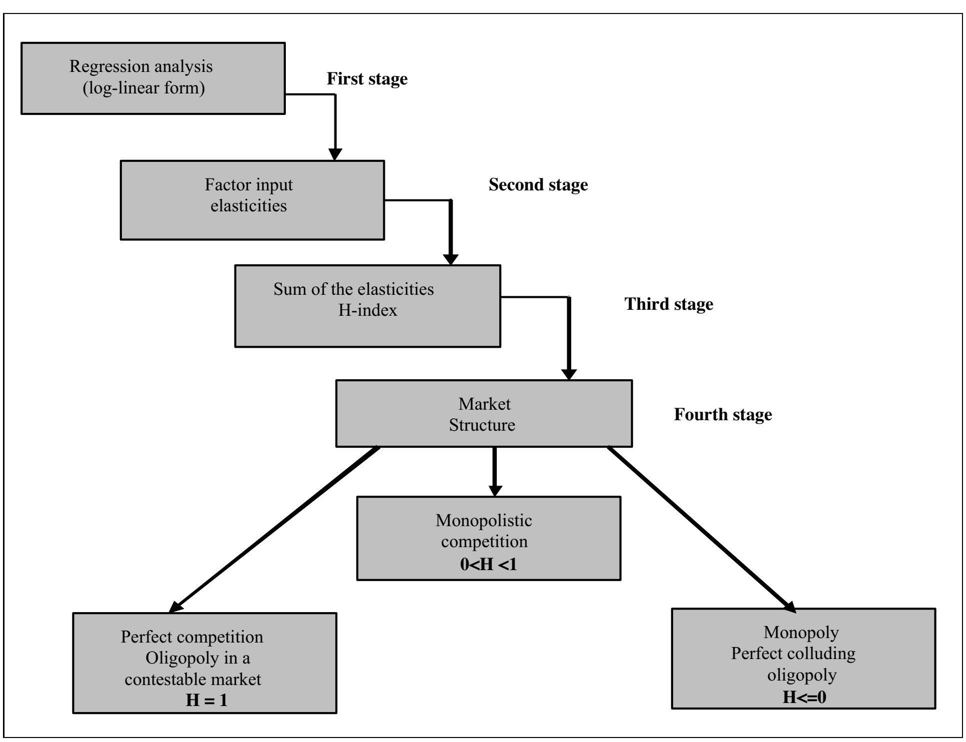

Fig. 1. Extraction and interpretation of the P–R index. *Source*: Authors' elaboration.

splitting the sample into three distinct time periods (i.e., 1996–2000, 2001–2007 and 2008–2011). The relevant time intervals examine the competitive conditions of the EMU banking sector on specific economic events (i.e., conditions before and after the adoption of Euro, and the first four years of the on-going financial crisis). Second, it confirms earlier studies in a more elaborated manner by using sophisticated econometric methodologies, to test the robustness of the empirical findings. Third, the paper contributes to the advancement of the empirical banking literature by providing evidence on the evolution of banking competition in the EMU-17 during the first years of the on-going financial crisis (i.e., 2008–2011 and given the availability of the dataset). Fourth, our findings enhance the managerial decisions in the banking sector, by providing a more explicit and elaborated framework of existing competition in this business. They also provide further business insight to managers and policy makers in the banking sector.

The remainder of the paper is structured as follows: Section 2 reviews the literature, while Section 3 discusses the data and methodologies applied. Section 4 illustrates and evaluates the results of the empirical analysis. Section 5 concludes the paper and offers some managerial and policy-making recommendations.

2. Review of the literature

A number of empirical studies have investigated the competitive conditions in various banking systems by applying the Panzar–Rosse *H*-index (Table 1). The majority of these studies conclude that banks operate in a monopolistic competitive environment (Shaffer, 2002; Claessens & Laeven, 2004; Beck, Demirguc-Kunt, & Levine, 2006; Gutierrez, 2007), while European banks seem to be less competitive than U.S. banks, with larger banks being more competitive than smaller banks (Gutierrez, 2007).

Another stream of scholars indicates differences in the level of competition across banks operating in different countries. De Bandt and Davis (2000) provide supportive evidence in favour of a monopolistic behaviour of small banks in France and Germany, while they find that monopolistic competition prevails both in the case of small banks in Italy and in the case of large banks across all countries in their sample. Their findings suggest that small banks, at least in their country sample, have more market power as they cater more to local markets.

Claessens and Laeven (2004) compute the *H*-statistic for 50 developed and developing countries over the period 1994–2001. According to their results, monopolistic competition describes best the markets under consideration. Subsequently, they draw attention on the factors underlying competition by means of regressing the estimated *H*-statistics on a number of country-specific characteristics, such as the role of foreign banks, activity restrictions, market entry conditions, market structure, the competition from non-banking sectors, general macroeconomic conditions, and the overall development level of the country. Their findings do not document any direct relationship between competition and concentration, but they find that fewer entry and activity constraints, i.e. higher contestability, result in higher competition.

Staikouras and Koutsomanoli-Fillipaki (2006) carry out the first multi-country analysis for the EU, following the enlargement to 25 member countries and spanning the period 1998 to 2002. They find evidence of monopolistic competition, with larger banks behaving more competitively than smaller ones and with new members showing higher levels of competition than older banks.

Overall, the majority of the empirical studies seem to provide strong evidence supporting the hypothesis that monopolistic competition is the prevailing environment across European banks (Bikker & Haaf, 2002; Bikker et al., 2006). In fact, monopolistic competition is quite a recurrent finding due to the wide range of values the *H*-statistic can take within this scenario (i.e., between zero and one). This context enhances the importance of certain methodological issues concerning the empirical implementation of

Table 1
Main empirical studies.

Author	Time period	Model	Countries	Main findings
Bikker and Groeneveld (2000)	1989–1996	Scaled revenue equations	5 EU countries	Monopolistic competition in all of the countries
De Bandt and Davis (2000)	1992–1996	Scaled revenue equations	Germany, France and Italy	Monopolistic competition for large banks in all of the countries. Monopolistic behaviour for small banks in France and Germany Monopolistic competition for small banks in Italy
Bikker and Haaf (2002)	1988–1998	Scaled price and revenue equations	23 Countries	Perfect competition for the large banks Monopolistic competition for the small and medium size banks
Claessens and Laeven (2004)	1994–2001	Scaled price and revenue equations	50 Countries	Monopolistic competition in all of the examined countries
Weill (2004)	1994–1999	Scaled revenue equations	5 EU countries	Monopolistic competition in all of the examined countries
Mamatzakis et al. (2005)	1998–2002	Scaled revenue equations	Bulgaria, Croatia, and FYROM	Monopolistic competition
Staikouras and Koutsomanoli-Fillipaki (2006)	1998–2002	Scaled revenue equations	EU-15 EU-10 (enlargement countries)	Monopolistic competition in the EU-15 (*H* index = 0.54)Monopolistic competition in the EU-10 (*H* index = 0.78)
Casu and Girardone (2006)	1997–2003	Scaled revenue equations	5 EU countries	Monopolistic competition in all of the examined countries
Gutierrez (2007)	1986–2005	Scaled revenue equation	Spain	Monopolistic competition More competitive environment among larger banks
Sun (2011)	1995–2009	Scaled price equation	EMU, US and United Kingdom	Monopolistic competition in all of the examined countries and regions
Bikker et al. (2012)	1986–2004	Scaled and unscaled price and revenue equations	67 Countries	Monopolistic competition in 40 countries

the Panzar and Rosse (1987) approach such as, inter alia, data, estimation techniques, and sample period under consideration.

3. Methodology and data

In this section we discuss the empirical model used to assess the level and the impact of competition on banks' efficiency measures across the EU banking sector.

3.1. The empirical model

Following the methodology of previous empirical studies (Molyneux, Lloyd-Williams, & Thornton, 1994; Bikker & Groeneveld, 2000; Bikker & Haaf, 2002; Claessens & Laeven, 2004; Yildirim & Phillipatos, 2007; Schaeck, Cihak, & Wolfe, 2009), we estimate the following reduced-form revenue equations[1].

$$\ln(P_{it}) = a + \beta_1 \ln(FUND_{it}) + \beta_2 \ln(WAGE_{it}) + \beta_3 \ln(CAP_{it}) + \gamma_1 \ln(LEV_{it}) + \gamma_2 \ln(RISK_{,it}) + \gamma_3 \ln(SIZE_{,it}) + \epsilon_{it} \quad (1)$$

$$\ln(ROA_{it}) = a + \beta_1 \ln(FUND_{it}) + \beta_2 \ln(WAGE_{it}) + \beta_3 \ln(CAP_{it}) + \gamma_1 \ln(LEV_{it}) + \gamma_2 \ln(RISK_{,it}) + \gamma_3 \ln(SIZE_{,it}) + \epsilon_{it} \quad (2)$$

$$\ln(Z_{it}) = a + \beta_1 \ln(FUND_{it}) + \beta_2 \ln(WAGE_{it}) + \beta_3 \ln(CAP_{it}) + \gamma_1 \ln(LEV_{it}) + \gamma_2 \ln(RISK_{,it}) + \gamma_3 \ln(SIZE_{,it}) + \epsilon_{it} \quad (3)$$

where α and ε_{it} are the constant and the error terms, respectively. All the relevant variables are on an annual basis. We have to stress, however, that due to absence of reliable quarterly or monthly data concerning our key variables, we focus solely on annual observations despite the fact that annual based datasets have a number of shortcoming (i.e., lower information on the data generation process and lack of consistency). All variables are in their natural logarithms. The data are obtained from the Bankscope database of Bureau van Dijk. This database reports published financial statements from financial institutions worldwide. These data are reported in euros and are homogenized in order to be comparable and therefore suitable for a panel approach (Mamatzakis et al., 2005; Andries and Capraru, 2014). It is worth mentioning that nearly all of the current empirical studies that measure the level of competition in the banking industry worldwide use this specific database.

The analysis employs a widely used non-structural methodology put forward by Panzar and Rosse (1987) and draws upon a comprehensive panel dataset of EU commercial and savings banks, spanning the period 1996–2011. The starting date for the empirical analysis was dictated by data availability. However, we must bear in mind that this could not raise any issue regarding the sample selection since minimal reforms in the EU banking sector occurred before prior to this date. The final date represents the last year for which data are available at the time the research was conducted. In addition, two were the main economic reasons for choosing the specific time period. First, unlike previous studies, we needed to assess the impact of the first years of the on-going financial crisis (i.e., 2008–2011) on the level of banking competition. Second, in order to measure the impact of the adoption of the common currency (euro) and its consequences to the level of financial stability within the EMU countries, we split the original sample into two sub-periods, accounting for the pre-EMU (1996–2000) and the post-EMU (2001–2007) period, respectively.

Summary statistics for the variables included in the empirical analysis are provided in Appendix A (Tables A1 and A2). From the relevant tables, it is evident that the sample data are well behaved showing limited variability in relation to the mean. On the other hand, the variables are not normally distributed, since the relative values of the skewness and kurtosis measures are not equal to zero and three, respectively.

Despite the great number of empirical studies on the topic, none of them, to the best of our knowledge, has investigated the competition level of the banking sector in the EU countries as a whole, by pooling the relevant banking variables (i.e., gross interest revenues, total assets, deposits, personnel expenses) across all member states (27 countries)[2]. The comparative advantage of the methodological approach of this paper vis-à-vis other traditional approaches (Andries and Capraru, 2012) is that it provides efficient estimators, despite the particular differences across the sample countries (Greene, 2000). However, it should be noted that many EU banks operate branches in other EU countries (overseas branches). Therefore, to avoid the double counting problem when we pool our data, we delete the branches that the EU banks operate in other EU countries in the estimations for the overall sample and the sub-samples, respectively. Since consolidated accounts are available, this recollection method is both reasonable and less complicated, given the data restrictions (Sun, 2011).

Furthermore, it should be stressed that any accounting differences across different EU countries, as well as over time, may affect comparability of the accounting data[3]. These differences in the banking accounting standards may influence the significance of losses/costs or gains in their balance sheets, which in turn may possibly distort the comparisons of the *H*-indices. In order to deal with this issue, we normalise the control variables of the empirical models with the total banking assets. Therefore, the pooled factor cost estimators (elasticities), used to construct the *H*-index, seem to be appropriate for our analysis.

P_{it} is the ratio of gross interest revenues to total assets as used a proxy for the output price of loans; $FUND_{it}$ is the ratio of interest expenses to total deposits and money market funding as a proxy for the average funding rate; $WAGE_{it}$, is the ratio of personnel expenses to total assets as an approximation of the wage rate; and, CAP_{it} is the ratio of other operating and administrative expenses to total assets as a proxy for the price of physical capital. Moreover, LEV_{it} is the ratio of equity to total assets as a proxy for the leverage, reflecting differences in the risk preferences across banking institutions (Bikker et al., 2012); $RISK_{it}$ is the ratio of net loans to total assets as an approximation of the credit risk; and, $SIZE_{it}$ represents total banking assets to control for the potential size effect[4]. It is noteworthy that controlling for scale is crucial since larger banking institutions earn more revenues in ways unrelated to variation in input prices. Therefore, if we estimate a reduced form revenue equation across banking firms of different sizes without controlling for scale, the standard measures of fit will be quite poor (Bikker et al., 2012).

ROA_{it} is the pre-tax return on assets and Z_{it} is the ratio of total revenues, including gross interest revenues plus other operating revenues to total assets. The sum of the three elasticities $\left(\sum_{i=1}^{3} \beta_i\right)$

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[1] In the case of the PGLS methodology, we allow fixed effects only for the sample countries (cross section dimension).

[2] The majority of the empirical studies (Molyneux et al., 1994; Bikker & Groeneveld, 2000; Bikker & Haaf, 2002; Mamatzakis et al., 2005; Bikker et al., 2012) measure the level of banking competition by estimating the *H*-index in each EU country. However, Andries and Capraru (2012, 2014) investigate competition in the banking sector of the EU-27 as a whole by estimating the mean value of the relevant *H*-indices for each member state.

[3] The Continental model (e.g., Italy, Greece, Spain, Portugal, etc.) focuses on debt holders, while the Anglo Saxon model (e.g., United Kingdom, Ireland, etc) favours share holders (Sun, 2011).

[4] In the case of the unscaled equations (Eqs. (4) and (5)) the control variable $SIZE_{it}$ represents the ratio of other non-earning assets to total assets as a proxy for the asset composition of banks notated by the variable $ASSET_{it}$ (Bikker et al., 2012).

yields the *H*-index in Eqs. (1) and (3). The reason for including Eq. (2) in the estimation process is that the P-R model is only valid if the market is in equilibrium (Claessens & Laeven, 2004). In this case, we define the equilibrium *E*-statistic as the sum of the three elasticities $\sum_{i=1}^{3}\beta_i$. In the next step, we test whether $E = 0$ using a Wald test. If rejected, the market is assumed not to be in equilibrium. The reason is that in free-entry equilibrium across homogenous banks, demand and supply forces will equalize ROA across banks, so that the level of ROA turns to be independent of input prices (Shaffer, 1982). However, Bikker et al. (2012) show that the ROA test is actually a joint test for competitive conduct and long-run structural equilibrium. They show that ROA would not equal zero, even if the market is in structural equilibrium. Because of the joint character of the test, it is difficult to interpret it by narrowing its applicability.

Bikker et al. (2012) point out that the scaled revenue and price functions (including total assets as a control variable) can lead to overestimation of the degree of competition in the banking industry. In order to check for the robustness of our findings, we estimate the *H*-indices generated from the following unscaled models (Eqs. (4) and (5)).

$$\ln(GIR_{it}) = a + \beta_1\ln(FUND_{it}) + \beta_2\ln(WAGE_{it}) + \beta_3\ln(CAP_{it}) + \gamma_1\ln(LEV_{it}) + \gamma_2\ln(RISK_{,it}) + \gamma_3\ln(ASSET_{,it}) + \epsilon_{it} \quad (4)$$

$$\ln(TR_{it}) = a + \beta_1\ln(FUND_{it}) + \beta_2\ln(WAGE_{it}) + \beta_3\ln(CAP_{it}) + \gamma_1\ln(LEV_{it}) + \gamma_2\ln(RISK_{,it}) + \gamma_3\ln(ASSET_{,it}) + \epsilon_{it} \quad (5)$$

where GIR_{it} is gross interest revenues (or interest income) and TR_{it} is total revenues (or total income) expressed as the sum of gross interest revenues plus other operating non-interest revenues.

The econometric methodology adopted in this paper uses two different sets of estimators. First, we assess the level of competition by using panel GLS fixed effects estimator (PGLS) that allows the unobserved country-specific factors to be filtered out. Thus, to check the robustness of our findings, we re-estimate the price and revenue equations by employing a Fully Modified OLS estimator (FMOLS) pioneered by Pedroni (2000). This methodology controls for the endogeneity that may arise in standard estimation methods, (e.g., OLS) often employed in practice (Apergis & Polemis, 2015). The latter can be a problem because, if unobserved variables jointly affect both the dependent and control variables, and in that case the coefficient estimates for the independent variables may be biased (Hausman and Ros, 2013). This estimator takes into account the unobserved time-invariant bilateral specific effects, while it can effectively deal with the potential endogeneity arising from the inclusion of several control variables (Polemis, 2015).

4. Results and discussion

4.1. Cross sectional dependence tests

We carry out the first part of the empirical analysis by examining the presence of cross-sectional dependence. Panel unit root tests of the first-generation can lead to spurious results (because of size distortions) if significant degrees of positive residual cross-section dependence exist and are ignored. Consequently, the implementation of second-generation panel unit root tests is desirable only when it has been established that the panel is subject to a significant degree of residual cross-sectional dependence. In the cases where cross-section dependence is not sufficiently high, a loss of power might result if second-generation panel unit root tests that allow for cross-section dependence are employed. Therefore, before selecting the appropriate panel unit root test, it is crucial to provide some evidence on the degree of residual cross-section dependence.

Table 2
Cross-section dependence (CD) test-cross-section correlations of the residuals in ADF(p) regressions.

	Lags			
Variables	1	2	3	4
P	[0.00][a]	[0.00][a]	[0.02][b]	[0.03][b]
ROA[+]	[0.00][a]	[0.00][a]	[0.01][a]	[0.00][a]
Z	[0.00][a]	[0.02][b]	[0.02][b]	[0.00][a]
GIR[+]	[0.00][a]	[0.00][a]	[0.03][b]	[0.02][b]
TR[+]	[0.01][a]	[0.03][b]	[0.01][a]	[0.04][b]
FUND	[0.03][b]	[0.00][a]	[0.02][b]	[0.02][b]
WAGE	[0.00][a]	[0.01][a]	[0.02][b]	[0.01][a]
CAP	[0.00][a]	[0.00][a]	[0.02][b]	[0.02][b]
LEV	[0.02][b]	[0.00][a]	[0.02][b]	[0.01][a]
RISK[+]	[0.01][a]	[0.00][a]	[0.01][a]	[0.01][a]
SIZE	[0.02][b]	[0.00][a]	[0.02][b]	[0.02][b]

Notes: Under the null hypothesis of cross-sectional independence the CD statistic is distributed as a two-tailed standard normal. Results are based on the test of Pesaran (2004). Figures in parentheses denote *p*-values. Significance levels: [a] (1%) and [b] (5%). Due to space limitations, we provide only the results for the SIZE and not for the ASSET variable.

The cross-sectional dependence (CD) statistic by Pesaran (2004) is based on a simple average of all pair-wise correlation coefficients of the OLS residuals obtained from standard augmented Dickey and Fuller (1979) regressions for each variable in the panel. Under the null hypothesis of cross-sectional independence, the CD test statistic follows asymptotically a two-tailed standard normal distribution. The results, reported in Table 2, uniformly reject the null hypothesis of cross-section independence, providing evidence of cross-sectional dependence in the data given the statistical significance of the CD statistics, regardless of the number of lags (from 1 to 4) included in the ADF regressions.

4.2. Stationarity tests

Two second-generation panel unit root tests are employed to determine the degree of integration in the respective variables. The Pesaran (2007) panel unit root test does not require the estimation of factor loading to eliminate cross-sectional dependence. Specifically, the usual ADF regression is augmented to include the lagged cross-sectional mean and its first difference to capture the cross-sectional dependence that arises through a single-factor model. The null hypothesis is a unit root for the Pesaran (2007) test. The bootstrap panel unit root tests by Smith, Leybourne, and Kim (2004) utilize a sieve sampling scheme to account for both the time series and cross-sectional dependence in the data through bootstrap blocks. All four tests by Smith et al. (2004) are constructed with a unit root under the null hypothesis and heterogeneous autoregressive roots under the alternative hypothesis. The results of these panel unit root tests are reported in Table 3 and support the presence of a unit root in all variables under consideration.

The next step is to ensure that Eqs. (1) through (5) represent valid long-run relationships, given that previous tests indicate the presence of unit roots. Table 4 reports Pedroni's (2000) panel cointegration test statistics. All seven test statistics reject the null hypothesis of no cointegration at the 1 percent significance level[5].

[5] Due to space limitations, we present only the results regarding the variables in the scaled models (Eqs. (1) and (3)). However, according to the panel cointegration test statistic, we find evidence that the variables in the unscaled models are also cointegrated.

Table 3
Panel unit root test results for the EU-27.

Variable	Pesaran CIPS	Pesaran CIPS*	Smith et al. *t*-test	Smith et al. LM-test	Smith et al. Max-test	Smith et al. Min-test
Levels						
Dependent variables						
P	−1.55	−1.35	−1.37	3.31	−1.26	1.32
ROA⁺	−1.28	−1.42	−1.25	2.35	−1.72	1.44
Z	−1.41	−1.34	−0.86	1.29	−1.25	1.45
GIR⁺	−1.27	−1.48	−1.35	3.26	−1.41	1.50
TR⁺	−0.85	−1.26	−1.29	4.14	−1.35	1.36
Control variables						
FUND	−1.42	−0.85	−1.17	1.57	−1.19	1.46
WAGE	−1.36	−0.92	−1.18	1.68	−1.25	1.32
CAP	−1.39	−0.82	−0.93	1.82	−1.29	1.39
LEV	−0.69	−1.12	−0.71	3.36	−1.42	1.14
RISK⁺	−1.43	−1.42	−1.04	3.49	−1.27	1.26
SIZE	−1.25	−1.37	−1.15	3.75	−1.53	1.45
First differences						
Δ(*P*)	−5.71*	−6.21*	−7.25*	32.56*	−6.65*	7.61*
Δ(*ROA*)⁺	−6.18*	−5.63*	−6.23*	17.34*	−6.14*	5.48*
Δ(*GIR*)⁺	−5.46*	−5.92*	−5.47*	22.67*	−5.48*	6.72*
Δ(*TR*)⁺	−5.92*	−5.91*	−5.42*	21.52*	−7.62*	7.15*
Δ(*Z*)	−6.45*	−5.44*	−7.61*	31.51*	−6.25*	6.39*
Δ(*FUND*)	−6.19*	−6.31*	−5.60*	22.18*	−6.60*	6.18*
Δ(*WAGE*)	−5.46*	−6.15*	−6.37*	25.64*	−7.15*	5.63*
Δ(*CAP*)	−5.64*	−5.82*	−6.75*	25.72*	−7.58*	6.58*
Δ(*LEV*)	−6.26*	−6.15*	−6.28*	25.24*	−7.50*	6.42*
Δ(*RISK*)⁺	−5.72*	−5.75*	−5.38*	18.69*	−7.16*	6.59*
Δ(*SIZE*)	−7.53*	−5.81*	−6.62*	19.57*	−5.73*	7.24*

Notes: Δ denotes first differences. A constant is included in the Pesaran (2007) tests. Rejection of the null hypothesis indicates stationarity in at least one country. CIPS* = truncated CIPS test. Critical values for the Pesaran (2007) test are −2.40 at 1%, −2.22 at 5%, and −2.14 at 10%, respectively.
* Denotes rejection of the null hypothesis. Both a constant and a time trend are included in the Smith et al. (2004) tests. Rejection of the null hypothesis indicates stationarity. For both tests the results are reported at lag = 4. The null hypothesis is that of a unit root. Due to space limitations, we provide only the results for the *SIZE* and not for the *ASSET* variable.

4.3. Estimation results

In this section, we present the results from the empirical analysis. For the sake of simplicity, we divide our analysis into two sub-sections. In the first section, we present the empirical results generated from the scaled models (Eqs. (1) and (3)) along with the equilibrium test (Eq. (2)). In order to check for the robustness of our results, we thoroughly compare and contrast the results from the scaled equations with the ones drawn by means of using the unscaled models.

4.3.1. Scaled equilibrium models

From the estimation results, it is obvious that the coefficients are statistically significant, the signs are the expected ones and the fit to the data is more than satisfactory (Table 5)[6]. The *H*-index, which is calculated as the sum of the three elasticities $\left(\sum_{i=1}^{3} \beta_i\right)$, is less than one across all specifications, implying that monopolistic competition is the best description of the extent of competition in the EU banking sector. No significant variation seems to exist across the three specifications.

It is worth noting that the magnitude of the *CAP* variable is lower compared to other two input price variables (*FUND* and *WAGE*, respectively) across the majority of specifications. In other words, the price of physical capital contributes less to the equation for the *H*-statistic than the remaining input prices. This finding is in alignment with similar studies (Molyneux et al., 1994; Molyneux, Lloyd-Williams, & Thornton, 1996; Bikker & Haaf, 2002; Delis, 2010), denoting that excess physical capital, including branches, does not generate abnormal revenues (Delis, 2010).

From the Wald tests, testing for $H = 0$, we conclude that only in the second specification (*E*-equilibrium statistic), the null hypothesis ($E = 0$) cannot be rejected. Therefore, we conclude that the banking sector is in long-run equilibrium and returns on bank assets should be related to input prices (Shaffer, 1982). However, according to Bikker et al. (2012), failure to reject $E = 0$ and $0 \leq H < 1$ does not necessary coincide with the presence of a long-run competitive equilibrium due to the misspecification of the test (i.e., presence of large standard errors). Therefore, we must be very cautious when interpreting the ROA test.

The empirical results seem to be quite robust and in alignment with other empirical studies (Claessens & Laeven, 2004; Weill, 2004; Mamatzakis et al., 2005; Staikouras and Koutsomanoli-Fillipaki, 2006; Casu & Girardone, 2006; Gutierrez, 2007), providing sufficient evidence of monopolistic competition. The relatively high level of the *H*-index in the EU-27 compared to the EMU-17 could be potentially explained by a number of policy initiatives, including the creation of an integrated market for banking services. In particular, certain regulatory restrictions have been removed over the last decade. This development triggered the consolidation and outward expansion processes across commercial banks, enhancing the level of integration of the European banking market (Gutierrez, 2007). In addition, the increasing mergers and acquisitions in the EMU, have positively affected the level of concentration in the banking sector, resulting to decreased levels of the *H*-index.

Fig. 2 presents the numbers of completed mergers and total deal values in some of the EMU countries (i.e., Spain, Italy, Belgium, Austria, Greece, The Netherlands) over the period 1995 to 2012. It is readily evident that most of the EMU countries experienced two main merger waves. The first wave seems to have hit the EMU countries at the end of the 1990s for reasons related to the preparations to adopt the euro. The second merger wave is evident at the mid-2000s and is more persistent in countries such as Belgium, Italy and Spain, while it has lower impact on the Mediterranean countries (i.e., Greece, Italy and Portugal).

[6] The Hausman test shows that the OLS fixed-effect model is superior for the empirical goals of our study.

Table 4
Panel cointegration tests.

Within dimension test statistics		Between dimension test statistics	
Eq. (1)—sample (EU-27)			
Panel ν-statistic	39.03*	Group ρ-statistic	−40.70*
Panel ρ-statistic	−40.78*	Group PP-statistic	−39.12*
Panel PP-statistic	−39.06*	Group ADF-statistic	−5.02*
Panel ADF-statistic	−5.21*		
Eq. (1)—sample (EMU-17)			
Panel ν-statistic	34.67*	Group ρ-statistic	−34.79*
Panel ρ-statistic	−35.13*	Group PP-statistic	−32.41*
Panel PP-statistic	−33.84*	Group ADF-statistic	−4.30*
Panel ADF-statistic	−4.78*		
Eq. (1)—sample (rest)			
Panel ν-statistic	33.17*	Group ρ-statistic	−32.16*
Panel ρ-statistic	−31.19*	Group PP-statistic	−30.27*
Panel PP-statistic	−32.80*	Group ADF-statistic	−4.62*
Panel ADF-statistic	−4.57*		
Eq. (2)—sample (EU-27)			
Panel ν-statistic	30.12*	Group ρ-statistic	−30.43*
Panel ρ-statistic	−30.00*	Group PP-statistic	−31.33*
Panel PP-statistic	−31.12*	Group ADF-statistic	−4.28*
Panel ADF-statistic	−4.10*		
Eq. (2)—sample (EMU-17)			
Panel ν-statistic	32.29*	Group ρ-statistic	−31.096*
Panel ρ-statistic	−31.18*	Group PP-statistic	−30.78*
Panel PP-statistic	−31.17*	Group ADF-statistic	−4.65*
Panel ADF-statistic	−4.45*		
Eq. (2)—sample (rest)			
Panel ν-statistic	33.74*	Group ρ-statistic	−33.66*
Panel ρ-statistic	−32.084*	Group PP-statistic	−32.57*
Panel PP-statistic	−33.74*	Group ADF-statistic	−4.93*
Panel ADF-statistic	−4.89*		
Eq. (3)—sample (EU-27)			
Panel ν-statistic	35.45*	Group ρ-statistic	−36.76*
Panel ρ-statistic	−34.84*	Group PP-statistic	−37.18*
Panel PP-statistic	−35.85*	Group ADF-statistic	−5.86*
Panel ADF-statistic	−5.71*		
Eq. (3)—sample (EMU-17)			
Panel ν-statistic	33.14*	Group ρ-statistic	−34.65*
Panel ρ-statistic	−34.23*	Group PP-statistic	−33.65*
Panel PP-statistic	−33.45*	Group ADF-statistic	−5.48*
Panel ADF-statistic	−5.32*		
Eq. (3)—sample (rest)			
Panel ν-statistic	36.14*	Group ρ-statistic	−35.54*
Panel ρ-statistic	−36.33*	Group PP-statistic	−34.54*
Panel PP-statistic	−35.35*	Group ADF-statistic	−4.84*
Panel ADF-statistic	−4.93*		

Notes: Of the seven tests, the panel ν-statistic is a one-sided test where large positive values reject the null hypothesis of no cointegration, whereas large negative values for the remaining test statistics reject the null hypothesis of no cointegration.
* Significant at 1%.

4.3.2. Unscaled models

Based on the main differences between the two approximations (scaled vs unscaled models) and the controversial results in the empirical literature described in the introduction of the paper, we re-estimate the main equations (Eqs. (1) and (3)) in an unscaled form (Eqs. (4) and (5)). In this way, we can check out the robustness of our findings.

Table 6 contains the estimation results when we do not account for the scale effect. It is evident that the coefficients are statistically significant, the signs are the expected ones and the fit is substantially high. The high value of the adjusted *R*-squared mostly evident in the PGLS method is an indication that the control variables might be correlated thus, resulting in numerically unstable estimates of the regression coefficients (multicolinearity). In order to investigate the presence of multicolinearity, we build the correlation matrix of the independent variables (Table A3 in Appendix A) and accordingly estimate the variance inflation factors (VIF)[7] generated by the two unscaled equations (Table A4 in Appendix A). From the relevant tables it is evident that the control variables are not correlated and most importantly the VIF for the two unscaled equations are negligible.

The *H*-index is less than one across all specifications, confirming the previous results and implying that monopolistic competition is the appropriate market structure in the EU banking sector. Moreover, it seems that there is no significant variation between the two specifications (Eqs. (4) and (5)). From the Wald tests, we argue that the null hypothesis cannot be accepted, pointing out that the banking sector in the EU is not characterized by a non-competitive (oligopolistic) behaviour. The two methodologies (PGLS and FMOLS) provide very little variation in similar results, revealing the robustness of our findings.

Regarding the magnitude of the relevant point elasticities and the subsequent *H*-indices, we conclude that unlike other previous studies (Bikker & Spierdijk, 2008; Bikker et al., 2012) which suggest that the scaled revenues function (including total assets as a control variable) can lead to overestimation of the degree of competition in the banking industry, our study reveals that the estimated results, generated by the unscaled price and revenue functions, are quite similar with the aforementioned ones. Similar findings can be traced in previous empirical studies (Sun, 2011).

When we restrict the sample to measure the level of competition across the EMU banks and the remaining countries, respectively, the findings do not seem to have significant variation compared against to the scaled equations. More specifically, in the EMU, the industry structure of the banking sector is characterised by monopolistic competition since the estimated *H*-index lies between zero and one across both specifications. Similarly, the level of banking competition outside the EU and the Euro-zone mechanism (i.e., Bulgaria, Czech Republic, Denmark, Hungary, Latvia, Lithuania, Poland, Romania, Sweden, and the U.K.) resembles monopolistic competitive conditions.

It is worth emphasising that when we introduce the PGLS fixed effects estimator in three out of six specifications (see, columns 3, 4 and 6), the coefficient of the WAGE input price is not statistically significant. One could argue that in estimating the *H*-index, insignificant values of input prices should not be taken into account. In the present study, this would imply that when the sample includes the EU-27 as a whole and the remaining EU countries, respectively, the degree of SMP may be underestimated (Shaffer, 2004; Delis, 2010). It is noteworthy to argue that the insignificant values of the relevant input price variable (e.g., *WAGE*) do not affect the magnitude of the *H*-statistic, confirming the presence of monopolistic competitive conditions across the EU-27.

A very interesting topic to study further was the responsiveness of the banking sector to the on-going financial crisis as well as the investigation of whether the bank competition pattern in the EU has changed due to the implementation of policies taken on a European level. The European banking sector has been hit hard by the financial crisis, which was triggered by different factors on a macro and microeconomic perspective (Polemis, 2014). From a macroeconomic perspective, prolonged low interest rates and large global imbalances, which emerged after the Asian crises, led to a bubble in both stock and in real estate markets. From a microeconomic perspective, high leverage, manager compensation and financial innovation are significant. In this scenario, considering the experiences of the different countries, it is not clear to what

[7] Variance inflation factors are used to detect multicolinearity. VIF are a scaled version of the multiple correlation coefficients between variable *j* and the rest of the independent variables. Specifically, $\text{VIF}_j = \frac{1}{1-R_j^2}$, where R_j is the multiple correlation coefficient. If R_j equals zero (i.e., no correlation between X_j and the remaining independent variables), then VIF_j equals 1. This is the minimum value. A value greater than 10 is an indication of potential multicolinearity problems (Neter, Kutner, Wasserman, & Nachtsheim, 1990).

Table 5
Empirical results for the scaled equations and the equilibrium model.

Control variables	EU-27 (Eq. (1))	EMU-17 (Eq. (1))	Rest (Eq. (1))	EU-27 (Eq. (2))	EMU-17 (Eq. (2))	Rest (Eq. (2))	EU-27 (Eq. (3))	EMU-17 (Eq. (3))	Rest (Eq. (3))
PGLS fixed effects									
Constant	0.708* (7.13)	−0.138 (−0.43)	1.371* (7.45)	1.441* (5.05)	1.239* (3.41)	0.879** (1.80)	1.252* (12.43)	1.063* (8.21)	1.574* (8.88)
FUND	0.563* (50.33)	0.161* (5.67)	0.568* (30.73)	−0.098* (−2.85)	−0.022 (−0.53)	−0.196* (−3.50)	0.362* (31.74)	0.370* (24.70)	0.352* (18.03)
WAGE	0.127* (3.80)	0.441* (3.31)	0.026 (0.68)	0.190** (2.24)	0.082 (0.71)	0.343** (2.53)	0.133* (4.05)	0.303* (5.38)	0.037 (1.02)
CAP	0.130* (3.54)	−0.193*** (−1.46)	0.312* (5.82)	−0.077 (−0.78)	−0.014 (−0.10)	−0.315** (−2.01)	0.280* (7.66)	0.032 (0.54)	0.490* (9.55)
LEV	−0.012 (−0.73)	−0.239* (−4.02)	0.097** (2.30)	0.175* (3.31)	0.182* (3.23)	0.410* (2.71)	0.050* (3.11)	0.032** (1.83)	0.118* (2.80)
RISK	0.101* (3.83)	0.125*** (1.52)	0.185* (4.04)	0.036 (0.55)	0.107 (1.41)	−0.382* (−2.63)	0.113* (4.35)	0.155* (4.83)	0.146* (3.07)
SIZE	−0.042* (−8.35)	−0.099* (−7.74)	−0.044* (−3.38)	−0.006 (−0.44)	0.007 (0.43)	0.022 (0.68)	−0.046* (−6.72)	−0.040* (−6.94)	−0.021*** (−1.64)
H-index++	**0.82**	**0.41**	**0.90**	**0.03**	**0.05**	**0.17**	**0.77**	**0.70**	**0.88**
Observations	403	253	148	386	244	142	403	255	148
Adjusted R^2	0.94	0.60	0.96	0.42	0.33	0.45	0.87	0.92	0.95
F-statistic	216.39* [0.00]	18.08* [0.00]	228.85* [0.00]	9.64* [0.00]	6.54* [0.00]	8.81* [0.00]	231.82* [0.00]	127.02* [0.00]	208.38* [0.00]
Wald test $H=0$	F=1445.22* [0.00]	F=42.91* [0.00]	F=389.79* [0.00]	F=0.09 [0.76]	F=0.61 [0.43]	F=1.82 [0.18]	F=1411.25* [0.00]	F=756.22* [0.00]	F=394.12* [0.00]
Wald test $H=1$	F=69.02* [0.00]	F=90.50* [0.00]	F=4.21** [0.04]	–	–	–	F=118.44* [0.00]	F=133.07* [0.00]	F=7.40* [0.00]
FMOLS									
Constant	0.687* (5.53)	−0.139* (6.65)	1.125* (6.32)	1.759* (7.58)	1.034* (4.58)	0.751* (2.95)	1.599* (14.9)	0.854* (9.44)	1.354* (10.65)
FUND	0.504* (54.9)	0.154* (10.23)	0.461* (15.3)	−0.150* (−5.73)	−0.031* (−3.14)	−0.242* (−5.11)	0.320* (33.4)	0.361* (18.31)	0.325* (15.71)
WAGE	0.163* (14.5)	0.319* (6.18)	0.124* (4.48)	0.185* (−9.69)	0.071* (6.31)	0.319* (3.90)	0.180* (6.4)	0.286* (7.31)	0.062* (5.34)
CAP	0.240* (16.6)	−0.179* (−3.94)	0.254* (7.31)	−0.008* (−7.58)	−0.021* (−3.04)	−0.295* (−4.10)	0.300* (8.29)	0.064* (4.10)	0.437* (10.39)
LEV	0.081* (12.3)	−0.223* (−7.31)	0.121* (5.36)	0.650* (9.10)	0.165* (4.61)	0.376* (3.19)	0.042* (5.52)	0.054* (3.94)	0.165* (4.91)
RISK	0.106* (9.00)	0.108 (1.14)	0.172* (5.46)	0.040* (5.69)	0.135* (4.14)	−0.362* (−4.51)	0.160* (8.67)	0.174* (6.14)	0.183* (4.41)
SIZE	−0.040* (13.5)	−0.115* (−8.31)	−0.051* (−3.64)	−0.013 (0.38)	0.011* (5.31)	0.045** (1.91)	−0.030*** (−1.43)	−0.071* (−8.32)	−0.012* (−3.94)
H-index++	**0.91**	**0.29**	**0.84**	**0.03**	**0.02**	**0.22**	**0.80**	**0.71**	**0.82**
Observations	403	253	148	386	244	142	403	255	148
Adjusted R^2	0.89	0.54	0.78	0.39	0.36	0.50	0.92	0.83	0.82
Wald test $H=0$	1324.31* [0.00]	F=91.82* [0.00]	F=314.25* [0.00]	F=0.14 [0.71]	F=0.95 [0.39]	F=2.11 [0.14]	F=1534.70* [0.00]	F=831.19* [0.00]	F=439.54* [0.00]
Wald test $H=1$	F=73.91* [0.00]	F=97.32* [0.00]	F=0.11 [0.61]	–	–	–	F=231.32* [0.00]	F=201.14* [0.00]	F=1.93* [0.15]

Notes: Figures in parentheses denote t-ratios. Figures in square brackets denote *p*-values.
++ The sum of the three elasticities in Eq. (2) denote the equilibrium *E*-statistic and not the *H*-index.
Significant at *1%, **5% and ***10%, respectively.

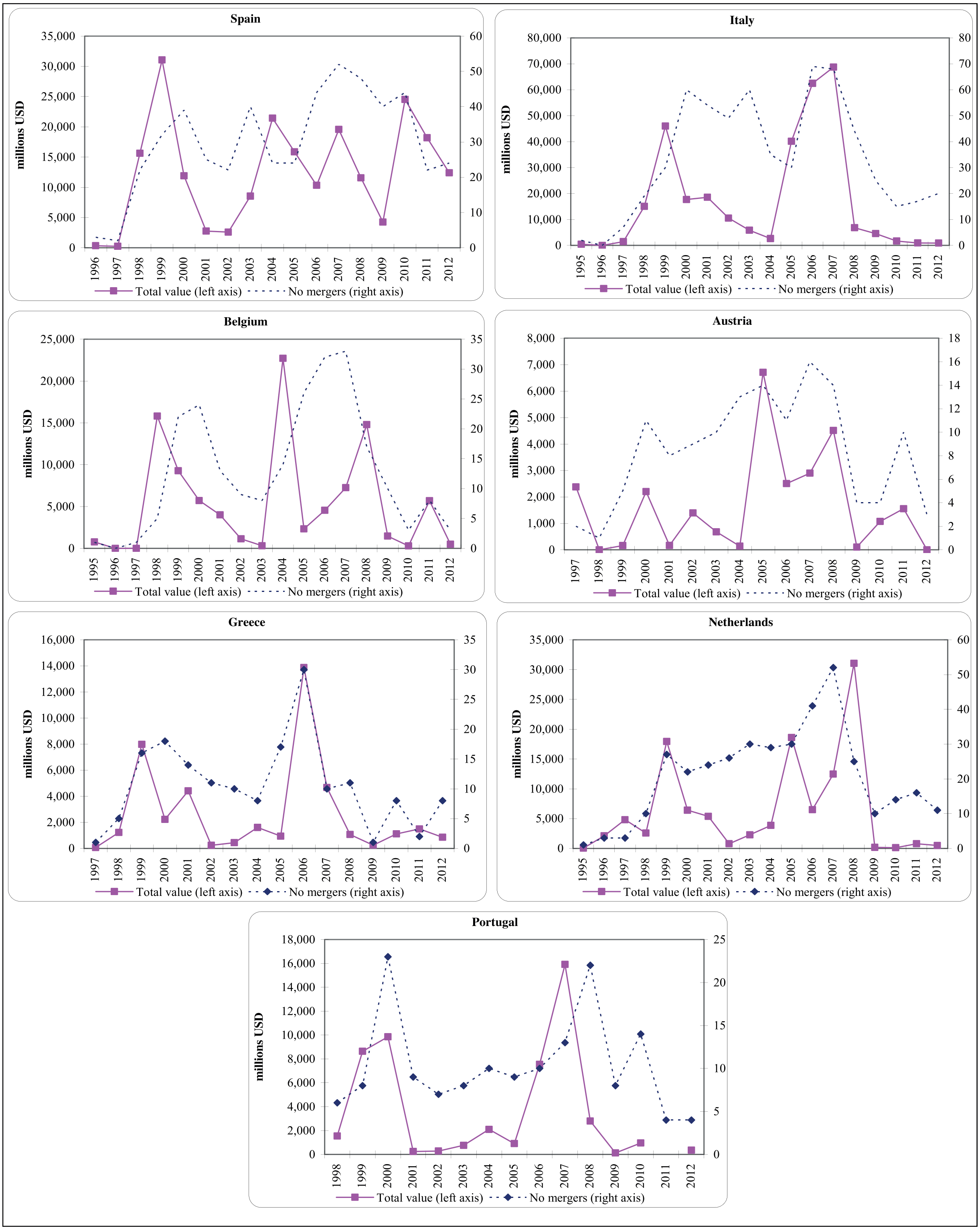

Fig. 2. Number and total deal values of mergers in the banking sector for some of the EMU countries. *Source*: Bloomberg.

extend competition itself contributed to the crisis event. Many banks incurred heavy losses and only a few managed to stay in business, with exceptional support from national governments and central banks. This process has started in several EMU countries (Greece, Portugal, Spain, Italy) and is being supported by EU mechanisms that require the restructuring of corporations, including banks, if certain state aid has been provided (Sun, 2011).

In order to assess the first impact of the on-going financial crisis, we divide the whole sample into three sub-periods. The first period covers the years from 1996 to 2000 (pre-EMU period), while the second period runs from 2001 to 2007 (post-EMU period). Finally, the third sub-sample covers the period 2008–2011 (post-crisis period). Table 7 displays the estimated average *H*-statistic across all three sub-periods. It is evident that the overall competition level

Table 6
Empirical results for the unscaled price and revenue equations.

Control variables	EU-27 (Eq. (4))	EMU-17 (Eq. (4))	Rest (Eq. (4))	EU-27 (Eq. (5))	EMU-17 (Eq. (5))	Rest (Eq. (5))
PGLS fixed effects						
Constant	2.113* (20.95)	2.542* (15.01)	1.374* (7.47)	2.413* (22.41)	2.732* (15.93)	1.571* (8.88)
FUND	0.575* (48.45)	0.543* (33.25)	0.567* (30.59)	0.365* (29.22)	0.380* (20.21)	0.352* (18.14)
WAGE	0.066* (2.10)	0.225* (2.97)	0.026 (0.70)	0.007 (0.27)	0.147* (2.36)	0.036 (1.00)
CAP	0.284** (7.45)	0.119*** (1.58)	0.310* (5.81)	0.432* (12.05)	0.180* (2.50)	0.488* (9.56)
LEV	−0.063* (−3.78)	−0.052** (−2.46)	0.097* (2.29)	0.020 (1.08)	−0.001 (−0.05)	0.119* (2.84)
RISK	0.147* (5.60)	0.172* (4.84)	0.185** (4.01)	0.106* (3.62)	0.059 (1.34)	0.143* (3.00)
ASSET	0.967* (171.80)	0.974* (134.09)	0.956* (73.99)	0.970* (164.51)	0.969* (143.69)	0.978* (77.23)
H-index	**0.92**	**0.89**	**0.90**	**0.80**	**0.71**	**0.87**
Observations	403	254	148	403	254	148
Adjusted R^2	0.99	0.99	0.99	0.99	0.99	0.99
F-statistic	5752.52* [0.00]	4098.37* [0.00]	2854.81* [0.00]	4560.48* [0.00]	3573.83* [0.00]	2940.06* [0.00]
Wald test $H=0$	$F=1264.17$* [0.00]	$F=531.61$* [0.00]	$F=388.78$* [0.00]	$F=932.93$* [0.00]	$F=425.79$* [0.00]	$F=394.77$* [0.00]
Wald test $H=1$	$F=8.41$* [0.00]	$F=8.69$* [0.00]	$F=4.43$** [0.04]	$F=55.05$* [0.00]	$F=72.92$* [0.00]	$F=7.95$* [0.00]
FMOLS						
Constant	1.897* (15.63)	1.907* (8.95)	1.328* (7.64)	2.017* (10.93)	1.719* (6.73)	1.628* (6.72)
FUND	0.518* (21.38)	0.549* (9.13)	0.487* (6.52)	0.416* (5.64)	0.483* (8.92)	0.442* (5.63)
WAGE	0.104* (4.57)	0.137* (6.94)	0.093* (5.43)	0.072* (5.26)	0.094* (5.84)	0.083* (6.41)
CAP	0.236* (4.16)	0.268* (7.11)	0.204* (4.61)	0.249* (5.66)	0.281* (6.58)	0.241* (5.38)
LEV	−0.077* (5.27)	−0.083* (6.48)	−0.057* (5.28)	−0.052* (4.26)	−0.060* (6.26)	−0.072* (5.49)
RISK	0.129* (4.36)	0.146* (5.68)	0.114* (5.13)	0.128* (5.95)	0.172* (5.46)	0.155* (5.48)
ASSET	0.638* (4.59)	0.581* (7.93)	0.529* (4.73)	0.502* (6.28)	0.541* (4.85)	0.498* (5.62)
H-index	**0.86**	**0.95**	**0.78**	**0.74**	**0.86**	**0.77**
Observations	403	254	148	403	254	148
Adjusted R^2	0.63	0.69	0.62	0.57	0.60	0.55
Wald test $H=0$	896.37* [0.00]	917.65* [0.00]	841.99* [0.00]	963.28* [0.00]	966.71* [0.00]	885.74* [0.00]
Wald test $H=1$	29.85* [0.00]	10.91*** [0.10]	47.88* [0.00]	50.92* [0.00]	47.82* [0.00]	45.28* [0.00]

Notes: Figures in parentheses denote *t*-ratios. Figures in square brackets denote *p*-values.
Significant at *1%, **5% and ***10%, respectively.

in the EMU countries dropped slightly after the formation of the EMU, from 0.77 to 0.68 (PGLS) and from 0.83 to 0.80 (FMOLS), respectively. These findings are in alignment with those in prior similar studies (Bikker & Spierdijk, 2008; Sun, 2011). While the small decline in the level of bank competition in the euro area is statistically significant, it is though somewhat larger than the estimates provided by Sun (2011) who reports a decline from 0.699 to 0.518. This could be explained by the fact that in contrast to our research approach, the previous study applied a scaled revenue function to measure the *H*-index for the two sub-periods

Table 7
Empirical results for the EMU countries (unscaled equations).

Control variables	[1996–2000] (Eq. (4))	[2001–2007] (Eq. (4))	[2008–2011] (Eq. (4))	[1996–2000] (Eq. (5))	[2001–2007] (Eq. (5))	[2008–2011] (Eq. (5))
PGLS fixed effects						
Constant	1.243* (4.04)	1.881* (7.93)	3.012** (2.27)	1.977* (6.71)	1.869* (8.05)	4.937* (7.19)
FUND	0.651* (22.54)	0.483* (21.96)	0.617* (20.72)	0.456* (17.07)	0.286* (11.99)	0.412* (16.40)
WAGE	0.067 (0.58)	0.013 (0.12)	0.200*** (1.73)	0.081 (0.70)	−0.124 (−1.20)	0.297* (17.59)
CAP	0.048 (0.38)	0.183*** (1.86)	−0.237*** (−1.73)	0.144 (1.21)	0.365* (3.57)	−0.197* (−2.58)
LEV	−0.168** (−2.91)	0.010 (0.26)	0.090 (0.82)	−0.106 (−1.38)	−0.007* (−0.22)	0.230* (2.74)
RISK	0.216* (3.65)	0.269* (5.20)	0.570** (2.17)	0.182* (2.99)	0.226* (3.86)	−0.063 (−0.43)
ASSET	0.981* (67.13)	0.969* (110.57)	0.915* (11.37)	0.983* (67.03)	0.976* (108.34)	1.018* (44.67)
H-index	**0.77**	**0.68**	**0.58**	**0.68**	**0.53**	**0.51**
Observations	84	117	53	84	117	53
Adjusted R^2	0.99	0.99	0.99	0.99	0.99	0.99
F-statistic	6814.90* [0.00]	3186.64* [0.00]	629.73* [0.00]	6250.77* [0.00]	3652.35* [0.00]	16317.73* [0.00]
Wald test $H=0$	$F=197.69$* [0.00]	$F=201.24$* [0.00]	$F=10.99$* [0.00]	$F=154.50$* [0.00]	$F=119.80$* [0.00]	$F=152.08$* [0.00]
Wald test $H=1$	$F=18.61$* [0.00]	$F=45.088$* [0.00]	$F=5.79$** [0.02]	$F=33.76$* [0.00]	$F=96.43$* [0.00]	$F=4.44$** [0.04]
FMOLS						
Constant	1.046* (5.41)	1.263* (4.86)	1.186* (4.58)	1.098* (4.82)	1.119* (4.39)	1.233* (5.48)
FUND	0.573* (7.15)	0.542* (6.19)	0.581* (6.73)	0.519* (5.84)	0.494* (5.84)	0.562* (6.93)
WAGE	0.081* (4.93)	0.103* (5.41)	0.137* (5.48)	0.126* (4.58)	0.120* (5.47)	0.147* (5.37)
CAP	0.178* (6.53)	0.155* (5.64)	0.148* (5.38)	0.133* (5.03)	0.129* (5.38)	0.140* (6.72)
LEV	−0.137* (5.83)	−0.141* (5.29)	−0.136* (4.92)	−0.108* (4.39)	−0.084* (4.55)	−0.114* (5.48)
RISK	0.186* (5.95)	0.162* (6.07)	0.149* (4.83)	0.116* (5.48)	0.094* (5.28)	0.116* (6.03)
ASSET	0.682* (7.18)	0.593* (5.83)	0.522* (5.38)	0.493* (6.52)	0.471* (5.62)	0.526* (6.55)
H-index	**0.83**	**0.80**	**0.87**	**0.78**	**0.74**	**0.85**
Observations	84	117	53	84	117	53
Adjusted R^2	0.63	0.60	0.56	0.58	0.53	0.55
Wald test $H=0$	84.39* [0.00]	65.49* [0.00]	53.82* [0.00]	42.39* [0.00]	52.39* [0.00]	62.39* [0.00]
Wald test $H=1$	56.31* [0.00]	52.38* [0.00]	49.06* [0.00]	58.96* [0.00]	64.38* [0.00]	74.48* [0.00]

Notes: Figures in parentheses denote *t*-ratios. Figures in square brackets denote *p*-values.
Significant at *1%, **5% and ***10%, respectively.

(i.e., pre- and post-EMU). However, the magnitude of the *H*-index is smaller when we use total revenues (TR) as the dependent variable (0.53 and 0.74, respectively).

The on-going financial crisis and the relevant policies seem to have affected the level of bank competition in the EMU countries, as indicated by the *H*-index before and after the crisis. We note that the level of bank competition during the first years of the on-going financial crisis shows a slight decline, since the *H*-index is estimated to 0.58 (column 3) and 0.51 (column 6), respectively. The outcome seems to be overturned when we use the FMOLS methodology. However, in both cases, we must bear in mind that these post-crisis estimates provide preliminary evidence only in view of the fact that the structural changes in the aftermath of the crisis may distort the long-run market equilibrium required for the validity of the *H*-index (Sun, 2011).

Regarding the statistically significant impact of the three input prices (i.e., FUND, WAGE and CAP) on revenues (i.e., gross and total interest revenues), we infer that when our econometric model is estimated with PGLS fixed effects, in four out of six specifications (see, columns 1, 2 4 and 5), the coefficient of (at least) one input price is not statistically significant. This implies that when the sample covers the pre-EMU (1996–2000) and the post-EMU period (2001 to 2007), the degree of SMP may be understated. However, this is not the case since the insignificant values of the relevant input price variables do not alter the final outcome, supporting the presence of a monopolistically competitive environment in the EMU.

Based on the fact that our analysis covers the period of the first years of the on-going financial crisis (2008–2011) with a large amount of entries and exits, it seems significant to test the equilibrium conditions for this period. From the ROA test applied to this sub-sample, we cannot reject the null hypothesis ($E = 0$), since the *F*-statistic of the PGLS methodology equals 0.27 with a corresponding *p*-value equal to 0.61[8]. This finding is in accordance with the adoption of $0 < H < 1$ and does not provide any clear evidence that the EMU banking sector is in long-run competitive equilibrium, notwithstanding the on-going financial crisis. In addition, the general decline in the competition level after the adoption of the euro zone mechanism and the first impact of the on-going financial crisis can be attributed to the process of consolidation as well as to the movement of bank activities from traditional financial business to off-balance sheet activities (Sun, 2011).

5. Conclusions and policy recommendations

The new evidence suggests that the lack of competition can undermine the stability of the banking sector, especially in cases where some banks become too big to fail. The debate on the matter is yet inconclusive; however, it is often correlated with the way competition in the banking sector is measured. The usual concentration ratios (market shares and the Hirschman–Herfindahl Index) do not always seem to be appropriate measures of competition due to some special characteristics of the banking sector, such as, information asymmetries in corporate borrowing, switching costs in retail banking, and network externalities in payment systems, which all increase the difficulties to apply the standard S–C–P paradigm to this sector.

In order to clarify the above issues we empirically estimated the level of banking competition in the EU-27 and across two sub-samples, i.e., Eurozone and the remaining European countries, by employing two different econometric techniques (PGLS and FMOLS) on two distinct empirical models (scaled and unscaled price and revenue equations). The empirical results were robust and, despite the presence of differences in the size of the sample under investigation, they were consistent with other previous empirical studies, providing evidence in favour of monopolistic competition. Our analysis did not reveal any significant differences in the magnitude of the relevant estimations between the scaled and the unscaled modelling approaches.

The EU banking sector is characterized by the presence of monopolistic competition. The relatively high level of the *H*-index in the EU-27 vis-à-vis the EMU-17 can be potentially explained by a number of policy initiatives taken as well as by certain regulatory restrictions, which were removed over the last decade, towards the creation of an integrated market for banking services. Further, the low level of the *H*-index regarding the EMU-17 countries can be explained by the increasing level of mergers and acquisitions activity occurred.

These findings reveal that the European banking sector is still extremely fragmented. However, it is characterised by a range of entry barriers that need further exploration (EC, 2007). Specifically, some of these barriers are generated by the presence of economies of scale or scope, consumption externalities and standardisation requirements with respect to networks, such as payment systems. Others are of an artificial nature resulting from specific regulation or conduct of firms (i.e., access to networks, discriminatory fee structures).

The results did not dramatically change when we split the sample into two distinctive regions (EMU-17 and the remaining countries). While the market structure across all countries was still monopolistic competition, the intensiveness of the competition level was differentiated. The euro zone countries experienced a slight, albeit significant, decline in bank competition after the formation of the EMU and the on-going financial crisis. The state aid provided to EU banks during the period of financial instability seems to have left the market structure unaltered (especially in the EMU countries) given that the *H*-index displayed a modest decrease, albeit remaining well less than one.

Following the results of other similar studies, this paper suggests that moving beyond the mere monitoring of banks' concentration rates and by using non-structural measures for assessing competition levels, such as the *H*-index, we can deliver a more informative insight on competition in the European banking sector. From a policy perspective, the results of this study support the need for a surge in the attention of policy makers and government officials towards a deeper reform of competition policies in the European banking sector. In this way, it is hoped that policy makers maybe able to introduce more proper competition policies, such as those that allow easier entry in the relevant markets, strengthen integration, and diminish the collusion practices of the existing competitors' in the relevant markets, thus, contributing to the stability of the banking system. In this respect, competition policy may act as a catalyst in order to support wider financial and business stability, while creating the appropriate conditions for more consolidated financial markets in both the short- and the long-run. In addition, such efforts can facilitate a better control over further turbulences introduced by bailout efforts. In this unstable competitive banking environment, competition authorities have a major role to play, at least for introducing a dialogue with the regulators to preserve a balance between competition and ex ante regulation. The latter, consists inter alia of limitations on business activities, managerial high competence's requirements, close monitoring of the flows across banking institutions and their shareholders, risk based capital, and adequate capital for the development of very large banking institutions. It is worth emphasizing that with adequate capital requirements and prudential regulation would make it possible to amend for the negative impact of competition on banking sector stability.

As our findings indicate, the presence of a monopolistic competitive environment in the banking sector across the EU-27 raises important managerial implications. First, the absence of entry barriers in tandem with a low level of SMP in the industry

[8] The same outcome applies when we use the FMOLS method.

might affect managerial decisions towards their engagement into cross-border transactions (i.e., mergers and acquisitions). Second, since monopolistic competition requires differentiation of product offerings, bank managers may expand their sources of earnings through diversification of assets and liabilities, and by reducing the operational costs and/or increasing non-interest revenues. Finally, our results provide a clearer and more updated insight to EU policy makers towards a more proper control regarding competition in the banking industry.

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Appendix A

See Tables A1–A4.

Table A1
Descriptive statistics for the dependent variables*.

Variables	*P*			*ROA*			*Z*			*GIR*			*TR*		
Statistical measures	*EU-27*	*EMU-17*	*Rest-10*	*EU-27*	*EMU-17*	*Rest-10*	*EU-27*	*EMU-17*	*Rest-10*	*EU-27*	*EMU-17*	*Rest-10*	*EU-27*	*EMU-17*	*Rest-10*
Observations	412	258	153	393	246	147	412	259	153	412	258	153	412	259	153
Cross sections	27	17	10	27	17	10	27	17	10	27	17	10	27	17	10
Mean	−2.98	−3.04	−2.84	0.58	0.49	0.73	−2.45	−2.58	−2.24	15.24	15.28	14.78	15.76	15.99	15.38
Median	−2.98	−3.00	−2.85	0.58	0.46	0.79	−2.49	−2.57	−2.23	15.10	14.96	14.95	15.60	15.74	15.39
Maximum	−0.94	−1.57	−0.94	2.43	2.23	2.43	0.09	−0.49	0.09	19.82	19.82	19.74	20.24	20.14	20.24
Minimum	−6.64	−6.64	−4.38	−1.35	−1.35	−0.87	−4.41	−4.41	−3.37	5.75	5.75	10.18	8.56	8.56	10.62
Standard deviation	0.51	0.47	0.56	0.44	0.41	0.45	0.49	0.41	0.55	2.23	2.41	1.94	2.10	2.20	1.87
Coefficient of variation	−0.17	−0.15	−0.20	0.76	0.84	0.62	−0.20	−0.16	−0.25	0.15	0.16	0.13	0.13	0.14	0.12
Skewness	−1.03	−2.65	0.30	−0.42	−0.62	−0.48	0.68	0.30	0.54	−0.41	−0.47	−0.01	−0.16	−0.34	0.07
Kurtosis	12.37	20.74	4.03	5.93	7.25	5.30	6.54	9.50	4.66	3.83	3.77	3.18	2.98	2.95	3.30

* Variables are in natural logarithms.

Table A2
Descriptive statistics for the control variables*(continued).

	FUND			*WAGE*			*CAP*			*LEV*			*RISK*			SIZE		
Statistical measures	*EU-27*	*EMU-17*	*Rest-10*	*EU-27*	*EMU-17*	*Rest-10*	*EU-27*	*EMU-17*	*Rest-10*	*EU-27*	*EMU-17*	*Rest-10*	*EU-27*	*EMU-17*	*Rest-10*	*EU-27*	*EMU-17*	*Rest-10*
Observations	409	256	153	407	259	148	412	259	153	411	258	153	410	257	153	412	259	153
Cross sections	27	17	10	27	17	10	27	17	10	27	17	10	27	17	10	27	17	10
Mean	−3.23	−3.22	−3.23	−4.69	−4.80	−4.50	−3.91	−4.12	−3.57	−2.70	−2.80	−2.53	−0.74	−0.76	−0.70	16.92	16.51	17.62
Median	−3.18	−3.13	−3.25	−4.68	−4.73	−4.56	−3.97	−4.12	−3.50	−2.71	−2.88	−2.46	−0.69	−0.69	−0.66	17.66	17.72	17.53
Maximum	−0.90	−1.97	−0.90	−1.04	−1.04	−3.14	−0.49	−0.54	−0.49	−0.16	−0.16	−1.53	−0.16	−0.21	−0.16	23.38	22.60	23.38
Minimum	−11.38	−11.38	−4.41	−8.77	−8.77	−5.65	−6.69	−6.69	−4.75	−3.75	−3.67	−3.75	−1.86	−1.86	−1.81	2.64	2.64	12.67
Standard deviation	0.71	0.75	0.63	0.62	0.65	0.50	0.68	0.63	0.62	0.50	0.50	0.45	0.31	0.32	0.29	4.56	5.45	2.19
Coefficient of variation	−0.22	−0.23	−0.20	−0.13	−0.14	−0.11	−0.17	−0.15	−0.17	−0.19	−0.18	−0.18	−0.42	−0.42	−0.41	0.27	0.33	0.12
Skewness	−4.49	−6.31	0.74	−0.13	−0.04	0.23	0.54	0.72	0.54	0.87	1.55	−0.21	−1.00	−1.02	−0.92	−1.98	−1.66	0.27
Kurtosis	50.81	63.45	3.98	12.93	15.50	2.63	7.10	10.74	5.79	5.39	8.17	2.55	4.10	4.05	3.96	6.89	4.74	3.28

* Variables are in natural logarithms. Due to space limitations, we provide only the results for the SIZE and not for the ASSET variable.

Table A3
Correlation matrix between the control variables (EU-27).

Variables	*FUND*	*WAGE*	*CAP*	*LEV*	*RISK*	*SIZE*
FUND	**1.00**					
WAGE	**0.02** (0.32) [0.75]	**1.00**				
CAP	**0.01** (0.22) [0.83]	**0.91*** (45.46) [0.00]	**1.00**			
LEV	**−0.20*** (−4.09) [0.00]	**0.51*** (11.99) [0.00]	**0.58*** (14.13) [0.00]	**1.00**		
RISK	**−0.08** (−1.55) [0.12]	**0.22*** (4.44) [0.00]	**0.17*** (3.43) [0.00]	**0.17*** (3.51) [0.00]	1.00	
SIZE	**−0.01** (−0.27) [0.78]	**−0.14*** (−2.97) [0.00]	**−0.14** (−2.85) [0.00]	**−0.12**** (−2.47) [0.02]	**0.10**** (2.06) [0.04]	**1.00**

Notes: Figures in parentheses denote *t*-ratios. Figures in square brackets denote *p*-values.
Significant at *1%, **5% and ***10%, respectively. Due to space limitations, we provide only the results for the SIZE and not for the ASSET variable.

Table A4
Variance inflation factors for the EU-27 unscaled equations (VIF).

Variables	Coefficient variance	Centered VIF
Price equation: dependent variable GIR		
Constant	0.010167	–
FUND	0.000141	1.089934
WAGE	0.000985	2.387438
CAP	0.001455	2.458326
LEV	0.000281	1.589980
RISK	0.000687	1.237794
ASSET	3.17E – 05	2.274000
Revenue equation: sependent variable *TR*		
Constant	0.011597	–
FUND	0.000156	1.074266
WAGE	0.000790	2.152403
CAP	0.001288	2.407038
LEV	0.000336	1.338325
RISK	0.000863	1.370595
ASSET	3.47E – 05	2.108906

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