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Exploring short- and long-run links from bank competition to risk

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Abstract

The current literature offers diverse findings on the bank competition-risk relationship. We seek to advance understanding by looking at both short- and long-run relationships for banks from 27 EU countries, using a six-year period before and since 2007 and employing both the H-statistic and the Lerner index as measures of competition. We thus highlight further nuances in the competition–risk relationship that are absent from the current literature. Both measures have a positive short-run relationship with risk, while long-run effects differ. Underlying this, the competition measures differ in their relationship to the volatility of profits, with important policy implications.

KEYWORDS

bank competition, EU banking markets, financial stability, Lerner index, Panzar–Rosse H-statistic, Z-score

JEL CLASSIFICATION

*G*21, *G*28

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1 | INTRODUCTION

The subject of bank competition and risk has returned to the fore with the global financial crisis (GFC) in 2008–2009, with a popular view being that competition between financial institutions during the preceding boom was an important feature underlying the crisis (e.g., the majority view of the Financial Crisis Inquiry Commission, 2011). This, in turn, would imply that the benefits of banking competition for economic growth and efficiency need to be considered in the balance. On the other hand, an extensive literature, generally estimated using pre-crisis data, finds diverse results for the relationship between competition and risk. This follows, on the one hand, the so-called franchise value or competition–fragility approach – that more competition reduces the value of a banking licence and thus induces firms to take more risk – and, on the other hand, the competition–stability view, that, with low levels of competition, banks may charge excessively high rates of interest on loans and hence generate adverse selection and moral hazard on their loan books. Both types of results have been found in the empirical literature. An emerging set of studies also suggests that both high and low levels of competition may be adverse for risk, that is, there is a U-shaped relationship, and that structural and regulatory features could affect country-level trade-offs.

Further ambiguity was introduced by the common use of the concentration of a banking system as a proxy for competition in earlier work in this area (e.g., Beck et al., 2006). However, the theory of contestable markets suggests a concentrated system may be highly competitive if there is sufficient potential competition from outside, as may be permitted by regulations allowing new entry, as in the European Single Market. The empirical literature mostly covers the period up to 2007, thus leaving open the interpretation of the post-crisis world, in which there are a diminishing number of banks, extensive government intervention, and – many would argue – less competition. The literature also suggests untested assumptions behind proposals for enforced structural change in banking, such as the Vickers proposal in the United Kingdom.

In this paper, we relate indices of banking market competition for the United Kingdom and other European Union (EU) countries to banking risk. Our aim is to test for dynamic as well as long-run links from competition to risk, before as well as after the GFC, thus highlighting further nuances in the relation of competition to stability that have not been emphasized in the literature to date. First, we use the Panzar–Rosse H-statistic to assess the changing nature of competition in individual markets over time. Among earlier studies using this approach is an analysis of the competition in the major Economic and Monetary Union (EMU) countries compared to the United States just prior to the EMU (De Bandt and Davis, 2000). Second, we investigate alternative approaches to measuring bank competition by using Iwata's Lerner index (Bikker, 2004).

This paper is structured as follows. In section 2, we briefly examine the theoretical literature and summarize recent empirical work on competition and risk in banking, highlighting the fact that the datasets used largely cover the boom period or before, with relatively little work on the post-GFC period. We note that empirical as well as theoretical work offers diverse findings and we probe the reasons why. We question, for example, whether global datasets are fully informative for policy in advanced countries. In section 3, we outline the data and methodology we use in the exercise, before discussing our empirical results. We test the impact of each measure of competition on risk, differentiating between static and dynamic aspects and between the results from the H-statistic and Lerner index and showing several robustness checks, before considering reasons for differences. The final section draws our conclusions.

2 | LITERATURE SURVEY

As mentioned, there are two broad approaches to the relationship between banking competition and risk. The theory of franchise value (Keeley, 1990), or competition–fragility, suggests that institutions in an uncompetitive banking system have incentives to avoid risk, because a banking licence is valuable in such a context, with restricted entry and probably large capital cushions (technically, the franchise value is high due to monopoly rents). This typifies the highly regulated situation of banking systems from World War II to the late 20th century, when banks had a great deal of market power and there was little financial instability. Then, when deregulation arises, the value of the licence declines as excess returns are competed away by new entrants (from abroad as well, where permitted) and by more intense competition between existing players. This situation gives incentives to increase balance sheet risk to recover the previous level of profitability, since banks effectively shift risks to depositors (or deposit insurers) and thus become more vulnerable to shocks. In a context of limited liability, there is also asymmetric risk for owners and managers, which can increase the positive effect of competition on risk. This effect could be intensified by an incentive to underinvest in screening and monitoring, since information rents from lending relationships are less valuable, as customers can switch banks more readily (Allen and Gale, 2000, 2004). Meanwhile, larger banks in a less competitive system may be better able to diversify risks and are easier to supervise (Allen and Gale, 2000; World Bank, 2013).

The alternative approach, which is due to Boyd and De Nicolo (2005), is often called the competition–stability approach. Whereas lower lending rates in competitive banking markets increase borrower scope for repayment, higher lending rates in uncompetitive markets lead to adverse selection, with only riskier borrowers seeking funds and moral hazard inducing borrowing firms to take greater risks (e.g., Stiglitz and Weiss, 1981). With perfect correlation of loan defaults, this naturally affects the entire portfolio. Large banks in uncompetitive markets can also be harder to supervise (Beck et al., 2006) and are vulnerable both to contagion and to 'too big to fail' incentives for risk taking, which can enhance the competition–stability effect (Mishkin, 1998).

An extension of Boyd and De Nicolo's (2005) approach allowing for imperfect correlation in loan defaults can instead generate a U-shaped relationship between risk and competition (Martinez-Miera and Repullo, 2010), since the initial benefit to lower probabilities of default from lower loan rates (risk-shifting effect) begins to be offset by lower revenues (margin effect), leading to instability. Wagner (2010) shows that, if banks can adjust their loan portfolios, the link from competition to risk taking could be reversed, since, for example, when borrowers become safer, banks shift their portfolios to higher-risk borrowers *per se*. There could be overcompensation due to loss of franchise value from competition. Hakenes and Schnabel (2011) show that the effect of capital adequacy on risk taking depends on whether the market is subject to competition–stability or competition–fragility, since capital requirements reduce competition, raise interest rates, and can lead banks to choose more correlated loan portfolios. Berger et al. (2009) argue that, even if loss of market power induces riskier loan portfolios, charter values may not fall if banks protect themselves with higher equity, lower risk securities, or the use of credit derivatives. This argument could help reconcile the two hypotheses.

The empirical literature is summarized in Table 1. For a more detailed discussion, see Davis and Karim (2013). Note at this point that the outstanding result from existing work is there are differing outcomes in terms of the competition–fragility and competition–stability paradigms. Some papers consistently favour the former, others the latter, while still others suggest that the outcome may differ between countries with structural and regulatory factors or even with the level of competition itself.

Unlike the studies cited above, we are able to cover comparable periods before and after the GFC, using Bankscope data up to 2012 and the 2011 wave of the World Bank Regulation and Supervision Database (Barth et al., 2012; Čihák et al., 2012b).

TABLE 1 Summary of recent empirical work on competition and risk in banking

This table reports a summary of recent empirical papers on bank competition, showing the authors and date of the study, the dependent variable, the competition variable(s), the dataset, the date of the sample, and the key result.

Study (date)	Dependent	Competition variable	Dataset	Date of sample	Key result
Anginer et al. (2012)	Distance to default	Lerner index	Global, bank by bank (publicly quoted institutions)	1996–2009	Competition enhances stability
Beck et al. (2006)	Banking crises	Concentration, competition-related regulations	Global, macroeconomic	1980–1997	More concentrated systems have fewer crises
Beck et al. (2013)	Z-score, St.Dev. Return on Assets, Equity/Total Assets	Lerner index	Global, bank by bank (1,600 US banks only)	1994–2009	The competition-risk trade-off varies with structural and regulatory conditions
Berger et al. (2009)	Z-score, Non-Performing Loans, Equity/Total Assets	Lerner index and concentration (loans and deposits)	Advanced countries (91% US banks)	1999–2005	Banks with more market power take less risk overall but more loan risk
Čihák and Schaeck (2010)	Banking crises	Financial soundness indicators	Global, macroeconomic	1994–2007	Some financial soundness indicators can help identify incipient systemic risks
Craig and Dinger (2009)	Z-score	Deposit rates	US banks	1997–2006	Deposit market competition raises risk
Demirgüç-Kunt and Huizinga (2010)	Z-score	Non-deposit funding share and fee income share	Global, bank by bank	1995–2007	Non-traditional funding and lending enhance risk
Fu et al. (2014)	Z-score, probability of bankruptcy	Lerner index, concentration	Asia-Pacific, bank by bank	2003–2010	Lower pricing power enhances risk exposure while concentration reduces risk

(Continues)

TABLE 1 (Continued)

Study (date)	Dependent	Competition variable	Dataset	Date of sample	Key result
Gropp et al. (2010)	Supervisory ratios	Concentration (control) and guarantees	OECD country banks	2003	Government guarantees raise the risk taking of competitor banks
Jimenez et al. (2010)	Non-performing loans	Lerner index	Spanish banks (by province)	1988–2003	Link from interest rate competition to loan risk
Liu et al. (2012)	Z-score and other measures (provisions/ loans. reserves/loans, ROA volatility)	H-statistic and concentration	Four Southeast Asian countries	1998–2008	H-statistic is either inversely related or unrelated to risk measures
Schaeck and Čihák (2012)	Capital ratios	H-statistic and banks/ population	European, bank by bank	1999–2005	Competition boosts capital ratios
Schaeck et al. (2009)	Banking crises	H-statistic	Global, bank by bank and macroeconomic	1980–2005	Competitive systems less prone to crisis
Tabak et al. (2012)	Z-score	Boone measure	Latin America, bank by bank	2003–2008	High and low competition enhances stability, medium competition reduces it
Uhde and Heimeshoff (2009)	Z-score	Concentration	EU countries, macroeconomic	1997–2005	Concentration raises risk
Weiss et al. (2013)	Marginal expected shortfall, lower tail dependence	Concentration	Global bank mergers	1991–2009	Merged banks have a greater impact on systemic risk
Yeyati and Micco (2007)	Z-score	H-statistic	Latin America, bank by bank	1993–2002	Positive effect of competition on risk

Beck et al. (2010) note that many results in the literature could be driven by the trend to consolidate during the Great Moderation of the early 2000s and/or may not hold in times of systemic global distress. Competitive systems might foster stability in normal times, for example, but contribute to bubbles, herding behavior, and the use of untested innovations in booms and credit crunches in times of recession or crisis. This implies a possible distinction between the long- and short-run relationships to competition, as well as before and after the GFC, that we test in our work.

Furthermore, competition may not have been the key factor underlying risk taking during the Great Moderation. It could have been the global liquidity glut and regulatory easing from Basel II (Barrell and Davis, 2008), as well as disaster myopia (e.g., measured by the time since the last crisis, as shown for East Asian banks by Craig et al., 2006) and risk-taking incentives on the part of the too big to fail. This can be tested by the relationship of competition to risk in our regressions. Furthermore, the broad issue of incentives has barely been applied in this field, although the work of Gropp et al. (2010) is an exception.

In this context, it could be suggested that the competition–fragility approach is characteristic of banking systems that have been recently deregulated, as in the case of the United States in the original franchise value study by Keeley (1990), or that are subject to other forms of structural change. However, as time passes, banks have become habituated to a deregulated system and the level of competition has become synonymous with stability. This pattern is in line with the industrial approach to financial instability proposed by Davis (1995, 1999), whereby a warning sign of financial instability is a change in entry conditions for institutions to financial markets, generating a rapid increase in competition after new entry. Mechanisms generating risk include the way new entrants to financial markets may induce borrowers to switch away from established credit relationships or offer extra credit to gain market share (by offering lower prices). Information-based linkages will thus be weakened and existing information devalued and risk increases overall (see Appendix A for more details). This effect will, however, be temporary and, in the long term, the market will settle down to a new level of competition.

Whereas this argument is clearly not fully consistent with the 2000–07 boom, the underlying relationship could be overlaid by these additional factors, which can be tested. Again, a corollary may be that the established level of competition is consistent with stability but abrupt changes in competition (e.g., after deregulation, as well as during periods of new entry, prolonged booms, or financial innovation) could have a negative impact. That is, a static–dynamic distinction arises again that has not been tested in the literature to date.

Furthermore, most studies are based on global samples. In Barrell et al. (2010) and Davis et al. (2011), we question whether this is the best approach for detecting the competition–risk trade-off, since the behavior of advanced countries (as in the EU) may differ from that of emerging market economies. It is telling that one of the most recent studies shows that 'an increase in competition is associated with a larger rise in banks' fragility in countries with stricter activity restrictions, lower systemic fragility, better developed stock exchanges, more generous deposit insurance and more effective systems of credit information sharing' (Beck et al., 2013, p. 219), some of which are advanced country characteristics. In addition, EU countries have a more homogeneous regulatory framework than a global sample does, which helps eliminate one cause of variation in the competition–stability trade-off.[1]

[1]In Figure 2 of Beck et al. (2013), all EU countries other than Latvia and Luxembourg have a positive trade-off between competition (measured by the Lerner index) and risk (measured by the Z-score) and the positive effect is significant, except for the Netherlands, Romania and Ireland (although the significant effects do vary from roughly 0.5 to 3.0).

There has not been much discussion in the recent literature of how macroprudential policies, such as loan-to-value limits, might relate to competition. The prevailing assumptions of macroprudential policies are generally based on the competition–fragility link and do not take into account variations in the trade-off shown in the literature survey. In addition, competition in wholesale funding has not been examined in most studies, particularly its relation to risk (with the exception of Demirgüç-Kunt and Huizinga, 2010). Studies typically focus on the banking system in a country (or country by country) *per se*. Intensive securities market competition and/or competition from shadow banking may interact with bank risk taking and competition in countries with more diverse financial systems. Hence, it is important to at least include indicators of securities market activity as control variables.

3 | DATA

We use data from Bankscope for the EU countries, ensuring a degree of commonality in terms of the regulatory framework. In particular, controls on entry should be low, helping to ensure a degree of contestability, while common minimum prudential standards are enforced across the EU. We include commercial, savings, cooperative and mortgage banks in our sample, but not investment banks. This approach is in line with the work of Schaeck and Čihák (2012), who test commercial banks versus a wider sample (not including mortgage banks, however) and conclude that ‘constraining the sample to profit maximizing institutions, although justified on theoretical grounds, is not necessary for the empirical tests’ (2012, p. 838). We have 6,008 banks from 27 EU countries over the period 1998–2012, thus including substantial periods both before and after the GFC. Usable observations typically number around 10 to 30,000. The banks are well distributed across countries, with Germany having the most banks, with just over 40% of the total, in contrast to studies such as that of Berger et al. (2009), where US banks account for 91% of the sample. Regression data using variables drawn from Bankscope are winsorized at the 1% level (as is common in the literature; e.g., Anginer et al., 2012). Details of the regression data and bank distributions by country are shown in Tables A1 to A3 in Appendix B.

Supplementing the Bankscope data, we use macroeconomic data from the World Bank's Financial Structure Database (Čihák et al., 2012a), which covers 1998–2011 in our sample. In particular, this information provides us with data on stock market value traded/percent of the gross domestic product to show the degree of securities market competition faced by banks. We also have dummies available for the legal origin of the country, from the revised dataset of La Porta et al. (2007).

Furthermore, we employ data from the World Bank's Bank Regulation and Supervision Surveys that took place in 1999, 2003, 2007 and 2011 and which are summarized in indices by Barth et al. (2012).[2] Potentially relevant data include, in particular, activity restrictions, limits on foreign banks, the fraction of applications denied, initial capital stringency, the overall index for capital regulation, supervisory power, a supervision index, multiple supervisors, a private monitoring index, a moral hazard index, the percentage of foreign banks, and an external governance index. For a discussion, see Barth et al. (2006). We construct a time series for these data following Beck et al. (2013), with each observation holding for the preceding year and the two following years.

[2]The indices are downloadable from http://faculty.haas.berkeley.edu/ross_levine/Regulation.htm.

4 | METHODOLOGY

We initially estimated revenue functions for the Panzar–Rosse H-statistic for each EU country. According to this approach, market power is measured by the extent to which changes in factor prices are reflected in revenues. With perfect competition and when banks operate within their long-run equilibrium, a proportional increase in factor prices (including the interest rate on liabilities) induces an equiproportional change in gross revenues. The output does not change in volume terms, while the output price rises to the same extent as the input price (i.e., demand is perfectly elastic). On the other hand, under monopolistic competition or where potential entry leads to a contestable market equilibrium, revenues will increase less than proportionally, since the demand for banking products facing individual banks is inelastic (Tirole, 1988). In the limiting case of a monopoly, there may be no response or even a negative response of gross revenues to changes in input costs.

Following Bikker et al. (2012) and in line with Panzar and Rosse (1987), we use an unscaled revenue function. Bikker et al. (2012) have shown that forms of scaling (e.g., including assets or equity on the right-hand side) or the use of a price and not a revenue variable on the left (e.g., revenue scaled by assets) upward-biases the H-statistic (i.e., imperfect competition is rejected too frequently). After extensive testing using 100,000 observations on 17,000 banks in 63 countries over 1994–2004, the authors find that price and scaled revenue functions cannot identify imperfect competition in the same way unscaled revenue functions can and that 'this conclusion disqualifies a number of studies since they apply a Panzar–Rosse test based on a price function or scaled revenue function' (Bikker et al., 2012, p. 1016).

Accordingly, our estimating equation for the H-statistic is as follows:

$$LogR_{it} = \Sigma^{J}{}_{(j=1)}\alpha_j Logw_{jit} + \Sigma^{N}{}_{(n=1)}\gamma_n LogX_{nit} + \varepsilon_{it} \tag{1}$$

for bank i at time t, where $t = 1, \ldots, T$, with T being the number of periods observed; $i = 1, \ldots, I$, with I being the total number of banks; and R_{it} is unscaled gross interest revenues. In our case, we have $J = 3$ inputs, so that w_{it} is a three-dimensional vector of factor prices (the logarithm of the ratio of interest expense to total debt funding, *IED*; the logarithm of the ratio of personnel expenses to total assets, *PTA*; and the logarithm of the ratio of other costs as a proportion of fixed assets, *OCF*), consistent with the intermediation approach to banking output measurement, where bank liabilities are inputs to produce loans and other earning assets. The term X_{it} is a vector of exogenous and bank-specific variables that may shift the cost and revenue schedule (business mix). In this context, we have $N = 4$, the logarithm of loans as a proportion of assets, *LAR*, showing credit risk (with an expected positive sign, since banks compensate for risk); the logarithm of the ratio of other non-earning assets to total assets (*OTA*), reflecting asset composition; the logarithm of customer deposits as a proportion of deposits plus money market liabilities (*CDT*), showing liquidity risk (but whose sign is ambiguous); and the logarithm of equity to total assets (*ETA*), showing leverage and hence risk preferences (expected to have a negative sign).

We first estimate the H-statistic by country and in subperiods, using the within estimator, with both bank and year fixed effects in line with the results of De Bandt and Davis (2000), as well as pooled feasible generalised least squares (FGLS) using White's (1980) method to reduce the impact of heteroskedasticity. We also test for market equilibrium in sub-periods (results for these estimates are in Davis and Karim (2013)). For the current exercise, we then estimate the H-statistic as an annual time series for each individual country. We apply the restriction of at least 12 banks per year.

We then estimate the Lerner index for the EU as a whole, following Anginer et al. (2012). Beck et al. (2013) and Weill (2013) and have tested it as a competition indicator. Accordingly, we first estimate the following translog cost function:

$$
\begin{aligned}
log(C_{it}) &= \alpha + \beta_1 log(TA_{it}) + \beta_2 \left(log(TA_{it})\right)^2 + \beta_3 log(W_{1,it}) + \beta_4 log(W_{2,it}) + \beta_5 log(W_{3,it}) \\
&+ \beta_6 log(TA_{it}) log(W_{1,it}) + \beta_7 log(TA_{it}) log(W_{2,it}) + \beta_8 log(TA_{it}) log(W_{3,it}) \\
&+ \beta_9 \left(log(W_{1,it})\right)^2 + \beta_{10} \left(log(W_{2,it})\right)^2 + \beta_{11} \left(log(W_{3,it})\right)^2 + \beta_{12} log(W_{1,it}) log(W_{2,it}) \\
&+ \beta_{13} log(W_{1,it}) log(W_{3,it}) + \beta_{14} log(W_{2,it}) log(W_{3,it}) + \Theta Year\ Dummies + \varepsilon it
\end{aligned} \tag{2}
$$

where C_{it} is total costs and TA_{it} is the quantity of output and is measured as total assets. Our input prices are $W_{1,it}$, which is the ratio of interest expenses to the sum of total deposits and money market funding (IES); $W_{2,it}$ is measured as personnel expenses divided by total assets (PTA); and $W_{3,it}$ is the ratio of other operating expenses to fixed assets (OCF). Having estimated this equation, we impose the following restrictions, again in line with the earlier authors, to ensure homogeneity of degree one in input prices:

$$
\beta_3 + \beta_4 + \beta_5 = 1;\ \beta_6 + \beta_7 + \beta_8 = 0;\ \beta_9 + \beta_{12} + \beta_{13} = 0;\ \beta_{10} + \beta_{12} + \beta_{14} = 0;\ \beta_{11} + \beta_{13} + \beta_{14} = 0 \tag{3}
$$

We then use the coefficient estimates from the previous regression to estimate the marginal costs for bank i in calendar year t:

$$
MC_{it} = \delta C_{it}/\delta TA_{it} = C_{it}/TA_{it} \times [\beta_1 + 2 \times \beta_2 \times log(TA_{it}) + \beta_6 \times log(W_{1,it}) + \beta_7 \times log(W_{2,it}) + \beta_8 \times log(W_{3,it})] \tag{4}
$$

The Lerner index for each bank–year is:

$$
Lerner_{it} = (P_{it} - MC_{it})/P_{it} \tag{5}
$$

where Pit is the price of assets and is equal to the ratio of total revenue to total assets.

To save space, we do not include the details of the estimates for the H-statistic and Lerner index as outlined above, although the descriptive statistics for the dependent and independent variables are included in Appendix B.[3] We then relate these annual competition variables to indicators of bank and systemic risk, controlling for relevant variables. Our core results, in line with the bulk of the literature, link competition each year to the logarithm of the Z-score for individual banks, which is defined as the return on assets (*ROA*) plus the leverage ratio, divided by the standard deviation of the return on assets over three years. As Liu et al. (2013) note, it is appropriate to use the logarithm of the Z-score, since the level is highly skewed while the logarithm is normally distributed. We assess the H-statistic and Lerner index as measures of competition in terms of both levels and differences, to distinguish between levels of competition and change in competition. To our knowledge, this has not been done in the literature and could capture important distinctions between long-run and dynamic aspects. The current difference of the H-statistic and the Lerner index is complemented by the second and third lags of their levels, thus avoiding any overlap between levels and differences and possible false conclusions. Since the H-statistic is a countrywide variable, we did not consider it to be correlated with bank-level risk and, accordingly, do not instrument the current difference, whereas we do so for the Lerner index.

[3]The regression results are available from the authors upon request.

We run three sets of estimates: with bank-level variables only, with bank-level variables and country dummies, and with additional macro-level control variables. In each case, we seek to shadow the best practices of Beck et al. (2013):

$$\begin{aligned} Log\, Z_{it} &= a_0\Delta H_{jt} + a_1 H_{jt-2} + a_2 H_{jt-3} + a_3 CDT_{it-1} + a_4 LAR_{it-1} + a_5 NIR_{it-1} + a_6 log(TA)_{it-1} \\ &+ a_7 PII_{it-1} + a_8\Delta log(TA)_{it-1} + a_9 SMT_{jt} + a_{10} CSI_{jt} + a_{11} ACT_{jt} + a_{12} LO_{jt} + \varepsilon_{it} \end{aligned} \tag{6}$$

and

$$\begin{aligned} Log\, Z_{it} &= a_0\Delta Lerner(instrumented)_{it} + a_1 Lerner_{it-2} + a_2 Lerner_{it-3} + a_3 CDT_{it-1} + a_4 LAR_{it-1} \\ &+ a_5 NIR_{it-1} + a_6 log(TA)_{it-1} + a_7 PII_{it-1} + a_8\Delta log(TA)_{it-1} + a_9 SMT_{jt} + a_{10} CSI_{jt} + a_{11} ACT_{jt} \\ &+ a_{12} LO_{jt} + \varepsilon_{it}. \end{aligned} \tag{7}$$

Accordingly, besides the H-statistic and the Lerner index, bank-level control variables in the risk function (denoted with the subscript *i*) are customer deposits as a proportion of deposits plus money market liabilities (*CDT*), the loan-to-asset ratio (*LAR*), the ratio of non-interest revenue to interest revenue (*NIR*), bank size (*log TA*), the ratio of provisions to interest income (*PII*), and the growth rate of assets (*d log TA*). All bank-specific variables (denoted with the subscript *i*) are lagged one year to avoid simultaneity.

Regarding macro-level control variables at the country level (denoted with the subscript *j*), after testing, we control for stock market turnover (*SMT*), which indicates the scope for securities market financing, and, in terms of regulation, capital stringency (*CSI*) and activity restrictions (*ACT*), as well as legal origin (*LO*). Since EU regulations are relatively homogeneous, we do not expect to find major effects of regulation *per se* in our work.[4] As for competition, we estimate using the within estimator with year fixed effects as well as pooled FGLS using White's (1980) method to reduce the impact of heteroskedasticity. Given our use of lags for bank-specific variables, we contend that this approach is more appropriate and reliable than the generalized method of moments.

We note that the literature on competition and risk in banking makes virtually no reference to panel unit roots. This omission likely relates to the fact that the time dimension is small and the number of cross sections is very large, while most panel unit root testing focuses on time series and cross section values of reasonable size, as in a cross-country macro dataset. Note that we did run tests on the stationarity of the key variables (logarithm of the Z-score at the bank level, the H-statistic at the country level, and the Lerner index at the bank level) and we find that they were stationary on the principal tests (Levin–Lin–Chu, Im–Pesaran–Shin, Fisher augmented Dickey–Fuller, and Fisher Phillips–Perron). This result justifies our specification with, for example, the level of the dependent variable.

As a robustness check, we run the various regressions with an alternative measure of risk, which is impaired loans as a proportion of total loans. In bank-dominated systems, as in much of the EU, this is a fairly accurate measure of risk for all but the largest banks (which hold proportionately more securities). There were, however, much fewer observations, especially before 2007, making the earlier estimates relatively unreliable. We also run further robustness checks with bank fixed effects, clustered standard errors and bootstrapped standard errors.

[4]Indeed, in a study of EU risk and competition down to the regional level, Liu et al. (2013) omit all regulatory controls.

5 | ESTIMATION

Having estimated the H-statistic year by year, as outlined above, to obtain a time series for the H-statistic for each country, we then winsorize the H-statistic at 95%, given the number of outliers resulting from the year-by-year estimation procedure and the lack of scaling for the revenue function. This approach gives a range from roughly +1 (behavior in line with perfect competition) to −2 (monopolistic behavior).[5] To compare and contrast with our results for the H-statistic, we go on to estimate the Lerner index, showing price–marginal cost margins for banks in the EU as described above, deriving the marginal cost for banks in each year of the sample.[6] The Lerner index then provides an alternative measure of competition in the banking system, with narrower margins tending to accompany an increase in competition.

Beginning with the results for the H-statistic, as noted, we estimated the current difference of the H-statistic and levels at lags 2 and 3. This approach also avoids spurious results from overlapping differences and levels. It has a natural interpretation in terms of the short-run relationship to changes in competition being estimated separately from the long-run relationship to levels of competition, in line with the discussion in sections 1 and 2. As noted, since the H-statistic is a macro variable, we do not consider it likely to be highly correlated with individual bank Z-scores, so we do not instrument its current difference. We start with work on the pre-crisis period where a link from competition to risk would have offered an early warning. Accordingly, the basic Z-score results for all the countries and the pre-crisis period 1998–2006 are shown in Table 2.

Recall that a higher Z-score indicates a less risky bank (i.e., with higher profitability and/or capital and less volatile profits). The core result shows that the difference of the H-statistic is negative and significant, whereas the lagged levels of the H-statistic are positive and significant. The short-run effect is shown by the coefficient of the difference, which is −0.036 in this case, while the long-run effect is the sum of the significant coefficients for the levels, which, in this case, is 0.263. Accordingly, a change in the level of competition is harmful to banks' solvency, consistent with slow adaptation to change and disaster myopia during periods of apparent high profitability. On the other hand, this result suggests that the level of competition *per se* is not a cause of risk. Banks can adapt to competitive conditions and remain solvent, for example, by holding more capital, as consistently found by Schaeck and Čihák (2012).

Regarding the other bank-specific variables, a higher share of deposits in total short-term funding reduces risk, a result that was strongly borne out later during the GFC, when wholesale funding dried up (a similar result was noted by Demirgüç-Kunt and Huizinga, 2010). A greater loan share in total assets also reduces risk, perhaps reflecting volatile securities holdings or greater risks run by large banks, which have lower loan-to-asset ratios. A higher ratio of non-interest income is negative for solvency (Demirgüç-Kunt and Huizinga, 2010), since the earlier promises that non-interest income could stabilize profits were not borne out over the data period. Again, it is large banks that tend to have greater non-interest income. The logarithm of total assets is not significant, but a rise in total assets strongly raises risk (perhaps reflecting adverse selection when assets rise sharply). Finally, the ratio of provisions to interest income is directly and strongly negatively related to risk. The results for the

[5]Testing for market equilibrium in sub-periods 1998–2006 and 2007–12, as well as 1998–2012 (Davis and Karim, 2013), we find most countries are shown to be in equilibrium in both the full period and the sub-periods. The key exceptions are Germany and Sweden, which fail the test in all sub-periods. Latvia and Lithuania also fail in the sub-periods, while, during the crisis, disequilibrium is shown for Estonia, Italy, Ireland, the Netherlands and Slovenia.

[6]The estimates are available from the authors upon request.

provisions ratio, change in assets, and non-interest share are also consistent with those of Beck et al. (2013, Table 5). However, for a global sample, Beck et al. find a positive link of size to stability that is not present in our EU sample. Note also that the authors use the Lerner index and not the H-statistic as a competition/market power indicator (we experiment with the Lerner index below).

Table 3 presents the results for a regression with country dummies as well, hence, capturing effects on the average Z-score that are country specific and not explained by other variables.

We leave out the dummy for one country (Germany), since it is necessary for identification; the dummies are not included in Table 3. Our results are consistent with those in Table 2 and the H-statistic results are again significant at the 95% level for the dynamic term and the second lag level term, although the third lag is insignificant. The coefficient for the growth of total assets also becomes insignificant in this case.

Our third main set of results, shown in Table 4, is for a regression including macro variables relating to financial structure and regulation. Following a search, we include the stock market turnover ratio, dummies for legal origin, and regulation variables for activity restrictions and stringency of capital regulations.

The results here are consistent with those in Tables 2 and 3, with the difference for the H-statistic significant at the 1% level, while the level effect is now significant at the second lag at the 10% level. Again, the total assets variable is insignificant. Regarding legal origin, we omit the French legal origin, since it covers the majority of EU countries. A British or German legal origin is shown to be associated

TABLE 2 Log Z-score results for the EU, 1998–2006 (dependent variable = log Z-score)

This table reports the regression results for EU banks from 1998 to 2006 with the logarithm of the Z-score as the dependent variable, estimated using the within estimator and pooled FGLS, with year fixed effects and White's cross-sectional standard errors and covariance (corrected for degrees of freedom). The variables are defined as follows: *H* is the Panzar–Rosse H-statistic for the country and year in question; *CDT* deposits as a share of short-term funding; *LAR* is the loan-to-assets ratio; *NIR* is the ratio of non-interest revenue to interest revenue; *TA* is bank size (total assets); and *PII* is provisions to interest income. The term Δ indicates the first difference. All variables are winsorized at 99%, except *H* (95%). The *t*-values are in parentheses. The superscripts ***, ** and * indicate significance at the 1%, 5% and 10% levels, respectively.

Variable	Coefficient
$\Delta H(t)$	−0.036** (2.0)
$H(t–2)$	0.148*** (3.6)
$H(t–3)$	0.115** (2.0)
$CDT(t–1)$	0.203*** (3.2)
$LAR(t–1)$	0.612*** (7.9)
$NII(t–1)$	−0.427*** (17.6)
$LogTA(t–1)$	0.0137 (0.7)
$\Delta LogTA(t–1)$	−0.269*** (2.7)
$PII(t–1)$	−0.916*** (4.3)
C	3.19*** (11.7)
Adj-R^2	**0.077**
Observations	**11,363**
Banks	**2,701**

TABLE 3 Log Z-score results for the EU with country dummies, 1998–2006

This table reports the regression results for EU banks from 1998 to 2006 with the logarithm of the Z-score as the dependent variable, estimated as shown in the header of Table 2. Country dummies (excluding Germany) are also included in the regression. The variables are defined as in Table 2. The term Δ indicates the first difference. All variables are winsorized at 99%, except *H* (95%). The *t*-values are in parentheses. The superscripts ***, ** and * indicate significance at the 1%, 5% and 10% levels, respectively.

Variable	Coefficient
$\Delta H(t)$	−0.052** (2.4)
$H(t–2)$	0.135** (2.6)
$H(t–3)$	0.093 (1.5)
$CDT(t–1)$	0.177*** (2.8)
$LAR(t–1)$	0.462*** (6.2)
$NII(t–1)$	−0.447*** (21.0)
$LogTA(t–1)$	0.0177 (0.8)
$\Delta LogTA(t–1)$	−0.168 (1.3)
$PII(t–1)$	−0.83*** (4.1)
C	3.25*** (10.3)
Adj-R^2	**0.1**
Observations	**11,363**
Banks	**2,701**

with higher Z-scores, on average, while a Scandinavian legal origin has a similar effect as a French one (during the pre-crisis period). Authorities imposing tighter activity restrictions are associated with lower Z-scores and hence less stable banks.

It is of interest to see whether the result shown in Tables 2–4 above of a positive H-statistic in the level and negative in the difference is stable in different samples. We show a variety of estimates, including for the later sub-period and the full sample, in Table 5.

Whereas the three specifications have consistent significance for the level and difference across the earlier period, 1998–2006 (reproducing the results in Tables 2–4), this is not the case for the full period (1998–2012) or the later sub-period (2007–12), where the level effect dominates, although the sign of the insignificant difference term remains negative. We contend that, in the short run, there remains a need for caution regarding risk when competition increases. This is in line with the competition–fragility approach, which finds wide support elsewhere in the literature. Meanwhile, the consistent finding of a positive long-run relationship of competition with soundness offers support for the competition–stability approach in the long run, in line with Anginer et al. (2012), among others. In terms of magnitude, the results imply that a one-year shift from a position of low competition ($H = 0.0$) to a higher level ($H = 0.5$)[7] would induce a drop in the Z-score of around 2–4% in the short run, while the Z-score would rise in the long run by 5–15%. We consider these magnitudes plausible.

[7]These values are typical of countries with low and high levels of banking competition over the data period, as shown in Davis and Karim (2013).

TABLE 4 Log Z-score results for the EU with macro variables, 1998–2006

This table reports the regression results for EU banks from 1998 to 2006 with the logarithm of the Z-score as the dependent variable, estimated as shown in Table 2., The variable *SMT* indicates stock market turnover, *LOBRIT* British legal origin, *LOSCAND* Scandinavian legal origin, *LOGER* German legal origin, *ACT* activity restrictions, and *CSI* capital stringency. The remainder of the variables are defined as in Table 2. The term Δ indicates the first difference. All variables are winsorized at 99%, except *H* (95%). The *t*-values are in parentheses. The superscripts ***, ** and * indicate significance at the 1%, 5% and 10% levels, respectively.

Variable	**Coefficient**
$\Delta H(t)$	−0.074*** (3.8)
$H(t–2)$	0.093* (1.8)
$H(t–3)$	0.04 (0.5)
$CDT(t–1)$	0.179*** (2.9)
$LAR(t–1)$	0.535*** (6.3)
$NII(t–1)$	−0.422*** (16.9)
$LogTA(t–1)$	0.0121 (0.6)
$\Delta LogTA(t–1)$	−0.211* (1.8)
$PII(t–1)$	−0.888*** (4.6)
C	3.55*** (12.9)
$SMT(t)$	0.000752 (1.4)
$LOBRIT(t)$	0.246** (2.6)
$LOSCAND(t)$	0.112 (1.0)
$LOGER(t)$	0.0802* (1.8)
$ACT(t)$	−0.0822*** (2.9)
$CSI(t)$	0.00268 (0.1)
Adj-R^2	0.084
Observations	**11,340**
Banks	**2,686**

A natural comparison with the results above is to then include the Lerner index instead of the H-statistic in the equation for the logarithm of the Z-score. In the work of Beck et al. (2013), for example, the current level of the Lerner index is consistently positive, since a high margin indicates a safer bank and hence a higher Z-score and vice versa. However, we contend that such a result is contestable: As Beck et al. point out, there is an element of circularity in the argument, since the Lerner index itself includes the return on assets, which, in turn, is strongly related to the price–cost margin, as indicated by the Lerner index (correlation of 0.5). Accordingly, the Lerner index should at least be instrumented to avoid bias from this simultaneity if the current level is used and otherwise lagged. Therefore and in line with our work for the H-statistic, we include the current difference and the second and third lags for the Lerner index, winsorized at the 99% level. To avoid simultaneity, we instrument the current difference of the Lerner index with the first and second lagged differences. The results are shown in Table 6.

The first difference term is consistent with the result for the differenced H-statistic, in the sense that, given a positive sign, a rise in competition (reduction in the Lerner index) links to a decline in the margin and less stable banks. This result applies consistently across the full and later period samples, in

TABLE 5 Results for the H-statistic for different periods (dependent variable = log Z-score)

This table reports the regression results for EU banks from 1998 to 2006, from 1998 to 2012, and from 2007 to 2012, with the logarithm of the Z-score as the dependent variable, estimated as indicated in Table 2. The term *H* is the Panzar–Rosse H-statistic for the country and year in question. Additional variables are as in Tables 2 to 4 and Δ indicates the first difference. All variables are winsorized at 99%, except *H* (95%). The *t*-values are in parentheses. The superscripts ***, ** and * indicate significance at the 1%, 5% and 10% levels, respectively.

	Basic (as in Table 2)			**Country dummies (as in Table 3)**			**Macro (as in Table 4)**		
Variable	1998–2006	1998–2012	2007–12	1998–2006	1998–2012	2007–12	1998–2006	1998–2012	2007–12
$\Delta H(t)$	−0.036** (2.0)	−0.052 (1.1)	−0.081 (1.0)	−0.052** (2.4)	−0.051 (1.3)	0.0 (0.0)	−0.074*** (3.1)	−0.035 (0.9)	−0.246 (0.4)
$H(t–2)$	0.148*** (3.7)	0.148*** (3.4)	0.11 (1.4)	0.135*** (2.7)	0.075 (1.5)	0.071 (0.8)	0.093* (1.8)	0.072* (1.8)	0.074* (1.9)
$H(t–3)$	0.115** (2.0)	0.177*** (4.2)	0.206*** (3.3)	0.093 (1.5)	0.126*** (3.4)	0.115** (2.5)	0.04 (0.6)	0.089 (1.6)	0.11* (1.8)
Adj-R^2	0.08	0.11	0.14	0.1	0.15	0.21	0.08	0.09	0.13
Observations	11,363	28,301	16,938	11,363	28,316	16,973	11,340	23,246	11,906
Banks	2,701	4,219	3,607	2,701	4,219	3,609	2,686	4,130	3,432

TABLE 6 Results for the Lerner index for different periods

This table reports the regression results for EU banks from 1998 to 2006, from 1998 to 2012, and from 2007 to 2012, with the logarithm of the Z-score as the dependent variable, estimated as indicated in Table 2. Here, Δ*Lerner_inst* is the current difference of the Lerner index instrumented by two lags of itself, while *Lerner* is the Lerner index itself. Additional variables are as in Tables 2 to 4 and Δ indicates the first difference. All variables are winsorized at 99%. The *t*-values are in parentheses. The superscripts ***, ** and * indicate significance at the 1%, 5% and 10% levels, respectively.

	Basic (as in Table 2)			**Country dummies (as in Table 3)**			**Macro (as in Table 4)**		
Variable	1998–2006	1998–2012	2007–12	1998–2006	1998–2012	2007–12	1998–2006	1998–2012	2007–12
Δ*Lerner_inst*(*t*)	0.174 (0.9)	0.33* (1.7)	0.41 (1.5)	0.19 (1.1)	0.33** (2.0)	0.4** (2.0)	0.17 (1.0)	0.39* (1.9)	0.54* (1.9)
Lerner(*t*–2)	0.744*** (3.1)	0.67* (1.6)	0.58 (0.9)	0.76** (2.7)	0.72** (2.1)	0.54 (1.1)	0.95*** (3.2)	0.69** (2.1)	0.4 (0.8)
Lerner(*t*–3)	−0.3 (1.3)	−0.34 (1.0)	−0.31 (0.6)	−0.3 (1.1)	−0.21 (0.7)	−0.17 (0.5)	−0.24 (0.8)	−0.02 (0.1)	0.16 (0.4)
Adj-R^2	0.07	0.09	0.12	0.1	0.13	0.18	0.08	0.1	0.13
Observations	10,193	24,263	14,070	10,193	24,263	14,070	10,161	19,558	9,397
Banks	2,298	3,795	3,179	2,298	3,795	3,179	2,291	3,707	3,062

contrast to the dynamic result for the H-statistic, which is largely confined to the 1998–2006 sample, although it is not significant for the basic specification in 2007–12. Meanwhile, concerning levels, it is the second lag that is significant, except for 2007–12, again with a positive sign, implying a long-run negative impact of competition on risk. Accordingly, the Lerner index favors the competition–fragility hypothesis in both the short and long run, the latter being in line with the work of authors such as Fu et al. (2014). In terms of magnitude, the results imply that a drop in margins of 0.1 (half of the mean level of 0.2) links to a reduction in the Z-score of around 3–5%, whereas, in the long run, the reduction is 7–9%. We consider these values economically plausible.

Hence, we have similar results between the two competition indicators for the impact of changes in competition on bank risk, namely, that a rise in competition leads to deterioration in bank soundness, as measured by the Z-score (competition–fragility). On the other hand, the H-statistic consistently shows a long-run negative relationship between competition and risk (i.e., banks, in the long run, are safer in competitive markets, or competition–stability), while the Lerner statistic indicates a negative long-run relationship for the same banks (i.e., banks, in the long run, are less safe in markets with narrower price–cost margins, or competition–fragility).

As a first robustness check of this result, we run the basic regressions for both the H-statistic and the Lerner index (i.e., with extra variables, as in Table 2) for a different dependent variable, namely, the ratio of impaired loans to total loans. This measure has a much smaller coverage than the Z-score, with less than 10,000 observations and rather few before 2007. The results for the H-statistic and the Lerner index are as shown in Table 7.

There is a broad tendency for the relationship of each variable to impaired loan ratios to be negative in the long run. The dynamics are less significant than with the Z-score, perhaps partly reflecting the lack of observations in the earlier sample. Accordingly, the data indicate that a higher level of competition in the long run (a higher H-statistic) leads to fewer impaired loans and hence less risk (competition–stability). On the other hand, the Lerner index results for both the short and long run imply that the narrower the margin (i.e., the greater the competition), the higher the impaired loan rate (competition–fragility). This pattern reverts to zero after three years in the later samples, however, probably reflecting cyclical patterns and the effect of the GFC.

We consider the macro equation most appropriate for further robustness checks, since it allows for regulatory and structural factors and, accordingly, controls for omitted variable bias. The checks are threefold. First, we provide estimates with bank-level fixed effects (bearing in mind the above results have time dummies and in some cases country dummies). It can be argued that this will help distinguish short-run impacts of competition from long-run impacts. Including bank fixed effects is also important because of the sample's high heterogeneity, as in different types of banks. Second, we provide estimates using country-clustered standard errors. The reason is that, since the H-statistic is a country-level variable, the error term should be clustered at the country level to allow for potential correlations in bank risk. Third, since the H-statistic and the Lerner index are calculated with estimated coefficients from first-stage regressions, one needs to pay attention to the extra variance introduced by the first-stage estimation. One way to correct the standard error in the second-stage regressions is to use bootstrapping, which we have also done.

Summarizing results shown in Tables 8 and 9, consistent with the baseline results above, we find a significant short-run link of the H-statistic to increased risk for all sub-periods with the bank dummies and for 1998–2006 and 1998–2012 for clustered standard errors and bootstrapping, respectively. In the long run, the H-statistic is significantly linked to reduced risk for the bank dummies for the second lag in the earlier period. For clustered standard errors and bootstrapping, the H-statistic is significant in the long run for all samples for the second lag and for all but the earlier period for the third lag. Regarding the Lerner index, the short-run effect linked to increased risk is

TABLE 7 Robustness check – Results for competition indicators using the impaired loan ratio as the risk variable

This table reports the regression results for EU banks from 1998 to 2006, from 1998 to 2012, and from 2007 to 2012, with the impaired loan ratio as the dependent variable, estimated as indicated in Table 2. The variable *H* is the Panzar–Rosse H-statistic for the country and year in question, Δ*Lerner_inst* is the current difference of the Lerner index instrumented by two lags of itself, while *Lerner* is the Lerner index itself. Additional variables are as in Table 2 and Δ indicates the first difference. All variables are winsorized at 99% except *H* (95%). The *t*-values are in parentheses. The superscripts ***, ** and * indicate significance at the 1%, 5% and 10% levels, respectively.

H-Statistic			
Variable	1998–2006	1998–2012	2007–2012
Δ*H*(*t*)	−0.00293 (0.4)	−0.00232 (0.4)	−0.0098 (0.1)
H(*t*–2)	−0.00367 (1.3)	−0.00839*** (3.5)	−0.00959*** (3.0)
H(*t*–3)	−0.0024 (0.3)	−0.00831** (2.2)	−0.00845** (1.9)
Adj-R^2	0.039	0.32	0.34
Observations	797	7,408	6,611
Banks	365	2,033	1,908
Lerner index			
Variable	**1998–2006**	**1998–2012**	**2007–2012**
Δ*Lerner_inst*(*t*)	−0.0227 (0.8)	−0.022* (1.9)	−0.023** (2.0)
Lerner(*t*–2)	−0.1377*** (6.1)	−0.063*** (2.6)	−0.05** (1.9)
Lerner(*t*–3)	−0.00437 (0.3)	0.067** (2.2)	0.083** (2.5)
Adj-R^2	0.12	0.33	0.36
Observations	636	5,784	5,148
Banks	229	1,800	1,683

present in all sub-periods except 1998–2006 as for the basic results, while for the long run, there is again a positive effect for the second lag except for the bank dummies in 1998–2012 where it is insignificant, and in 2007–2012 where it has a significant negative relationship. Generally, the robustness checks strongly support the main results.

6 | ASSESSING THE DIFFERENCES IN RESULTS

Given the consistent differences in long-run relationships of the H-statistic and the Lerner index to risk, we conclude by investigating the reasons for such differing long-term results. There are conceptual differences between the H-statistic and the Lerner index, as shown by Carbo et al. (2009), in that the former is a difference term (elasticity of revenue to prices) and the latter is a level effect (price to marginal cost margin). Furthermore, the H-statistic is a macroeconomic index describing the situation in the banking sector as a whole, while the Lerner index describes the price cost margin of an individual bank.

We seek to find the reasons for the differing predictions by regressing the level of each indicator of competition on the sub-components of the Z-score. These sub-components are the return on assets (*ROA*), which as noted is a measure of performance; capital adequacy, a measure of safety and soundness and the volatility of the return on assets, a measure of risk. We estimated these relationships

TABLE 8 Robustness checks with banking dummies, clustered standard errors, and bootstrapping, results for the H-statistic with macro variables (dependent variable = log Z-score)

This table reports the regression results for EU banks from 1998 to 2006, from 1998 to 2012, and from 2007 to 2012, with the logarithm of the Z-score as the dependent variable, 1) estimated as indicated in Table 2, with bank fixed effects; 2) estimated as indicated in Table 2, with standard errors clustered at the country level; and 3) estimated using bootstrapping with year fixed effects. The variable *H* is the Panzar–Rosse H-statistic for the country and year in question. Additional variables are as in Table 4, except the legal origin dummies for the regression with bank dummies. The term Δ indicates the first difference. All variables are winsorized at 99%, except *H* (95%). The *t*-values (*Z*-values for bootstrapping) are in parentheses. The superscripts ***, ** and * indicate significance at the 1%, 5% and 10% levels, respectively.

	(1) Bank dummies			**(2) Clustered standard errors**			**(3) Bootstrapping**		
Variable	1998–2006	1998–2012	2007–12	1998–2006	1998–2012	2007–12	1998–2006	1998–2012	2007–12
$\Delta H(t)$	−0.057** (2.2)	−0.062** (2.5)	−0.075** (2.1)	−0.074*** (2.8)	−0.035** (2.1)	−0.025 (1.0)	−0.074*** (3.2)	−0.035** (2.0)	−0.025 (1.0)
$H(t–2)$	0.073* (1.8)	−0.036 (0.7)	−0.081 (0.9)	0.093*** (3.5)	0.072*** (4.4)	0.074*** (3.4)	0.093*** (3.4)	0.072*** (4.0)	0.074*** (3.8)
$H(t–3)$	0.083 (1.1)	−0.005 (0.1)	−0.058 (0.8)	0.039 (1.4)	0.089*** (6.1)	0.111*** (6.1)	0.04 (1.4)	0.089*** (6.5)	0.111*** (6.7)
Adj-R^2	0.425	0.376	0.45	0.082	0.094	0.123	0.082	0.094	0.122
Observations	11,340	23,246	11,906	11,340	23,246	11,906	11,340	23,246	11,906
Banks	2,686	4,130	3,492	2,686	4,130	3,492	2,686	4,130	3,492

TABLE 9 Robustness checks with banking dummies, clustered standard errors, and bootstrapping, results for the Lerner index with macro variables (dependent variable = log Z-score)

This table reports the regression results for EU banks from 1998 to 2006, from 1998 to 2012, and from 2007 to 2012, with the logarithm of the Z-score as the dependent variable, 1) estimated as indicated in Table 2, with bank fixed effects; 2) estimated as indicated in Table 2, with standard errors clustered at the country level; and 3) estimated using bootstrapping, with year fixed effects. The term Δ*Lerner_inst* is the current difference of the Lerner index instrumented by two lags of itself, while *Lerner* is the Lerner index itself. Additional variables are as in Table 4, except the legal origin dummies for the regression with bank dummies. The term Δ indicates the first difference. All variables are winsorized at 99%. The *t*-values (*Z*-values for bootstrapping) are in parentheses. The superscripts ***, ** and * indicate significance at the 1%, 5% and 10% levels, respectively.

	(1) Bank dummies			(2) Clustered standard errors			(3) Bootstrap		
Variable	1998–2006	1998–2012	2007–12	1998–2006	1998–2012	2007–12	1998–2006	1998–2012	2007–12
Δ*Lerner_inst* (*t*)	0.08 (0.6)	0.389** (3.3)	0.556*** (3.0)	0.169 (1.4)	0.389*** (4.9)	0.542*** (9.6)	0.169 (1.4)	0.389*** (4.0)	0.542*** (5.7)
Lerner(*t*–2)	0.467* (1.8)	−0.147 (0.5)	−1.27** (2.3)	0.952*** (4.1)	0.691*** (4.5)	0.399** (3.9)	0.952*** (4.4)	0.691*** (4.1)	0.399** (2.0)
Lerner(*t*–3)	−0.294 (1.1)	−0.031 (0.1)	−0.377 (1.0)	−0.24 (1.1)	−0.022 (0.2)	0.16 (0.8)	−0.239 (1.2)	−0.022 (0.1)	0.16 (0.9)
Adj-R^2	0.405	0.375	0.478	0.079	0.12	0.188	0.082	0.0916	0.127
Observations	10,161	19,558	9,397	10,161	19,559	9,397	10,161	19,558	9,397
Banks	2,291	3,707	3,062	2,291	3,706	3,061	2,291	3,706	3,061

using a specification similar to that in Table 2 (i.e., including the basic control variables) and a lagged level of the relevant competition indicator estimated over the whole sample (we find the same results hold over each sub-sample as well). The results are shown in Table 10.

We multiply the Lerner index by −1 so that, in both cases, an increase in the indicator is consistent with higher competition. The results show that both the H-statistic and the Lerner index* −1 are negatively related to both *ROA* and capital adequacy. In other words, increased competition on both measures is related to lower profits and lower capital cover, thus reducing the Z-score. On the other hand, we find a difference in the relationship to the standard deviation of profitability, with the H-statistic having a negative relationship (higher levels of market competition are related to a lower volatility of profits), and *Lerner* * −1 having a positive relationship (narrower price–cost margins for an individual bank accompany an increase in the volatility of profits). The outcome of these different effects for the Z-score, the overall measure of risk, is that greater market competition (as indicated by the H-statistic) links to less risky positions in the long run, while narrower price–cost margins (indicated by the Lerner index) lead to increased risk. The same results, including the contrast for profit volatility, apply in all cases and with similar significance for the two sub-periods (1998–2006 and 2007–12). In other words, it is not an artefact of either the boom period or the later financial crisis.

A full investigation of reasons for these differences is beyond the scope of this paper. However, note, first, that the H-statistic is a countrywide indicator, while the Lerner index is a bank-specific one. While there is a general tendency for greater competition to lead to narrower margins, this need not always be the case (e.g., existing monopolists seeking a 'quiet life' and obtaining their rents by allowing their banks to remain inefficient; see Koetter et al., 2012). Numerically, narrower margins are likely to lead to greater profit volatility, since they are closer to zero, while the link between profit volatility and market competition is less direct, given the latter is measured by the response of revenue to input prices.

Taking the results at face value, this section suggests overall a need for caution in drawing policy conclusions from risk–competition studies without carefully considering the likely impact of a given policy shift. For example, the separation of retail and wholesale banking or certain macroprudential policies could have different effects on margins (as shown by the Lerner index) as opposed to market competition (as shown by the H-statistic). Margin effects consistently increase risk, whereas our results suggest that more competition is favorable for soundness in the long run.

TABLE 10 Relationship between the H-statistic, the Lerner index, and the components of the Z-score (1998–2012)

This table reports the sign of the coefficient of the lagged H-statistic (*H*) and the lagged Lerner index (*Lerner*) in regressions on the dependent variables for bank return on assets (*ROA*), bank capital adequacy, the standard deviation of the return on assets over the previous three years (*St. Dev. (ROA)*), and the Z-score. The estimation method and other independent variables are as in Table 2. All effects are significant at the 1% level.

Dependent variable	*Lag H*	*Lag Lerner* (* −1)
ROA	–	–
Capital adequacy	–	–
St. Dev. (*ROA*)	–	+
Z-Score	+	–

7 | CONCLUSIONS

This is, to our knowledge, the first empirical study of banking competition and risk to allow changes in competitive conditions, as well as levels, to impact on risk. It is also one of the first studies to assess comparable periods before and after the GFC and to compare and contrast the results using two competition indicators: the H-statistic (which indicates the scope of competition in the country concerned) and the Lerner index (which indicates profit margins bank by bank). In doing so, we highlight further nuances in the relationship between competition and stability in banking that have not been emphasized in the literature to date. We employ an extensive dataset of banks from 27 EU countries over the period from 1998 to 2012, with typically around 10 to 30,000 usable observations, and we use control variables similar to those of Beck et al. (2013).

The results can be summarized as follows:

1. In the short run, we find that a change in competition has a consistently positive relationship to risk, regardless of whether we use the H-statistic or the Lerner index as a measure of competition.
2. In the long run, the results differ between the two competition measures, with the H-statistic showing a negative relationship of competition to risk, while the Lerner index correlates positively with risk.
3. Estimates made over periods before and after the GFC as well as over the full sample yield broadly consistent results.
4. Robustness checks with impaired loans as a dependent variable and with bank dummies, clustered standard errors, and bootstrap estimations underpin the main results of the analysis.
5. The differing long-run relationships between the risk measure, the logarithm of the Z-score, and the competition measures link to the fact that a rise in competition as indicated by the H-statistic is negatively related to the volatility of bank profits, while the Lerner index has a positive such link.

We conclude that the most important contribution of this study is the consistent short-run link from competition to risk, which has not been tested in the literature to date.

These results have important implications for policymaking, showing that considerable caution is warranted by regulators in the initial period after a rise in competition, since the indicators consistently show an accompanying rise in bank risk. In the longer term, our results suggest the need for careful consideration of the likely impact of a given policy shift, such as the separation of retail and wholesale banking or certain macroprudential policies. A direct impact on margins (as indicated by the Lerner index) is shown to be deleterious to risk, while enhancing competition (as indicated by the H-statistic) generally tends to enhance soundness.

An interesting topic for future research is the identification of factors underlying changes to competition. Such research could be feasible, for example, with US data, focusing on bank deregulation in the 1980s, or in emerging market economies where pro-competition reforms have been introduced. Studies in which such a shock affects competition but is exogenous to bank risk could shed light on the causal relationship between bank risk and competition. There could also be further testing of factors that could influence the trade-off between competition and risk in the short and long run, including potential variation by bank type, for Eastern European countries as opposed to Western countries, and depending on bank leverage and on excess capacity in banking.

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APPENDIX A

An Industrial Approach to Financial Instability

Davis (1995, 1999) suggests that an industrial analysis based on the effects of changes in entry conditions in financial markets can provide a supplementary set of underlying factors and transmission mechanisms to existing theories of financial instability. The easing of entry barriers can be due to deregulation, technical progress or market developments reducing the comparative advantage of incumbents over new entrants. Note that such a framework does not require an actual new entry; rather, the key is that the sunk (irrecoverable) costs of market entry should decline (Tirole, 1988). This could be reflected in the more competitive behavior of incumbents to protect themselves from the threat of entry. It is commonly observed that such changes in entry barriers do not merely entail reductions in profit and/or the smooth elimination of excess capacity. Rather, reductions in spreads and increases in quantities of credit go beyond the equilibrium level (i.e., the level at which lenders can make normal profits on their lending business, on average, over the cycle), leaving the institutions involved vulnerable to financial instability. Note that this pattern is, by nature, a short-run phenomenon and, in due course, normal behavior will reassert itself.

Drawing on theories of financial instability and applying the logic of market competition, we argue that the following mechanisms, among others, could play a role:

- To the extent that new entrants to financial markets can induce borrowers to switch away from established credit relationships or offer extra credit (by offering lower prices), information-based linkages will be weakened and existing information devalued. Conceptually, new lenders may be seen as 'cannibalizing' existing market information and structure, to the detriment of existing firms. Nonetheless, new lenders are still likely to lend on the basis of inadequate or asymmetric information.

- Uncertainty may be increased by new entry. Incumbents may be unable to predict accurately the responses of new entrants to changing conditions and their existing knowledge of market dynamics will be rendered less useful. Entrants, inexperienced in the market, will face even greater uncertainty. Unaware of the dynamics of supply and demand in the market, they may be prone to herd-like behavior, for example, all lending to the same type of client. When the market itself is new (or after liberalization, when interest rate controls that previously prevented lending to risky borrowers have been removed), all institutions will face uncertainty.

- Competition may cause firms to make inadequate provisions for uncertain events, such as financial crises, because firms that make adequate provisions are undercut by those disregarding such possibilities for reasons of ignorance or competitive advantage. New entrants may be particularly prone to such undercutting.

Sufficiently short time horizons may even make firms disregard systematic risks such as the economic cycle in their risk appraisals and, so, again, via the process of competition, help to reduce prudential standards for the whole market. Hit-and-run entry, as predicted by the theory of contestable markets, must, by nature, have a short time horizon during the initial stages. Thus, for both types of lenders, entry may lead to a lowering of credit standards.

- Competition for market share, as stressed by managerial theories of the firm – an approach frequently adopted by entrants or in new and developing markets – may lead to cumulative reductions in market prices until checked by the losses of participants and withdrawal or retrenchment. Such competition may persist if participants can cross-subsidize their operations from others making excess profits (i.e., there is market failure elsewhere) and are relatively immune to takeovers, as is the case for banks in most countries. The evaluation of loan officers over a short period on the basis of current lending performance is typical of market share-oriented banks.

- Besides the features outlined above, which are of particular importance in financial markets, several more general features of competitive processes may cause the competitive equilibrium to be overshot. Firms earning normal profits on their existing products may all be simultaneously attracted to situations offering potential for growth, but individual firms are unable to predict whether rivals will follow. Such tendencies will be particularly marked if there is no clear ordering of firms in terms of likelihood of success. Once investments are sunk, entry decisions may be difficult to reverse. Moreover, if there are sunk costs, firms may find it optimal to stay in the market for some time, even if they incur losses, as they will lose sunk costs of reputation, and so forth, if they leave. During this period, they may be vulnerable to adverse conditions in financial markets.

APPENDIX B

Variable Statistics

TABLE A1 Number of banks, by country

This table reports the countries included in the sample, the number of banks for each country, and the percentage of the total sample that they represent.

Country	Number of banks	Percent of total	Country	Number of banks	Percent of total	Country	Number of banks	Percent of total
Austria	361	6.01	Germany	2,571	42.79	Netherlands	74	1.23
Belgium	106	1.76	Greece	31	0.52	Poland	76	1.26
Bulgaria	31	0.52	Hungary	49	0.82	Portugal	49	0.82
Cyprus	27	0.45	Ireland	41	0.68	Romania	38	0.63
Czech Republic	39	0.65	Italy	902	15.01	Slovakia	28	0.47
Denmark	147	2.45	Latvia	32	0.53	Slovenia	33	0.55
Estonia	14	0.23	Lithuania	16	0.27	Spain	257	4.28
Finland	25	0.42	Luxembourg	148	2.46	Sweden	125	2.08
France	483	8.04	Malta	13	0.22	UK	292	4.86

TABLE A2 Variables for the H-statistic and the Lerner index

This table reports the statistical properties of the independent variables used in the regressions to determine competition, showing the variable definitions and their abbreviation, the number of observations, and the mean, standard deviation, minimum and maximum for each variable. Variables marked * are winsorized at 99%.

Variable definition/code	Observations	Mean	Standard deviation	Maximum	Minimum
R (Total interest revenue)*	46,148	270,522	1,041,025	8,541,040	744.9
IED (interest expense/debt)*	46,142	0.027	0.014	0.1	0.003
PTA (personnel expenses/assets)*	45,189	0.014	0.007	0.053	0.0004
OCF (other costs/fixed assets)*	45,224	2.2	5.18	38.0	0.18
CDT (deposits/short-term funding)*	43,563	0.78	0.21	1.0	0.02

(Continues)

TABLE A2 (Continued)

Variable definition/code	Observations	Mean	Standard deviation	Maximum	Minimum
LAR (loan/asset ratio)*	43,861	0.59	0.18	0.94	0.04
OTA (other non-earning assets/total assets)*	43,948	0.016	0.023	0.158	0.001
ETA (equity/total assets)*	43,971	0.079	0.052	0.356	0.016
IES (interest expense/total deposits plus money market funding)*	42,190	0.028	0.014	0.1	0.006
TOTALC (total cost)*	44,971	268,570	1,034,403	8,508,045	1,178

TABLE A3 Variables for risk/competition

This table reports the statistical properties of the variables used in the regressions to determine risk, showing the variable definitions and their code, the number of observations, and the mean, standard deviation, minimum and maximum for each variable. Variables marked * are winsorized at 99%, except for the H-statistic (95%).

Variable definition/code	Observations	Mean	Standard deviation	Maximum	Minimum
Log Z (Log Z-score)*	32,118	3.76	1.06	6.56	0.69
H (H-Statistic)*	86,774	−0.07	0.82	1.04	−1.9
Lerner (Lerner index)*	40,706	0.2	0.106	0.49	−0.15
CDT (deposits/short-term funding)*	43,563	0.78	0.21	1.0	0.02
LAR (loan/asset ratio)*	43,861	0.59	0.18	0.94	0.04
NIR (non-interest revenue/interest revenue)*	46,026	0.26	0.44	3.64	−0.06
LogTA (log total assets)*	46,842	13.6	1.8	19.3	9.8
PII (provisions/interest income)*	44,128	0.09	0.12	0.69	−0.26
SMT (stock market turnover)	83,664	106.9	50.7	259.6	0.14
CSI (capital stringency)	89,392	1.7	0.69	3	0
ACT (activity restrictions)	81,915	5.5	1.74	11	3
LOGER (German legal origin)	90,120	0.54	0.5	1	0
LOFR (French legal origin)	90,120	0.35	0.48	1	0
LOBRIT (UK legal origin)	90,120	0.06	0.23	1	0
LOSCAND (Scandinavian legal origin)	90,120	0.05	0.22	1	0
IMP (impaired loan ratio)*	10,202	0.067	0.072	0.43	0.0003