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The concentration–stability controversy in banking: New evidence from the EU-25

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ABSTRACT

This study explores whether the concentration–stability relation is affected by the level of analysis; i.e., bank-level versus country-level stability. The diverging results in the literature suggest that we may indeed expect differences between the two levels. With the z-score as the measure of financial stability, our theoretical analysis confirms that we may find such differences. Yet our empirical analysis for the EU-25 during the 1998–2014 period finds no economically significant effect of concentration on either the bank-level or the country-level z-score. The finding that concentration hardly affects stability at both levels of analysis is an indication of robustness in the empirical concentration–stability relation not previously established in the literature. This finding further suggests that neither supervisory restructuring, nor normal market-driven mergers, are likely to be substantially harmful to financial stability.

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1. Introduction

The recent financial crisis has witnessed massive government interventions in the financial sector, which have consisted mainly of the arrangement of restructuring mergers and bailouts. Although these measures have been successful in preventing an immediate collapse of the financial system during the recent crisis, they might have detrimental long-run consequences for the financial system. Bailout expectations have been shown to lead to additional risk taking by banks (Dam and Koetter, 2012). Moreover, while restructuring mergers are less likely to lead to the above-mentioned moral hazard problem, they reinforce a trend of an increasing degree of market concentration in the financial sector (Vives, 2011). Some theoretical studies indicate that this trend of consolidation is likely to make the financial system more fragile (Boyd and De Nicoló, 2005; Nier et al., 2007; De Nicoló and Lucchetta, 2009). Others, however, believe that a positive association exists between the degree

of banking market concentration and financial stability (Marcus, 1984; Keeley, 1990; Repullo, 2004).

Bank failures directly increase subsequent market concentration, whether handled by acquisition or liquidation. It is therefore crucial for regulators and policymakers to know the impact of banking market concentration on financial stability. The recent focus on systemic risk and macroprudential regulation indicates that regulators are not only concerned with the stability of individual banks, but also with the stability of a country's financial system as a whole. It is therefore important to explore whether the level of analysis – bank-level versus country-level stability – affects the observed concentration–stability relation.

Surprisingly, the role of the level of analysis has been ignored by the existing literature on the concentration–stability nexus. Existing studies typically focus on either bank-level or country-level stability, instead of analyzing the concentration–stability nexus at both levels. Furthermore, the literature yields conflicting evidence on the impact of concentration on stability (e.g., Boyd et al., 2010; Jiménez et al., 2013; Fiordelisi and Mare, 2014). Among the bank-level studies that use the z-score as the measure of financial stability, both negative and positive effects have been found. The

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only study using the country-level aggregate z-score as the measure of financial stability establishes a negative effect of concentration (Uhde and Heimeshoff, 2009).

The goal of our study is to explore the role of the level of analysis in the concentration–stability controversy, both theoretically and empirically, using the z-score at both the bank level and the country level as the measure of financial stability. Our study is, to our best knowledge, the first to systematically analyze the relation between banking concentration and financial stability at both levels in a coherent way and to make the comparison between both levels of analysis.

The theoretical part of our analysis shows that the aggregate z-score, unlike the bank-level z-score, incorporates the correlations across banks' returns on assets and thereby accounts for systemic risk. In this way, the bank-level and country-level z-scores measure different aspects of financial stability in the common scenario of imperfect return correlations. We may therefore expect empirical differences in the way concentration affects the z-score at both levels of analysis.

The empirical part of our analysis builds upon Uhde and Heimeshoff (2009) and focuses on commercial banks in the EU-25. We analyze the relation between banking market concentration and the z-score at both the bank level and the country level, using as much as possible the same model variables and econometric methodology for both approaches. In some of our models, we find a significantly negative effect of concentration on stability. Nevertheless, all estimated models indicate that the effect of concentration on stability is economically speaking small at both the bank level and the country level. Our finding that concentration hardly affects stability at both levels of analysis is an indication of robustness in the empirical concentration–stability relation not previously established in the literature.

Our findings are somewhat reassuring for regulators. They suggest that restructuring mergers, which are often arranged in order to restore financial stability during banking crises, will not contribute substantially to instability, nor will ordinary market-driven mergers and acquisitions.

The rest of the paper is organized as follows: in Section 2, the existing theoretical and empirical literature is reviewed. Our methodology is elaborated upon in Section 3 and Section 4. Section 5 presents the empirical results. Section 6 concludes.

2. Literature review

As mentioned in the introduction, the theoretical literature is inconclusive about the relation between banking market concentration and financial stability. Moreover, the empirical literature does not paint a clear picture either. We briefly review the literature below.

The concentration–stability controversy can be summarized as follows. Advocates of the so-called *concentration–stability view* argue that banks in more concentrated markets tend to be more stable for one of the following reasons. First, the *charter value hypothesis* maintains that a bank's charter is more valuable when the bank operates in a less competitive environment with high expected future profits. Banks in more concentrated markets will therefore engage less in excessively risky lending (Marcus, 1984; Chan et al., 1986; Keeley, 1990; Allen and Gale, 2000, 2004; Repullo, 2004) and will screen loan applicants better (Cordella and Yeyati, 2002; Hauswald and Marquez, 2006) to protect the value of their charter. Both outcomes are beneficial for financial stability.¹ Second, in more concentrated markets, banks become informed about

a larger proportion of borrowers. As a result, they make more informed decisions and are less exposed to credit risk (Marquez, 2002). Third, when the failure of a bank threatens the stability of the system, banks in more concentrated markets may find it easier to reach an agreement to rescue the troubled bank to prevent contagion. In more diffuse markets, an agreement is less likely to be reached because of a coordination problem. Hence, contagion is less likely to occur in more concentrated markets (Sáez and Shi, 2004). Finally, some argue that it is easier to monitor a system with only a few large banks than one with many small banks.²

Proponents of the *concentration–fragility view*, on the other hand, argue that banking market concentration is *detrimental* to financial stability. First, if the level of competition decreases with the degree of market concentration, banks in more concentrated markets can charge higher loan rates. This aggravates moral hazard problems on the part of borrowers, who will be induced to invest in more risky projects. As a result, the riskiness of the bank's asset portfolio increases (Boyd and De Nicoló, 2005; De Nicoló and Lucchetta, 2009). Second, banks in concentrated markets are more likely to be too-big-to-fail, which gives rise to a moral hazard problem on the part of bank managers (Mishkin, 1999).³ Third, the ex-ante risk of financial contagion is higher in more concentrated markets, since the probability that a particular bank is large enough to impact the rest of the system increases with the degree of market concentration (Nier et al., 2007). Finally, some argue that the supervision of concentrated banking markets is more difficult because banks in such markets tend to be larger and more complex than their counterparts operating in more diffuse markets (De Nicoló, 2004; Beck et al., 2006).

Empirical studies of the relation between banking market concentration and financial stability tend to focus solely on either the bank level or the country level. Analyses at the country level typically look at real episodes of financial crises. Using a crisis indicator variable, Beck et al. (2006) and Schaeck et al. (2009) find that higher levels of banking market concentration lower the probability of a financial crisis. Boyd et al. (2010), on the other hand, show that banking market concentration is positively associated with the probability of a sharp decline in lending, which is indicative of a crisis. Although a focus on real episodes of crises is intuitively appealing, it has the important drawback that an indicator variable does not provide information about the *intensity* of a crisis or about the fragility of the system in the *absence* of a crisis. For this reason, most bank-level analyses use the z-score as a proxy of the solvency of individual banks. As shown by the overview in Table 1, however, these studies tend to give conflicting evidence. Finally, Uhde and Heimeshoff (2009) combine aspects of the country-level and bank-level approaches mentioned above. They aggregate bank-level data to obtain an *aggregate z-score*, which can be interpreted as measuring the solvency of a country's financial sector as a whole. Looking at the EU-25 in the period between 1997 and 2005, Uhde and Heimeshoff (2009) obtain strong results which indicate a *negative* relation between banking market concentration and financial stability. The next sections will explore the role of the level of analysis in more detail, using the z-score as the measure of financial stability.

(Baumol, 1982; Corvoisier and Gropp, 2002) and the *efficiency theory* (Demsetz, 1973; Smirlock, 1985; Berger, 1995).

² Another often-mentioned argument in favor of the concentration–stability view is that banks in more concentrated markets are larger and therefore better able to diversify idiosyncratic risk. However, De Vries (2005) and Wagner (2010) show that diversification cannot raise the stability of the system as a whole, even though it may increase the stability of individual banks.

³ Dam and Koetter (2012) provide evidence that bank managers who expect to be bailed out in case of failure engage in more risky behavior.

¹ A crucial assumption behind this line of reasoning is that concentrated markets are less competitive. This view is challenged by the *contestability theory*

Table 1

Overview of empirical analyses of the effect of banking market concentration on financial stability.

Paper	Level	Dependent variable	Effect	Instrument(s)
Demirgüç-Kunt and Detragiache (2002)	Country	Crisis dummy	Positive	No
De Nicolis et al. (2004)	Bank	z-Score	Negative	No
Beck et al. (2006)	Country	Crisis dummy	Positive	Yes
De Nicolis and Loukoianova (2007)	Bank	z-Score	Negative	No
Berger et al. (2009)	Bank	z-Score	Positive	Yes
Boyd et al. (2009)	Bank	z-Score/loan losses	Positive	Yes
Schaek et al. (2009)	Country	Crisis dummy	Positive	No
Uhde and Heimeshoff (2009)	Country	Aggregate z-score	Negative	Yes
Boyd et al. (2010)	Country	“Crisis indicators”	Negative	Yes
Jiménez et al. (2013)	Bank	Non-performing loans	Non-linear	Yes
Fiordelisi and Mare (2014)	Bank	z-Score	Ambiguous	Yes
Fu et al. (2014)	Bank and country	z-Score/crisis dummy	Negative	Yes

The reported effects are significant at the 10% level at the least. The last column indicates whether the study controls for reverse causality by instrumenting the concentration measure with an exogenous variable.

3. Measuring concentration and financial stability

This section explains how we measure concentration and how we calculate the bank-level and country-level z-scores. We also make a comparison between the bank-level and country-level z-scores.

3.1. Measuring market concentration

In both the country-level and the bank-level analysis, we use the *five-bank concentration ratio* (CR₅) and the *Herfindahl–Hirschman index* (HHI) as measures of the degree of market concentration. The CR₅ is defined as the combined market share in terms of assets of the largest five banks operating in a country. Higher values thus indicate a more concentrated market. Although the CR₅ is a straightforward measure, its drawbacks are that the cut-off point of five banks is arbitrary and that it ignores the market shares of all other banks in the country. As a result, the CR₅ could be the same for markets with rather different structures (Bikker, 2004). The HHI does not suffer from an arbitrary cut-off point, but it has the drawback of being sensitive to the entrance of a large number of small banks (Rhoades, 1995). According to Bikker (2004), differences in the CR₅ across countries are mainly determined by the skewness of the size distribution of banks, whereas differences in the HHI result mainly from differences in the number of banks operating in the market. We therefore use both concentration measures. Note that both the CR₅ and the HHI are measured at the country level. As such, we assume that countries represent banking markets, as is common in the literature (e.g., Bikker et al., 2012).

3.2. The bank-level and country-level z-scores

We use the z-score as our measure of financial stability, which is based on bank balance sheet data (Roy, 1952; Hannan and Hanweck, 1988; Boyd et al., 1993). The z-score is a widely used solvency measure, which combines information about *profitability*, *capital buffers* and *return volatility*. The z-score of bank *i* at time *t* is defined as

$$z_{it}^B = \frac{E(r_{it}) + k_{it}}{\sigma(r_{it})}. \quad (1)$$

Here $E(r_{it}) = E\left(\frac{\pi_{it}}{a_{it}}\right)$ is the bank's expected return on assets at time *t*, calculated as expected net income in the coming period ($E(\pi_{it})$) divided by total assets (a_{it}). Furthermore, $k_{it} = \frac{e_{it}}{a_{it}}$ is the equity ratio at time *t*, calculated as total equity (e_{it}) divided by total assets.

Finally, $\sigma(r_{it})$ is the standard deviation of the return on assets.⁴ If returns are normally distributed, the bank's probability of default is equal to $1 - \Phi(z_{it}^B)$, where Φ is the standard normal cdf.

We now turn to a country's *aggregate z-score*. Without loss of generality, we assume that only two banks are active in the country. Similar to the individual z-score, the aggregate z-score of two banks, say *i* and *j*, is defined as

$$z_{ij,t}^C = \frac{E(r_{ij,t}) + k_{ij,t}}{\sigma(r_{ij,t})}, \quad (2)$$

where the subscript *ij* refers to aggregate values. Here

$$E(r_{ij,t}) = E\left(\frac{\pi_{it} + \pi_{jt}}{a_{it} + a_{jt}}\right) = \frac{a_{it}E(r_{it}) + a_{jt}E(r_{jt})}{a_{it} + a_{jt}} = w_{it}E(r_{it}) + w_{jt}E(r_{jt}) \quad (3)$$

where $w_{it} = \frac{a_{it}}{a_{it} + a_{jt}}$ and $w_{jt} = \frac{a_{jt}}{a_{it} + a_{jt}}$ are the asset weights of bank *i* and *j*, respectively, at time *t*. Using a similar exercise, we can write $k_{ij,t}$ as

$$k_{ij,t} = w_{it}k_{it} + w_{jt}k_{jt}. \quad (4)$$

The standard deviation $\sigma(r_{ij,t})$ equals

$$\begin{aligned} \sigma(r_{ij,t}) &= \sigma(w_{it}r_{it} + w_{jt}r_{jt}) \\ &= \sqrt{w_{it}^2\sigma^2(r_{it}) + w_{jt}^2\sigma^2(r_{jt}) + 2w_{it}w_{jt}\sigma(r_{it})\sigma(r_{jt})\rho(r_{it}, r_{jt})}, \end{aligned} \quad (5)$$

where $\rho(r_{it}, r_{jt})$ is the correlation between the banks' returns on assets. Substitution of (3)–(5) into (2) gives

$$z_{ij,t}^C = \frac{w_{it}(E(r_{it}) + k_{it}) + w_{jt}(E(r_{jt}) + k_{jt})}{\sqrt{w_{it}^2\sigma^2(r_{it}) + w_{jt}^2\sigma^2(r_{jt}) + 2w_{it}w_{jt}\sigma(r_{it})\sigma(r_{jt})\rho(r_{it}, r_{jt})}}. \quad (6)$$

The aggregate z-score of two banks can be interpreted as the z-score corresponding to a portfolio that consists of a weighted combination of the two individual banks. The aggregate z-score indicates the number of standard deviations by which the return of the portfolio could fall below its expected value before exhausting the portfolio's capital buffer.

⁴ For the empirical calculation of z-scores, we need estimates of expected returns and standard deviations. Throughout, we proxy these variables by realized returns in the last period and the standard deviation of the realized return over the sample period, respectively. This approach is common in the literature. Alternatively, some studies use a rolling window to obtain a time-varying estimate of the standard deviation of the return. Even though this approach is conceptually attractive, a rolling window has the drawback that it requires a relatively long sample to deliver reliable estimates. As such, Lepetit and Strobel (2013) find that using the entire sample period for calculating the standard deviation of returns gives better estimates of the return volatility.

We thus see that the aggregate z-score is generally *not* a weighted average of the banks' individual z-scores due to the presence of imperfect return correlation between banks. The return correlation is a measure of banks' interconnectedness. Because the aggregate z-score uses banks' return correlations, it reflects systemic risk, which has become a primary focus of prudential regulation in recent years. Indeed, the aggregate z-score goes to infinity when two equally sized banks with the same return standard deviation have a perfectly negative return correlation, even when the individual z-scores of the two banks are finite. More formally, we observe that $\sigma(r_{it}) > 0$, $\sigma(r_{jt}) > 0$ and $-1 \leq \rho(r_{it}, r_{jt}) \leq 1$. We must therefore have:

$$\begin{aligned} \frac{w_{it}(E(r_{it}) + k_i) + w_{jt}(E(r_{jt}) + k_j)}{w_{it}\sigma(r_{it}) + w_{jt}\sigma(r_{jt})} &\leq z_{ij,t}^C \\ &\leq \frac{w_{it}(E(r_{it}) + k_i) + w_{jt}(E(r_{jt}) + k_j)}{|w_{it}\sigma(r_{it}) - w_{jt}\sigma(r_{jt})|} \end{aligned} \quad (7)$$

Consequently, the aggregate z-score of two banks could be very high even though both banks are quite fragile, as long as their return correlation is low enough. It is readily seen that, if $\rho(r_{it}, r_{jt}) = 1$, the aggregate z-score in Eq. (6) can be written as $z_{ij,t}^C = v_{ij,t}z_{it}^B + (1 - v_{ij,t})z_{jt}^B$, where $v_{ij,t} = \frac{w_{it}\sigma(r_{it})}{w_{it}\sigma(r_{it}) + w_{jt}\sigma(r_{jt})} = \frac{a_{it}\sigma(r_{it})}{a_{it}\sigma(r_{it}) + a_{jt}\sigma(r_{jt})}$. Appendix A provides a complete proof of the next result.

Result 3.1. The aggregate z-score is a weighted average of banks' individual z-scores if and only if banks' returns on assets are perfectly correlated.

We can thus distinguish two cases. In the unlikely scenario that banks' returns are perfectly correlated, the aggregate z-score equals the asset-weighted average of the individual z-scores. Then the aggregate z-score uses exactly the same information as the individual bank-level z-scores. In the common scenario that banks' returns are imperfectly correlated, the country-level z-score is no longer equal to the weighted average of banks' individual z-scores. The aggregate z-score then additionally incorporates the return correlation across banks and thereby accounts for systemic risk. In this way, the bank-level and country-level z-scores measure different aspects of financial stability. We may therefore expect empirical differences in the way concentration affects stability at both levels. This will be explored in the next sections.

4. Econometric models

This section discusses the bank-level and country-level models that we will use in our empirical analysis to estimate the impact of concentration on financial stability as measured by the z-score.

4.1. Bank-level model

Our bank-level analysis will be based on the following model:

$$z_{it}^B = \beta_i + \beta \text{CONC}_{it}^* + \sum_k \beta_{x,k} x_{it,k} + \sum_\ell \beta_{y,\ell} y_{it,\ell} + \sum_t \beta_{time,t} d_t + \varepsilon_{it}, \quad (8)$$

where the subscripts i and t denote the bank and year, respectively. The variable z_{it}^B refers to the (logarithmically transformed) z-score of bank i in year t , while CONC_{it}^* represents one of the two (logarithmically transformed) measures of banking market concentration in the country where bank i is located. Moreover, $x_{it,k}$ refers to the k -th bank-specific control variable for bank i in year t and $y_{it,\ell}$ to the ℓ -th country-specific control variable for bank i in year t . In addition, the model includes a bank-specific individual effect β_i that will

be further specified later, year dummies d_t and a zero-mean error term ε_{it} . The main coefficient of interest is β , which measures the percentage change in the individual z-score following a percentage change in the degree of market concentration.

Similar to Uhde and Heimeshoff (2009), we control for the following macroeconomic variables: the rate of real GDP growth, the level of GDP per capita, the rate of inflation, and the real interest rate. In addition to the macroeconomic controls, five bank controls are included in the analysis. These are the bank's total assets, net interest margin, ratio of loan loss provisions (LLP) to total assets, cost-income ratio, and loan-assets ratio.

On the basis of the existing literature, we expect the following coefficient signs. GDP growth and GDP per capita are expected to have a positive effect on financial stability (Laeven and Majnoni, 2003), whereas the effects of the rate of inflation and the real interest rate are theoretically ambiguous (Uhde and Heimeshoff, 2009). We expect a negative coefficient for total assets (Boyd and Runkle, 1993; De Nicoló, 2000). The net interest margin is a measure of profitability, which is expected to have a positive effect on stability. The ratio of loan loss provisions to total assets, on the other hand, is a measure of credit risk and is expected to have a negative effect. The cost-income ratio is a measure of bank inefficiency and is expected to negatively affect financial stability (Uhde and Heimeshoff, 2009). Finally, the loan-asset ratio measures the extent to which banks are specialized in making loans as opposed to obtaining other sources of income. Its effect is not a-priori clear (Berger et al., 2009; Beck et al., 2013).

4.2. Country-level model

We specify the following country-level model:

$$z_{jt}^C = \gamma_j + \gamma \text{CONC}_{jt}^* + \sum_k \gamma_{x,k} x_{jt,k}^C + \sum_\ell \gamma_{y,\ell} y_{jt,\ell} + \sum_t \gamma_{time,t} d_t + \eta_{jt}, \quad (9)$$

where z_{jt}^C is the (logarithmically transformed) aggregate z-score of country j in year t and CONC_{jt}^* is the (logarithmically transformed) degree of market concentration in country j 's banking sector in year t . In addition, $x_{jt,k}^C$ refers to the k -th country-aggregated bank control variable and $y_{jt,\ell}$ to the ℓ -th macroeconomic control variable in country j at time t . The model includes a country-specific individual effect γ_j that will be specified later, year dummies d_t and a zero-mean error term η_{jt} . The coefficient of interest is γ , which measures the percentage change in the aggregate z-score following a percentage change in the degree of market concentration.

We include the same control variables as used in the bank-level analysis. Yet the country-level model includes country-aggregated values of total assets, net interest margin, loss provisions ratio, cost-income ratio, and loan-assets ratio. These bank controls are calculated as an asset-weighted average of the banks in a country, with the weight of bank i in year t equal to $\frac{a_{it}}{\sum_k a_{kt}}$. Table 2 lists the variable definitions and data sources for both the country-level and the bank-level model.

5. Empirical analysis

We use Bankscope EU-25 bank data to run both a bank-level and a country-level analysis. We assess the impact of banking market concentration on financial stability as measured by the z-score. This section discusses the data and the empirical results based on the bank-level and country-level models.

Table 2
Variable definitions and data sources.

Variable	Definition	Source
<i>Dependent variable</i>		
z-Score	$(\text{Capital ratio} + \text{return on assets}) / \text{SD}(\text{return on assets})$	Authors' calculations
Capital ratio	Ratio of equity to total assets	Bankscope
Return on assets	Return on average assets before taxes	Bankscope
SD (return on assets)	Standard deviation of return on assets over the sample period	Authors' calculations
<i>Explanatory variables</i>		
CR ₅	Combined market share in assets of largest 5 banks in the country	ECB statistics
HHI	Sum of squared market shares in assets of all banks in the country	ECB statistics
<i>Country controls</i>		
GDP growth	Rate of real GDP growth (%)	World Development Indicators
GDP per capita	Ratio of nominal GDP to population (US \$)	World Development Indicators
Inflation	Rate of growth of GDP deflator (%)	World Development Indicators
Real interest rate	3-Month money market interest rate minus GDP deflator (%)	Eurostat, OECD
<i>Bank controls</i>		
Assets	Total assets (US \$)	Bankscope
Net interest margin	Bank's net interest revenue as a share of interest-bearing assets (%)	Bankscope
LPP ratio	Loan loss provisions / total assets (%)	Bankscope
Cost-income ratio	Ratio of overhead costs to total revenue (%)	Bankscope
Loan-asset ratio	Ratio of outstanding loans to total assets (%)	Bankscope

Fig. 1. Aggregate z-scores of the banking sector in the five major European economies (1998–2014).

5.1. Data description

Our empirical analysis uses bank balance sheet data from the EU-25 in the period between 1998 and 2014, obtained from Bankscope. The bank-level sample includes commercial banks in the EU-25 for which Bankscope provides data for at least five different years. We drop observations with statements that are under processing by Fitch ratings, of branches with no statement, of no longer existing banks without statements, and of banks with no statement. We use consolidated data where possible (Bikker et al., 2012) and apply the Duprey–Lé algorithm to iteratively drop duplicates for any given bank while keeping the time series for each bank as long as possible (Duprey and Lé, 2015).

Fig. 1 shows the evolution over time of the aggregate z-scores of the five major European economies in the period from 1998 to 2014. The figure shows substantial variation in the z-score; some over time but especially between countries. The recent financial crisis resulted in a sharp decrease in the z-scores of France, Germany and the UK in 2008 and a quick subsequent recovery. In Italy and Spain, on the other hand, the effects of the crisis are reflected in a steadier decline in the z-score. Fig. 2 shows the degree of banking market concentration in the five major European economies, as measured by the CR₅ (panel a) and the HHI (panel b). The process of

Fig. 2. Five-bank concentration ratio (a) and Herfindahl–Hirschman index (b) of the banking sector in the five major European economies (1998–2014).

increasing market concentration in these countries, especially after the 2007–2008 financial crisis, can clearly be seen. The exception is France, where neither of the two concentration measures exhibit a tendency to increase over time.

5.2. Preliminary analysis

Descriptive statistics of all model variables are shown in Tables 3 and 4. We divide the sample into a pre-crisis and a (post-)crisis period to show the structural change due to the global and financial crisis that started with the fall of Lehman Brothers in September 2008. We thus distinguish between the full sample period (1998–2014), the pre-crisis period (1998–2007) and the (post-)crisis period (2008–2014). The pre-crisis and (post-)crisis sample differ in various ways. For instance, we observe a noticeable drop in GDP growth and real interest rates after the onset of the crisis and a substantial increase in banks' average total assets.⁵

Unit roots. We have applied the Im–Pesaran–Shin (IPS) panel unit root test (Im et al., 2003) to both the bank-level and country-level data sets. The distribution of the IPS test statistic is derived under fixed- $T/N \rightarrow \infty$ asymptotics, which may be problematic given our small- T /small- N country-level data set. We therefore report the results for the country-level IPS tests with some caution. At the bank-level though, fixed- $T/N \rightarrow \infty$ asymptotics seem reasonable. Only for (log) per-capita income, the IPS test cannot reject the null hypothesis of non-stationarity. Given the evidence for non-stationarity of the log per-capita income variable, we will exclude this variable from the analysis.

Return correlations. As explained in Section 3.2, we expect imperfect return correlation across banks. To get a feel for these correlations, we run an explorative analysis. For each bank i in country j in year t , we calculate the equally-weighted return on assets of the other banks in country j in year t (denoted r_{jt}^{-i}). We interpret this as a 'market return' (in a factor-model sense). Subsequently, we calculate, for each country j , the sample correlation between the market return and banks' individual return on assets as an estimate of $\rho_j = \text{Cor}(r_{it}, r_{jt}^{-i})$. The ρ_j s give a rough indication of the return correlation across banks in a country and are reported in Table 5. We have only done this for the full sample period to ensure a sufficient amount of observations for calculating the sample correlation. Table 5 shows that the ρ_j s are positive for all countries with the exception of Malta, the Netherlands and Portugal. The ρ_j s range between -0.62 (Malta) and 0.65 (Ireland) and its average value over all 25 countries equals 0.21 .

5.3. Bank-level model

Since the z-score is heavily skewed to the right, we use a logarithmic transformation of the z-score as the dependent variable.⁶ To avoid losing observations with negative z-scores, we add a constant equal to the sample median before taking the logarithm. The transformed z-score is thus calculated as $\tilde{z}_{it}^B = \log(z_{it}^B + z^m)$, where z^m is the sample median of the z-score. The effect of a unit change in the degree of banking market concentration on the z-score is therefore

$$\frac{\partial \tilde{z}_{it}^B}{\partial \text{CONC}_{it}^*} = \frac{dz_{it}^B}{dz_{it}^B} \frac{\partial z_{it}^B}{\partial \text{CONC}_{it}^*} = \beta(z_{it}^B + z^m). \quad (10)$$

⁵ We notice that the countries Cyprus and Malta are only available in the post-crisis period; we will later investigate the consequences of this in our robustness analysis. Furthermore, initially we also considered a credit growth variable taken from the World Development Indicator. However, this variable had relatively many missing values, thus reducing the sample. Because of the limited significance of this variable in our preliminary model estimations, we have excluded this variable altogether.

⁶ As shown by Lepetit and Strobel (2015), the logarithm of the z-score is negatively proportional to an upper bound of the log odds of insolvency.

Table 3
Sample statistics (bank level).

	1998–2014					1998–2007					2008–2014				
	Mean	SD	Min	Max		Mean	SD	Min	Max		Mean	SD	Min	Max	
z-score	21.68	21.50	−0.22	165.42		21.25	20.72	−0.02	165.42		22.12	22.28	−0.22	163.89	
CR ₅	0.47	0.17	0.19	0.99		0.45	0.19	0.19	0.99		0.48	0.16	0.23	0.95	
Herfindahl	0.07	0.06	0.01	0.41		0.07	0.06	0.01	0.41		0.07	0.05	0.02	0.37	
Per capita income	37,437	19,254.32	2976	116,613		32,292	16,356	2976	104,841		42,832	20,547	11,320	116,613	
GDP growth	1.65	2.94	−14.81	11.90		3.05	2.17	−1.13	11.90		0.18	2.92	−14.81	7.58	
Inflation rate	1.99	1.89	−9.69	20.13		2.35	2.14	−1.26	20.13		1.61	1.48	−9.69	11.84	
Real interest rate	0.77	2.18	−9.54	25.21		1.44	2.05	−9.54	15.34		0.07	2.09	−5.30	25.21	
Total assets	41,810,743	202,153,100	3270	3,065,092,958		25,748,503	141,413,993	3270	2,833,801,543		58,653,545	249,404,592	10,610	3,065,092,958	
Net interest margin	2.52	1.62	0.03	10.58		2.71	1.75	0.03	10.52		2.31	1.44	0.03	10.58	
LTP-asset ratio	0.51	0.87	−1.24	6.23		0.38	0.73	−1.21	6.23		0.66	0.97	−1.24	6.09	
Cost-income ratio	0.67	0.25	0.12	2.38		0.66	0.24	0.12	2.38		0.67	0.26	0.12	2.38	
Loan-asset ratio	0.53	0.25	0.01	0.98		0.50	0.26	0.01	0.98		0.55	0.25	0.01	0.98	
# obs.	9485					4855					4630				
# banks	1051					883					858				
# countries	25					23					25				
# years	17					10					7				

Table 4
Sample statistics (country level).

	1998–2014					1998–2007					2008–2014				
	Mean	SD	Min	Max		Mean	SD	Min	Max		Mean	SD	Min	Max	
z-score	14.37	9.82	0.29	79.82		14.30	8.91	2.57	79.82		14.49	10.98	0.29	53.91	
CR ₅	0.58	0.19	0.19	0.99		0.57	0.20	0.19	0.99		0.60	0.18	0.23	0.95	
Herfindahl	0.11	0.08	0.01	0.41		0.11	0.09	0.01	0.41		0.11	0.07	0.02	0.37	
Per capita income	30,095	19,849	2976	116,613		23,047	15,217	2976	88,680		36,232	21,145	11,248	113,727	
GDP growth	2.16	3.65	–14.81	11.90		3.76	2.55	–1.13	11.90		0.51	4.11	–14.81	11.09	
Inflation rate	2.54	2.61	–9.69	20.13		3.25	2.75	–1.26	19.96		2.07	2.69	–9.69	20.13	
Real interest rate	0.79	2.84	–9.54	25.21		1.41	2.73	–7.27	15.34		0.42	3.01	–9.54	25.21	
Total assets	234,493,859	415,695,703	32,564	1,890,742,062		103,726,527	226,944,769	114,297	1,262,991,515		378,712,055	510,288,454	3,035,837	1,766,824,936	
Net interest margin	2.42	1.27	0.48	6.55		2.80	1.48	0.28	9.52		2.07	1.06	0.54	5.43	
LIP-asset ratio	0.52	0.68	–0.46	3.51		0.37	0.67	–2.11	6.81		0.75	1.11	–0.71	7.82	
Cost-income ratio	0.63	0.18	0.37	2.38		0.67	0.29	0.35	3.75		0.60	0.18	0.38	1.86	
Loan-asset ratio	0.54	0.14	0.16	0.89		0.51	0.14	0.15	0.88		0.57	0.13	0.25	0.80	
# obs.	378					205					173				
# country	25					23					25				
# years	17					10					7				

Table 5

Sample correlations per country (1998–2014).

Country	Correlation
Austria	0.11
Belgium	0.07
Cyprus	0.29
Czech Republic	0.18
Denmark	0.49
Estonia	0.52
Finland	0.17
France	0.07
Germany	0.01
Greece	0.46
Hungary	0.37
Ireland	0.65
Italy	0.22
Latvia	0.39
Lithuania	0.44
Luxembourg	0.11
Malta	–0.62
The Netherlands	–0.03
Poland	0.09
Portugal	–0.04
Slovak Republic	0.15
Slovenia	0.52
Spain	0.09
Sweden	0.16
United Kingdom	0.26

Notes: For each bank in country *i* in year *t*, we have calculated the equally-weighted return on assets of the other banks in country *i* in year *t* ('market return'). Subsequently, we have calculated for each country the sample correlation between banks' individual return on assets and the market return. The resulting correlations are reported in the table.

For a bank with a z-score equal to the median, this gives:

$$\frac{\partial z_{it}^B / \partial \text{CONC}_{it}^*}{z_{it}^B} = 2\beta \quad (11)$$

Hence, multiplying the estimate of β in Eq. (8) by 2 gives the estimated percentage change in the z-score due to a percentage change in the degree of market concentration, for a bank with a z-score equal to the sample median.

The fixed-effects (FE) estimator for Eq. (8) is attractive, because it allows for unobserved bank-specific heterogeneity that is correlated with the observed covariates. However, similar to Uhde and Heimeshoff (2009) we find that the amount of time variation in the concentration measures is limited. We therefore incur the risk that part of the effect of concentration on stability is absorbed by the fixed effect. This could lead to a seemingly minor estimated effect of concentration on stability, while the true effect is much larger. We therefore also use the random-effects (RE) estimator. In contrast to the FE estimator, (nearly) time-invariant variables pose no problem for the RE estimator. However, the RE estimator assumes that any unobserved bank-specific heterogeneity is uncorrelated with the observed covariates, which may be unrealistic. We will later show that the point estimates obtained by the FE estimator are very similar to those generated by the RE estimator, suggesting that the limited amount of time variation in the concentration measures poses no problem for the FE estimator.

For the moment we assume that a sufficient way to deal with various sorts of endogeneity in Eq. (8) is to account for time-invariant and bank-invariant omitted variables correlated with the observed variables. We do this by using the FE estimator in combination with year dummies. We will verify the validity of this assumption in Section 5.5.

The RE and FE models are estimated for the full sample period 1998–2014 and contain either the CR₅ or the HHI as the concentration measure. The estimated coefficients are reported in Table 6. For

Table 6
Estimation results (RE and FE estimator, bank-level, 1998–2014).

	CR ₅								HHI							
	RE				FE				RE				FE			
	Est.	S.E.	t-Value	p-Value	Est.	S.E.	t-Value	p-Value	Est.	S.E.	t-Value	p-Value	Est.	S.E.	t-Value	p-Value
Intercept	5.2809	0.1612	32.7536	0.0000					5.1789	0.1619	31.9888	0.0000				
Concentration (log)	–0.1068	0.0294	–3.6368	0.0003	–0.1006	0.0335	–3.0011	0.0027	–0.0658	0.0165	–3.9885	0.0001	–0.0641	0.0195	–3.2929	0.0010
GDP growth	–0.0021	0.0012	–1.8026	0.0715	–0.0024	0.0012	–1.9483	0.0514	–0.0022	0.0012	–1.8170	0.0692	–0.0024	0.0012	–1.9646	0.0495
Inflation rate	0.0002	0.0020	0.0807	0.9357	0.0013	0.0021	0.6222	0.5338	0.0001	0.0020	0.0518	0.9587	0.0012	0.0021	0.5980	0.5498
Real interest rate	–0.0003	0.0018	–0.1468	0.8833	0.0003	0.0019	0.1470	0.8831	–0.0002	0.0018	–0.1230	0.9021	0.0003	0.0019	0.1677	0.8669
Total assets (log)	–0.1404	0.0109	–12.9319	0.0000	–0.1857	0.0153	–12.1529	0.0000	–0.1405	0.0108	–12.9548	0.0000	–0.1859	0.0153	–12.1832	0.0000
Net interest margin	0.0329	0.0047	7.0263	0.0000	0.0315	0.0050	6.2784	0.0000	0.0331	0.0047	7.0731	0.0000	0.0318	0.0050	6.3206	0.0000
LLP-asset ratio	–0.0333	0.0035	–9.4515	0.0000	–0.0304	0.0035	–8.6510	0.0000	–0.0331	0.0035	–9.3810	0.0000	–0.0302	0.0035	–8.5820	0.0000
Cost-income ratio	–0.1031	0.0163	–6.3104	0.0000	–0.1119	0.0170	–6.5838	0.0000	–0.1020	0.0163	–6.2479	0.0000	–0.1108	0.0170	–6.5371	0.0000
Loan-asset ratio	0.0427	0.0320	1.3338	0.1823	0.0362	0.0334	1.0831	0.2788	0.0421	0.0320	1.3137	0.1889	0.0354	0.0335	1.0588	0.2897
Adj. R ²	0.45				0.23				0.45				0.23			

Notes: This table reports the estimation results for the bank-level model. The standard-errors are robust for time-series correlation and heteroskedasticity. The time fixed-effects are not reported.

Table 7
Estimation results (RE and FE estimator, country level, 1998–2014).

	CR ₅								HHI							
	RE				FE				RE				FE			
	Est.	S.E.	t-Value	p-Value	Est.	S.E.	t-Value	p-Value	Est.	S.E.	t-Value	p-Value	Est.	S.E.	t-Value	p-Value
Intercept	4.3036	0.3384	12.7191	0.0000					3.7538	0.3439	10.9163	0.0000				
Concentration (log)	–0.4866	0.1117	–4.3547	0.0000	–0.3795	0.1661	–2.2845	0.0223	–0.1437	0.0155	–9.2680	0.0000	–0.2814	0.0794	–3.5463	0.0004
GDP growth	0.0134	0.0095	1.4055	0.1599	0.0145	0.0090	1.6140	0.1065	0.0134	0.0094	1.4170	0.1565	0.0144	0.0090	1.5967	0.1103
Inflation rate	–0.0088	0.0083	–1.0696	0.2848	–0.0062	0.0079	–0.7829	0.4337	–0.0087	0.0078	–1.1081	0.2678	–0.0060	0.0076	–0.7856	0.4321
Real interest rate	0.0110	0.0100	1.1007	0.2710	0.0129	0.0098	1.3083	0.1908	0.0111	0.0098	1.1338	0.2569	0.0130	0.0095	1.3703	0.1706
Total assets (log)	–0.1485	0.0164	–9.0648	0.0000	–0.1509	0.0166	–9.0950	0.0000	–0.3044	0.0467	–6.5181	0.0000	–0.1450	0.0153	–9.4642	0.0000
Net interest margin	0.1581	0.0212	7.4711	0.0000	0.1618	0.0201	8.0488	0.0000	0.1668	0.0204	8.1842	0.0000	0.1716	0.0188	9.1070	0.0000
LLP-asset ratio	–0.1744	0.0527	–3.3103	0.0009	–0.1680	0.0536	–3.1343	0.0017	–0.1710	0.0515	–3.3226	0.0009	–0.1637	0.0521	–3.1438	0.0017
Cost-income ratio	–0.1946	0.0846	–2.3005	0.0214	–0.1584	0.0868	–1.8256	0.0679	–0.1836	0.0840	–2.1869	0.0288	–0.1491	0.0847	–1.7592	0.0785
Loan-asset ratio	–0.4918	0.2861	–1.7190	0.0856	–0.3353	0.2878	–1.1650	0.2440	–0.5177	0.2723	–1.9015	0.0572	–0.3900	0.2672	–1.4600	0.1443
Adj. R ²	0.57				0.57				0.58				0.58			

Notes: This table reports the estimation results for the country-level model. The standard-errors are robust for time-series correlation and heteroskedasticity. The time fixed-effects are not reported.

Table 8

Signs of covariates' coefficients in the bank-level and country-level models.

	Expected sign	CR ₅				HHI			
		RE		FE		RE		FE	
		Bank	Country	Bank	Country	Bank	Country	Bank	Country
Concentration (log)	?	–	–	–	–	–	–	–	–
GDP growth	+	–	NS	–	NS	–	NS	–	NS
Inflation rate	?	NS	NS	NS	NS	NS	NS	NS	NS
Real interest rate	?	NS	NS	NS	NS	NS	NS	NS	NS
Total assets (log)	–	–	–	–	–	–	–	–	–
Net interest margin	+	+	+	+	+	+	+	+	+
LLP-asset ratio	–	–	–	–	–	–	–	–	–
Cost-income ratio	–	–	–	–	–	–	–	–	–
Loan-asset ratio	?	NS	–	NS	NS	NS	–	NS	NS

Notes: This table summarizes the signs of the covariates in the bank-level and country level models. 'NS' stands for 'not significant' at the 10% level. The column 'expected effects' indicates what the expected sign is according to the literature, where a question mark indicates a lack of consensus in the literature about the expected effect.

all models, the reported standard errors are robust to time series correlation and heteroskedasticity.

The negative point estimates of the FE and RE estimators are of a similar magnitude, suggesting that the limited amount of time variation is not a problem for the FE estimator. Because the RE models are rejected by a Hausman test, we proceed with the FE models, whose adjusted R^2 s equal 0.23. The FE models indicate that the effect of concentration on stability is *negative* and significant at the 1% level. In terms of the CR₅, a 1% increase in concentration leads to a *decrease* in stability of about $2 \times 0.1006 = 0.20\%$ with associated 95% confidence interval [0.07, 0.33]%, while for the HHI a 1% increase in concentration leads to a *decrease* in stability of about $2 \times 0.0641 = 0.13\%$ with corresponding 95% confidence interval [0.05, 0.20]%. We thus see that the economic significance of the effect of concentration on stability is limited, despite the statistical significance.⁷ With the exception of GDP growth, the signs of the control variables' coefficients are in line with our theoretical predictions whenever they are significant; see Table 8.

The estimated models rely on the standard assumption in the literature that banking markets coincide with individual countries; see e.g. Bikker et al. (2012). In this way, the focus of our models is on the relation between bank stability and *national* concentration measures, which – from a policy perspective – seems the most relevant focus. For some commercial banks in our country the standard assumption about the extent of the market may be incorrect, such as banks with a regional focus or banks that are active in multiple countries. For the first type of banks, the impact of a national concentration measure will probably be weaker than that of a regional concentration measure. Regarding the latter type of bank, we notice that the individual effects in our models will correct for time-invariant missing information about the extent of the market.

5.4. Country-level model

We estimate the country-level model of Eq. (9) with the logarithmically-transformed z-score as the dependent variable. Using similar arguments as for the bank-level analysis, we apply FE and RE estimators. We estimate the country-level models for the full sample period 1998–2014 using either the CR₅ or the HHI as the concentration measure. By definition, the country-level models focus on the relation between bank stability and *national* concentration measures. The estimated coefficients are reported in Table 7.

⁷ Statistical significance is a necessary but not a sufficient condition for economic significance. For more details about statistical versus economic significance, see e.g. Granger (1998).

As before, the reported standard errors are robust to time series correlation and heteroskedasticity.⁸

Again the RE and FE models provide negative point estimates of a similar magnitude, so there is no need for concerns about the FE estimator. Because the RE models are rejected by a Hausman test, we proceed with the FE models, whose adjusted R^2 s equal 0.57–0.58. The effect of concentration on stability is *negative* and significant at the 5% level. In terms of the CR₅, a 1% increase in concentration leads to a *decrease* in stability of 0.38%, with associated 95% confidence interval [0.05, 0.71]%. In terms of the HHI, a 1% increase in concentration leads to a *decrease* in stability of 0.28%, with associated 95% confidence interval [0.13, 0.44]%. The economic relevance of the effect of concentration on stability is larger than in the bank-level model, but economically speaking still modest. Throughout, the signs of the control variables' coefficients are in line with theory whenever they are significant; see Table 8.

5.5. Robustness checks

The control variables in both the bank-level and country-level models include year dummies. In each model, the number of dummies equals the number of years in the sample minus one; so $17 - 1 = 16$ in total. With only 378 bank-years, the country models may be overfit due to the large amount of year dummies (Babak, 2004), thus capturing noise rather than a meaningful economic effect. We therefore re-estimate the FE country-level models including a single time dummy for the (post-)crisis period (2008–2014); see Table 9. The estimated coefficients of concentration are all negative as before (–0.0312 for the CR₅ and –0.1317 for the HHI), but no longer statistically significant (p -values of 0.87 and 0.19, respectively). The significance of some of the control variables also changes, but the signs are still in line with theory.

We verify the robustness of the estimates to the exclusion of currently inactive banks. Second, we analyze separately the subsample of the largest 25% of banks in terms of total assets. Third, we redo the main analysis without the countries Cyprus and Malta, since they are only available during the post-crisis period. In all cases, we find similar estimation results as before.⁹

We also estimate the effect of concentration on stability separately for the pre-crisis period (1998–2007) and the (post-)crisis

⁸ Because the number of countries is relatively small in the country-level analysis, the formula-based clustered standard errors may be problematic (since they are based on the assumption that the number of countries is large). We therefore report the most conservative standard errors, where we choose between the standard errors based on a wild panel bootstrap (Cameron et al., 2008) and the formula-based standard errors.

⁹ More detailed estimation results for the robustness checks are available upon request.

Table 9

Estimation results with fewer time dummies (FE estimator, country level, 1998–2014).

	CR ₅				HHI			
	Est.	S.E.	t-Value	p-Value	Est.	S.E.	t-Value	p-Value
Concentration (log)	−0.0312	0.1973	−0.1583	0.8742	−0.1317	0.1001	−1.3160	0.1882
GDP growth	0.0060	0.0049	1.2351	0.2168	0.0055	0.0050	1.1044	0.2694
Inflation rate	−0.0436	0.0080	−5.4400	0.0000	−0.0443	0.0085	−5.2386	0.0000
Real interest rate	−0.0298	0.0082	−3.6466	0.0003	−0.0308	0.0083	−3.7148	0.0002
Total assets (log)	−0.0896	0.0178	−5.0475	0.0000	−0.0836	0.0163	−5.1353	0.0000
Net interest margin	0.1460	0.0330	4.4232	0.0000	0.1539	0.0317	4.8533	0.0000
LLP-asset ratio	−0.1486	0.0577	−2.5749	0.0100	−0.1447	0.0557	−2.5960	0.0094
Cost-income ratio	−0.1488	0.1294	−1.1500	0.2502	−0.1423	0.1261	−1.1290	0.2589
Loan-asset ratio	−0.0197	0.2932	−0.0673	0.9464	−0.1043	0.2785	−0.3746	0.7080
Adj. R ²	0.44				0.44			

Notes: This table reports the estimation results for the country-level model, with a single time dummy to avoid overfitting. The standard-errors are robust for time-series correlation and heteroskedasticity. The time fixed-effect is not reported.

Table 10

Robustness check: the effect of the crisis on the concentration–stability relation.

	CR ₅				HHI			
	Est.	S.E.	t-Value	p-Value	Est.	S.E.	t-Value	p-Value
<i>Bank-level</i>								
Concentration	−0.0968	0.0316	−3.0632	0.0022	−0.0607	0.0180	−3.3724	0.0007
Concentration $\times I(t \geq 2008)$	0.0143	0.0197	0.7247	0.4686	0.0098	0.0109	0.8974	0.3695
<i>Country-level</i>								
Concentration	−0.3762	0.1621	−2.3206	0.0203	−0.2813	0.0790	−3.5603	0.0004
Concentration $\times I(t \geq 2008)$	0.0279	0.1277	0.2187	0.8269	0.0010	0.0712	0.0134	0.9893

Notes: This table reports the estimated coefficient of the (log) concentration variable in the FE difference-in-difference models. The standard-errors are robust for time-series correlation and heteroskedasticity.

Table 11

Robustness check: results for FE-2SLS estimation.

	CR ₅				HHI			
	Est.	S.E.	t-Value	p-Value	Est.	S.E.	t-Value	p-Value
<i>Bank-level</i>								
Concentration	0.0811	0.2561	0.3164	0.7517	0.0906	0.2900	0.3126	0.7546
First-stage F-statistic	129.8				94.9			
<i>Country-level</i>								
Concentration	−0.6632	1.0660	−0.6221	0.5339	−0.3000	0.4652	−0.6450	0.5190
First-stage F-statistic	12.5				13.8			

Notes: This table reports the estimated coefficient of the (log) concentration variable in the FE models, estimated by means of 2SLS. The standard-errors are robust for time-series correlation and heteroskedasticity. The first-stage F-statistics are also reported.

period (2008–2014) using a difference-in-difference approach; see Table 10. The estimation results show that the effect of concentration on stability does not significantly differ between the two periods. We also draw this conclusion in the country-level model with a single time dummy, in which case the effect of concentration is not significant either.

Simultaneity of stability and concentration is a potential issue. This would arise if a drop in the stability of a (too-big-to-fail) bank results in a restructuring merger and an increase in the CR₅ or HHI. Such simultaneity would result in omitted, unobserved variables correlated with the concentration measure in Eqs. (8) and (9). In our previous estimations, we have assumed that the FE estimator, in combination with year dummies, is able to deal with this sort of endogeneity. We investigate the validity of this assumption by means of a FE-2SLS approach; see Table 11. At the bank level, we use log population as an instrument in the preferred specification, the FE model. The underlying motivation is that small countries are expected to be more concentrated than larger ones and that the instability of individual banks will not affect a country's population. At the country level, we use investment freedom from the Heritage Foundation as the instrument. The intuition is that this freedom indicator is an exogenous feature of the (financial) economy, which

determines its development and thereby the degree of banking concentration. At both the bank level and the country level the first-stage F-statistic is above the value 10. Hence, the instruments are strong enough (Stock et al., 2002). At the bank level, the coefficients of concentration equal 0.08 (CR₅) and 0.09 (HHI), but they are not statistically significant. At the country level, we find coefficients of −0.66 (CR₅) and −0.30 (HHI). Also here the coefficients of concentration are no longer statistically significant. Throughout, the magnitude and significance of the control variables is similar as before. The loss of statistical significance of the concentration coefficient in the FE-2SLS models reinforces our prior conclusion that the effect of concentration on stability is economically speaking limited.

In sum, our robustness checks confirm that the effect of concentration on stability is economically speaking limited, although sometimes statistically significant.

6. Conclusion

The recent focus on systemic risk and macroprudential regulation indicates that regulators are not only concerned with the stability of individual banks, but also with the stability of a coun-

try's financial system as a whole. It is therefore important to explore whether the level of analysis – bank-level versus country-level stability – affects the observed concentration–stability relation. The diverging results in the literature suggest that we may indeed expect differences between the two levels.

Our theoretical analysis has shown that the country-level aggregate z-score, unlike the bank-level z-score, incorporates the correlations across banks' returns on assets and thereby accounts for systemic risk. In the common scenario of imperfect return correlations, the bank-level and country-level z-scores measure different aspects of financial stability. We may therefore expect differences in the empirical concentration–stability relation between both levels of analysis.

The empirical part of our study has used bank data from the EU-25 during the 1998–2014 period to investigate the causal relation between banking market concentration and the z-score at both the bank level and the country level. At both levels of analysis, we find that the effect of concentration on stability is economically speaking limited. Our finding that concentration hardly affects stability at both levels of analysis is an indication of robustness in the empirical concentration–stability relation not previously established in the literature.

Our findings are somewhat reassuring for regulators. They suggest that restructuring mergers, which are often arranged in order to restore financial stability during banking crises, will not substantially contribute to instability, nor will ordinary market-driven mergers and acquisitions. Since existing research has indicated that bailouts lead to increased bank risk taking by raising expectations of future bailouts, our findings support the idea that restructuring mergers are a viable alternative to bailouts when a troubled bank is deemed too important to fail.

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Appendix A. Proof of Result 1

Result 1. The aggregate z-score is a weighted average of banks' individual z-scores if and only if banks' returns on assets are perfectly correlated.

Proof. Without loss of generality, we confine the proof to the case of two banks. If $\rho(r_{it}, r_{jt}) = 1$, the aggregate z-score in Eq. (6) can be written as $z_{ij,t}^C = v_{ij,t} z_{it}^B + (1 - v_{ij,t}) z_{jt}^B$, where $v_{ij,t} = \frac{w_{it}\sigma(r_{it})}{w_{it}\sigma(r_{it}) + w_{jt}\sigma(r_{jt})} = \frac{a_{it}\sigma(r_{it})}{a_{it}\sigma(r_{it}) + a_{jt}\sigma(r_{jt})}$. Conversely, assume that $z_{ij,t}^C = \tilde{v}_{ij,t} z_{it}^B + (1 - \tilde{v}_{ij,t}) z_{jt}^B$. From Eq. (6) it becomes clear that we must have

$$\tilde{v}_{ij,t} = \frac{w_{it}\sigma(r_{it})}{\sqrt{w_{it}^2\sigma^2(r_{it}) + w_{jt}^2\sigma^2(r_{jt}) + 2w_{it}w_{jt}\sigma(r_{it})\sigma(r_{jt})\rho(r_{it}, r_{jt})}}. \quad (\text{A.1})$$

However, because the weights must sum to unity, we have to get rid of the square root in the denominator of (A.1). This is only possible for $\rho(r_{it}, r_{jt}) = 1$, yielding $\tilde{v}_{ij,t} = v_{ij,t}$. □

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