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BANK-LEVEL ESTIMATES OF MARKET POWER

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ABSTRACT

The aim of this study is to provide an empirical methodology for the estimation of market power of individual banks. The new method employs the well-known model of Panzar and Rosse (1987) and proposes its estimation using the local regression technique. Thus, a number of restrictive assumptions regarding the properties of the production function of banks are relaxed, while the method proves successful in providing reasonable estimates of bank-level market power when applied to a large panel of banks of transition countries. The empirical results suggest that many banks in the sample deviate significantly from competitive practices and that market power varies substantially across banks in each country. Country averages of the bank-level results exhibit a very close relationship with standard, industry-level Panzar-Rosse estimates.

Keywords: Market power, bank-level, local regression
JEL classification: G21, L11, C14

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1. Introduction

Evaluating competition at the industry level is a standard preoccupation of industrial organization in general and a central concern of banking research in particular. Accordingly, several studies have assessed the level of competition in banking markets at different levels of aggregation (for a recent review of this literature, see Delis et al., 2008). In many circumstances, however, the researcher may be interested in obtaining bank-level measures of market power, so as to address questions regarding the potential relationship of market power with certain elements of the behavior of banks, the structure of the industry and the macroeconomic environment. Most of these studies either employ concentration as a measure of competition or estimate industry-level measures of competition. Only few studies construct Lerner indices (e.g. Jimenez et al., 2007) or Tobin's q (e.g. Keeley, 1990) to measure the market power of individual banks. Yet, use of Lerner indices entails the rather restrictive assumption of a constant marginal cost for the banking industry or for classes of banks, which is then used to calculate bank-specific price-cost margins. Furthermore, estimation of the marginal cost requires further assumptions regarding the functional form of the underlying production relationship. Tobin's q, even though quite useful as a proxy for market power, requires information on the market value of assets, which may not be generally available, and additionally it does not originate from standard industrial organization theory.

In an effort to make some progress with the estimation of bank-level market power using widely available sources of bank data, this paper combines two well-established theoretical and econometric frameworks. In particular, it utilizes the well-known model of Panzar and Rosse (PR hereafter) (1987) and proposes its estimation with the local regression (LR) method, as put forth by Cleveland and Devlin (1988). LR has been widely used in econometrics, and in the context of the present analysis it has the great advantage of producing observation-specific coefficients, thus dispelling concerns about the degrees of freedom. In addition, the restrictive assumption of a global parametric functional form (such as the Cobb-Douglas, translog or Fourier) that is needed to estimate marginal cost is avoided and, hence, the model is robust to such potential misspecification. Finally, since the parameters are localized at each observation, flexibility of the functional form is not an issue and the use of a general linear form gives a clear economic meaning to each and every coefficient that is made bank-specific through localization (see Kumbhakar et al., 2007).

This technique may be particularly valuable in exploring theoretical relationships in banking that require data on the market power of individual institutions. For example, studies exploring the relationship between competition and (i) the risk-taking behavior of banks (e.g. Boyd et al., 2006), (ii) regulation (e.g. Brissimis et al., 2008), (iii) the interest rate margins (e.g. Maudos and de Guevara, 2004), (iv) privatizations or M&As (e.g. Gugler and Siebert, 2007), (v) other industries' structure (e.g. Cetorelli and Strahan, 2006) and/or (vi) financial crises (e.g. Boyd et al., 2004) may benefit from the suggested approach. This is mainly because the number of observations will be considerably increased and, therefore, single-country studies are possible, while a number of concerns regarding distributional assumptions of the underlying production relationships are addressed. It is noteworthy that similar local regression techniques with more or less the same advantages over their parametric equivalents have been recently employed to measure bank efficiency (see Kumbhakar et al., 2007).

We opt for an application of the new method to a large panel of banks operating in 20 transition countries over the period 1999-2006. This choice is motivated by the existence of a recent body of literature that offers good priors regarding the competitive conditions in the banking sectors of these countries (see e.g. Delis, 2008; Brissimis et al., 2008; Yildirim and Philippatos, 2007), as well as by the rapid institutional changes that characterized them and may yield differences in market power at an inter- and/or intra-country level.

The rest of the paper is organized along the following lines. Section 2 comments on the theoretical background and presents the estimation method. Section 3 discusses the estimation results, and Section 4 offers some concluding remarks.

2. Background and empirical methodology

Recent studies that evaluate competitive conditions in the banking industry employ the so-called non-structural approaches, which have emerged under the impulse of the New Empirical Industrial Organization (NEIO) literature. These approaches, pioneered by Iwata (1974) and strongly enhanced by Appelbaum (1982), Bresnahan (1982, 1989), Lau (1982), Panzar and Rosse (1987) and Roeger (1995), test for the presence of market power by analyzing deviations from competitive pricing (marginal cost pricing). Their major advantage is formal grounding on explicit optimization models and equilibrium conditions and, as in every empirical framework, each of them has its advantages and disadvantages.[1]

The Panzar and Rosse (PR) model[2] relies on the premise that banks will employ different pricing strategies in response to a change in input costs. In other words, market power is measured by the extent to which changes in factor prices are reflected in revenue. PR define a measure of competition, the 'H-statistic', which represents the percentage change of the equilibrium revenue resulting from an infinitesimal percent increase in the price of all factors used by the firm, i.e. the sum of the input price elasticities. Owing to its relative simplicity, this methodology has been extensively applied to the banking sector, both in regional and single-country studies (see e.g. Molyneux et al., 1994; De Bandt and Davis, 2000; Bikker and Haaf, 2002; Claessens and Laveven, 2004, 2005).

Panzar and Rosse assert that the H-statistic is negative when the market structure is a monopoly, a perfectly colluding oligopoly, or a conjectural variations short-run oligopoly; an increase in input prices will increase marginal costs, reduce equilibrium output, and subsequently reduce revenue.[3] Under perfect competition, where banks' products are regarded as perfect substitutes, the Chamberlinian model, based on free entry of banks and determining not only the output level but also the equilibrium number of banks, produces the perfectly competitive solution as demand elasticity approaches infinity. In this case, the H-statistic is equal to unity. Shaffer (1982) shows that the H-statistic is also unity for a natural monopoly operating in a perfectly contestable market and for a sales-maximizing firm that is subject to break-even constraints. Consequently, an increase in input prices raises both marginal and average costs without altering the optimal output of a bank. Exit from the market will

[1] For a review of the advantages and disadvantages of the approaches followed in these studies, see Shaffer (2004a).
[2] Note that this paper has its origins in Rosse and Panzar's (1977) work.
[3] In the case where the monopolist faces a demand curve of constant price elasticity $e>1$ and where a constant returns to scale Cobb–Douglas technology is employed, PR proved that the H-statistic is equal to e-1. Hence, apart from the sign, the magnitude of the H-statistic may also be of importance, as the H-statistic yields an estimate of the Lerner index of monopoly power $L = (e-1)/e = H/(H+1)$ (Bikker and Haaf, 2002; Shaffer, 1983). It should be noted, however, that Shaffer (2004b), among others, is somewhat skeptical about how robustly the H-statistic maps into a range of oligopoly solutions.

evenly increase the demand faced by each of the remaining banks, thereby leading to an increase in prices and total revenue by the same amount as the rise in costs (i.e. demand is perfectly elastic). Finally, if the *H*-statistic is between zero (inclusive) and unity (exclusive), the market structure is characterized by monopolistic competition. Under monopolistic competition, potential entry leads to contestable market equilibrium, and income increases less than proportionally to input prices, as the demand for banking products that individual banks face is inelastic.

The PR model is a valuable tool in assessing market conditions, mainly owing to its simplicity and transparency, while it does not lack efficiency. Moreover, data availability becomes much less of a constraint, since revenue is more likely to be observable compared to output prices (needed by the other NEIO methods). Also, by utilizing bank-level data, this approach allows for bank-specific differences in the production function. In addition, the non-necessity to define the location of the market *a priori* implies that the potential bias caused by the misspecification of market boundaries is avoided; hence for a bank that operates in more than one market, the H-statistic will reflect the average of the bank's conduct in each market.

The H-statistic is derived using the following specification of the reduced-form revenue equation for a panel dataset:

$$\ln rtr_{it} = a_0 + a_1 \ln w_{1,it} + a_2 \ln w_{2,it} + a_3 \ln w_{3,it} + a_4 \ln b_{it} + a_5 \ln m_t + \varepsilon_{it} \qquad (1)$$

where *it* is the subscript indicating bank *i* at time *t*, *rtr* is a bank's real total revenue, w_1, w_2 and w_3 are the three input prices, *b* stands for other bank-specific characteristics and *m* stands for a number of country-specific control variables, the latter being included to capture differences of bank revenue owing to structural characteristics of the banking industry or different macroeconomic conditions in cross-country studies.

Estimation of Eq. (1) using conventional econometric techniques and bank-level data has been carried out in a number of studies (for a recent review, see Delis et al., 2008). The results obtained describe the competitive conditions that characterize the banking industry (or industries) examined. What remains as a challenge is to provide estimates of market power at the individual bank level. To conduct such an analysis, this paper draws on a non-parametric estimation technique, the local regression (LR) method, as put forth by Cleveland and Devlin (1988).[4] LR estimation is a consistent way to allow for nonparametric effects within the parametric model, and this is accomplished as follows. The underlying model for local regression is $Y_i = \mu(x_i) + \varepsilon_i$, where *x* are predictor variables and *Y* is the response variable. The unknown function *μ*(*x*) is assumed to be smooth and is estimated by fitting a polynomial model (a quadratic in our case, as in most of the literature) within a sliding window. Differently phrased, no strong assumptions are made about *μ* globally, but locally around *x* we assume that *μ* can be well approximated. For a fitting point *x*, define a bandwidth *h* that controls the smoothness of the fit and a smoothing window (*x-h(x)*, *x+h(x)*). To estimate *μ*, only observations within this sliding window are used. Therefore, for each fitting point a locally weighted least squares criterion of the following form is considered:

$$\sum_{i=1}^{n} W\left(\frac{x_i - x}{h}\right)\left(Y_i - (a_0 + a_1(x_i - x))\right)^2 \qquad (2)$$

where *W* is the weight function that assigns largest weights to observations close to *x*, and takes the form

$$W(u) = \begin{cases} (1-|u|^3)^3 & \text{if } |u| < 1 \\ 0 & \text{otherwise} \end{cases} \qquad (3)$$

The local least squares criterion of Eq. (2) is minimized to produce estimates $\hat{a}_0$ and $\hat{a}_1$.[5, 6] In terms of the PR model given by Eq. (1), the above discussion implies that we can obtain *n* estimates of each of the coefficients, naturally corresponding to each of the banks in the sample. Then, the H-statistic is calculated for each bank *i* from the equation $H_i = a_{1i} + a_{2i} + a_{3i}$.

Estimation of Eq. (1) using the aforementioned technique presents some considerable advantages, besides the obvious one of deriving observation-specific estimates through localization. First, no assumption regarding the functional form of

[4] For a thorough discussion of local regression, see Loader (1999).

[5] This discussion relates to the bivariate local regression. The multivariate local regression simply adds further terms to the right hand-side of the formula for *Y*. Estimations are carried out using the program Locfit.

[6] An important issue in the implementation of LR is the choice of an optimal bandwidth. Many alternatives have been proposed, like plug-in methods and cross-validation (see Kumbhakar et al., 2007). Here we used the generalized cross-validation method (see Loader, 1999), which in our case yields a bandwidth equal to 0.612.

the underlying production relationship is needed, and it is well-known that it is quite difficult for the researcher to be certain that the “correct” functional form has been chosen. Second, and given this qualification, economic hypotheses are not rejected simply because an “improper” functional form has been chosen. Third, localization implies that, besides obtaining bank-level H-statistics, bank-level elasticities of revenue with respect to input prices and structural and macroeconomic conditions are also obtained, which may be quite useful information for managers and policy-makers. For the above and possibly other reasons, a very recent literature has employed similar non-parametric techniques to measure bank efficiency (see e.g. Kumbhakar et al., 2007).

It should be noted, however, that estimation of the PR model using a non-parametric technique may also have some drawbacks. First, it is well-known that non-parametric techniques have to be applied to larger datasets to avoid the so-called “curse of dimensionality”. Luckily, this is not an issue for micro-level studies, where datasets are quite large. Second, the PR model should be estimated on observations that are in long-run equilibrium. To test for equilibrium, one can calculate another H-statistic (H_n) using the rate of return (return on assets), instead of total revenue, as the dependent variable in the regression equation.[7] In this framework, H_n=0 indicates that banking systems are in equilibrium. In the present analysis, testing for long-run equilibrium for every single observation separately is not feasible because one need to identify whether all observations one by one are in equilibrium. However, prior to estimation of the model using LR, long-run equilibrium has been examined at the industry level using exactly the same methodology as in previous studies (e.g. Claessens and Laeven, 2004). The results (not reported but available on request) suggest that the hypothesis of equilibrium is confirmed for all but one of the banking systems examined.[8]

[7] The empirical test for equilibrium is justified on the grounds that competitive capital markets will equalize the risk-adjusted rate of return across banks, so that (in equilibrium) the rate of return should not be statistically correlated with input prices.

[8] This is the Estonian banking system (p-value of the test for the Hn=0 is equal to 0.000) and thus the results for the H-statistic for Estonian banks may have to be treated with caution. Exclusion of Estonia from the LR regressions does not affect the results for the rest of the banks.

3. Estimation results

As discussed above, the dataset of the present study consists of bank- and country-level variables. All bank-level data are obtained from BankScope and include 2768 observations from 465 commercial banks (unbalanced panel) operating in the 20 transition countries reported in Table 1 for the period 1999-2006.[9] Some banks were excluded from the empirical analysis on account of their unreasonably high/low input price data or because some of the required data were missing. Following standard practice in banking industry studies, input prices are calculated by dividing interest expenses by total deposits (w_1), depreciation and other capital expenses by fixed assets (w_2) and personnel expenses by total assets (w_3).[10] The bank-specific variables, *b*, include the ratio of equity to total assets (*ea*) and the ratio of loans to total assets (*la*).[11] Finally, the country-specific control variables, *m*, include the European Bank for Reconstruction and Development (EBRD) index of banking sector reform (*ebrd*), a 3-bank concentration ratio (*conc3*), the asset share of majority state- and foreign-owned banks (denoted by *state* and *foreign*, respectively), the inflation rate (*inf*) and GDP per capita (*gdpcap*). Table 1 provides a formal definition, the sources and average values (on a country-specific basis) for all the variables included in Eq. (1).

Estimation of Eq. (1) using LR is carried out twice (corresponding to Models 1 and 2 below, respectively), the first time controlling only for *ea* and *la* and the second time using the full set of control variables.[12] Figure 1 illustrates the distribution of the estimated H-statistic for the full sample and for Models 1 and 2, along with the distribution of the associated disturbances. In addition, more detailed country-specific average results for all the coefficients and the H-statistic are provided in Tables 2 and

[9] We have decided to restrict the analysis to commercial banks only, so as to avoid comparing banks with different products, clientele and objectives.

[10] A better measure for the price of labor would be obtained if we divided personnel expenses by the number of employees; however, the latter variable is not available for many banks.

[11] Note that we avoid measuring bank size in terms of total assets because adding this measure to the revenue equation would make it *de facto* a price equation and might lead to a systematic bias in the estimation of the price parameters and therefore the H-statistic (see Vesala, 1995). Other bank-specific factors were used as additional explanatory variables to reflect e.g. differences in credit risk (measured by the ratio of loan loss provisions to assets). However, differences in the results were not statistically significant and therefore we decided to limit this analysis to the use of the *ea* and *la* variables.

[12] To capture country heterogeneity, both models include country dummies, but the results are not reported to save space. A number of robustness tests were carried out, including outlier analysis (i.e. trimming observations in the upper and lower 5 per cent of the distribution of the error term), scaling the nominal dependent variable by total assets (in the fashion of e.g. Claessens and Laeven, 2004) and inclusion or exclusion of a number of control variables (e.g. the loan loss provisions to loans ratio, the GDP growth rate and the short-term interest rate instead of the inflation rate). The results remained practically unchanged.

3, while country-specific distributions of the H-statistic are also illustrated in Figure 2. Noticeably, the noise component of both models shows small variations among banks, suggesting very good fit of the data. Model 2 slightly outperforms Model 1 in this respect, which may suggest that inclusion of country-level control variables is necessary.

A first interesting result is that the country averages of the estimated coefficients on input prices vary widely between countries, whereas the price of labor tends to have a negative effect on bank revenue in many of the banking systems examined. This implies that many banks function with excess labor capacity, a situation representing an efficiency problem that is a common element of banking systems of transition countries (see Brissimis et al., 2008). Moreover, the control variables exert the expected impact on bank revenue. Specifically, *ea* is negatively related to *rtr*, indicating on the one hand that equity capital is an expensive form of financing because there is an expected rate of return embedded in what someone is willing to pay today for future increases in equity prices, and on the other hand that capital requirements in the banking systems examined keep capital levels higher than optimal in the period under consideration.[13] Also, *la* bears a positive sign on average, especially for relatively developed banking systems (e.g. Czech Republic, Latvia and Slovenia). The impact of banking sector reform (*ebrd*) on bank revenue is clearly positive, while majority state-owned banks generate on average lower revenue and majority foreign-owned banks higher revenue. Finally, both inflation and GDP per capita have a positive effect on bank revenue, possibly because inflation is associated with wider margins, while rising GDP per capita leads to increased lending and thus revenue.

Given the considerable differences in the elasticities of input prices, the H-statistics vary extensively on an inter- and intra-country basis. The results from Model 1 (see column H1 in Table 2) indicate that the banking systems of most of the countries are characterized by monopolistic competition (H1 is between 0 and 1); however, banks operating in Croatia, Estonia, FYROM, Kazakhstan and Slovakia earn (on average) monopolistic profits (H1 is negative). Figure 2a shows that even the intra-country variation of market power of banks is quite significant (the distributions

[13] This is a common result among studies examining banking systems under reform (see e.g. Brissimis et al., 2008).

are scattered), implying that certain banks have greater either monopoly or monopsony power compared to other banks operating in the same country.[14] The average H-statistics for the countries examined are remarkably similar to the ones found by Delis (2008) and also conform to those derived with the Bresnahan's (1987) method by Brissimis et al. (2008). This greatly enhances confidence in the suitability and applicability of the new procedure.

When we incorporate the variables associated with the structural and macroeconomic environment, the H-statistics slightly increase on average, without significantly altering the findings about the competitive conditions (see Table 3). The results suggest that banks in Estonia, Kazakhstan, Slovakia and Slovenia are characterized by significant market power, as the respective H-statistics are clearly negative. Also, it can be noted that Model 2 generates a wider distribution of results (see Figure 2b), indicating a wider intra-country variation of bank market power. This finding, which is obviously related to the inclusion of the control variables *m*, may suggest that certain banks are able to better insulate their portfolios against structural and macroeconomic changes and thus earn higher rents or that certain banking systems have an oligopolistic core of banks with a competitive fringe.[15]

4. Concluding remarks

This paper proposes a new method for measuring the market power of individual banks, by combining well-established econometric and theoretical frameworks. Specifically, the local regression principle is used to estimate the model of Panzar and Rosse (1987) and thus bank-level coefficients on input prices are obtained, which are then summed up to calculate the H-statistic. The method is applied to bank-level data from 20 transition countries so as to get some insight into the power of the new method. In particular, the choice of the sample is motivated by (i) the existence of a recent body of literature on industry-level competition for these

[14] In fact, the results suggest that there is a trend toward more competitive practices in virtually all of the banking systems examined, as most of the observations with low values of the H-statistic are identified at the beginning of our sample period. This is an expected result given the regulatory reforms in the banking systems of the region and the increased privatization and foreign ownership. Since a more detailed discussion of the specific features of the banking systems examined is beyond the scope of the present analysis, graphs with a time dimension are not reported here, but are available on request.

[15] This may also be viewed as the failure of some bank managers to fully anticipate inflation, implying that above-normal revenue of some banks could be due to asymmetric information.

countries that allows comparison of the results and (ii) the well-documented transitional characteristics of these banking systems, which are usually associated with considerable differences of conduct across banks. Our findings suggest that country averages for the H-statistic are very close to their parametric equivalents, as derived in recent literature, implying that the proposed methodology is a useful tool for future analysis of the competitive conditions of the banking industry. In addition, the intra-industry bank-specific estimates suggest a significant variation of market power estimates across banks, mainly reflecting wide differences in elasticities with respect to the price of loans and labor.

Finally, note that estimation of other models of market power for a number of industries, such as the ones suggested by Bresnahan (1987), Roeger (1995) and Berry et al. (1995), is also possible using similar non-parametric techniques. In any case, this may be a desideratum for future research.

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Table 1
Descriptive statistics by country

Country	No of banks	No. of observations	rtr	w1	w2	w3	roa	ea	la	ebrd	conc3	state	foreign	inf	gdpcap
Albania	11	57	12642.4	0.030	1.048	0.011	0.016	0.126	0.315	2.41	76.81	40.69	58.00	2.38	1391.8
Armenia	15	71	3260.6	0.050	1.304	0.032	0.018	0.217	0.388	2.40	62.26	0.91	50.58	2.40	895.4
Azerbaijan	21	102	4795.6	0.081	1.098	0.025	0.022	0.270	0.528	2.23	83.39	60.24	5.15	3.33	1007.1
Belarus	17	92	1195.8	0.120	1.224	0.040	0.024	0.222	0.507	1.44	77.36	66.71	11.76	78.71	1928.7
Bulgaria	26	181	824.8	0.044	1.605	0.016	0.013	0.172	0.491	3.30	53.88	14.08	73.11	5.86	1879.5
Croatia	38	235	45577.8	0.043	1.007	0.018	0.013	0.158	0.535	3.63	57.38	8.58	82.93	3.19	5336.9
Czech Rep.	24	169	167943.5	0.109	2.896	0.010	0.008	0.103	0.377	3.68	64.81	11.00	77.38	2.46	6598.8
Estonia	10	55	72537.3	0.029	4.150	0.018	0.011	0.173	0.558	3.81	98.14	0.99	97.04	3.68	5739.5
FYROM	15	84	12838.8	0.036	0.964	0.022	0.011	0.303	0.456	2.70	78.02	1.73	44.85	2.26	1925.5
Georgia	11	66	5431.5	0.060	0.795	0.031	0.036	0.256	0.548	2.45	67.29	0.00	39.46	7.68	678.9
Hungary	30	180	102191.4	0.070	4.101	0.018	0.015	0.123	0.537	4.00	62.10	7.96	74.10	6.64	5706.1
Kazakhstan	24	140	24168.7	0.190	1.437	0.022	0.026	0.205	0.541	2.75	66.40	5.55	20.98	8.15	1928.0
Latvia	24	168	20547.0	0.023	1.441	0.017	0.012	0.123	0.410	3.48	54.59	3.69	59.85	4.00	4303.7
Lithuania	9	68	35020.6	0.028	0.756	0.021	0.005	0.128	0.508	3.25	82.23	11.63	79.50	1.24	4032.4
Moldova	16	92	2516.3	0.059	0.764	0.032	0.039	0.260	0.496	2.45	69.85	13.63	32.14	16.83	336.8
Poland	53	289	134323.2	0.067	4.419	0.019	0.012	0.145	0.520	3.40	55.21	23.74	69.53	4.05	4750.3
Romania	27	180	7609.9	0.090	1.328	0.033	0.006	0.180	0.444	2.81	67.25	31.23	56.88	23.91	2112.3
Slovakia	16	112	82663.3	0.046	4.302	0.011	0.007	0.096	0.417	3.34	77.48	13.95	77.06	7.06	4104.2
Slovenia	21	129	87437.3	0.038	0.871	0.014	0.012	0.108	0.530	3.30	64.29	24.61	17.93	5.64	10474.2
Ukraine	57	298	10899.8	0.078	1.157	0.024	0.018	0.166	0.582	2.33	49.77	10.54	15.81	12.56	912.0
Total	465	2768	49504.9	0.123	2.183	0.021	0.015	0.167	0.495	3.00	63.66	17.70	50.47	9.91	3211.2

Note: The table presents descriptive statistics of the sample on a country-specific basis. The variables are as follows: rtr is real total bank revenue (in thousand euros), w1 is interest expenses over total deposits, w2 is overheads minus personnel expenses over fixed assets, w3 is personnel expenses over total assets, roa is total profits over total assets, ea is equity over total assets, la is total loans over total assets, ebrd is the index of banking sector reform, conc3 is the 3-bank concentration ratio, state is the asset share of majority state-owned banks, foreign is the asset share of majority foreign-owned banks, inf is the inflation rate and gdpcap is GDP per capita in euros. All bank-level data are obtained from BankScope. ebrd, foreign, inf and gdpcap are obtained form the EBRD's Transition Reports and conc3 is obtained from the Beck et al. (2000) database as updated in 2007.

Table 2
Average coefficients of Model 1

Country	lnw1	lnw2	lnw3	lnea	lnla	cons	H1
Albania	0.636	0.036	-0.349	-1.685	0.475	6.385	0.324
Armenia	-0.175	0.317	0.223	-0.330	0.235	8.808	0.366
Azerbaijan	0.115	0.296	0.261	-0.911	-0.089	8.339	0.671
Belarus	0.404	0.287	-0.116	-0.515	-0.055	8.756	0.575
Bulgaria	0.306	0.107	-0.408	-0.297	0.222	7.311	0.005
Croatia	0.228	0.084	-0.447	-0.140	0.240	7.759	-0.135
Czech Rep.	0.399	0.125	-0.214	-0.161	0.294	9.252	0.310
Estonia	0.075	-0.301	-0.411	-0.074	0.173	7.584	-0.636
FYROM	-0.616	0.237	0.109	-0.952	-0.150	5.404	-0.270
Georgia	0.388	0.007	-0.072	-0.390	0.026	9.072	0.323
Hungary	0.329	0.059	-0.488	-0.194	0.099	7.793	-0.099
Kazakhstan	0.145	-0.023	-0.777	-0.053	0.161	6.268	-0.655
Latvia	0.476	0.039	-0.547	-0.088	0.117	8.108	-0.032
Lithuania	0.270	0.169	-0.525	-0.227	0.156	7.360	-0.086
Moldova	0.270	0.062	0.330	-0.486	-0.166	9.984	0.662
Poland	0.390	0.192	0.282	-0.367	0.079	10.662	0.864
Romania	0.747	0.076	-0.413	-0.533	-0.232	8.338	0.411
Slovakia	0.071	0.344	-1.501	1.733	-0.128	6.371	-1.086
Slovenia	-0.015	0.175	-0.252	-0.494	0.192	7.017	-0.092
Ukraine	0.380	-0.022	-0.265	-0.378	0.367	8.562	0.093
Average	0.287	0.105	-0.189	-0.282	0.121	8.173	0.203

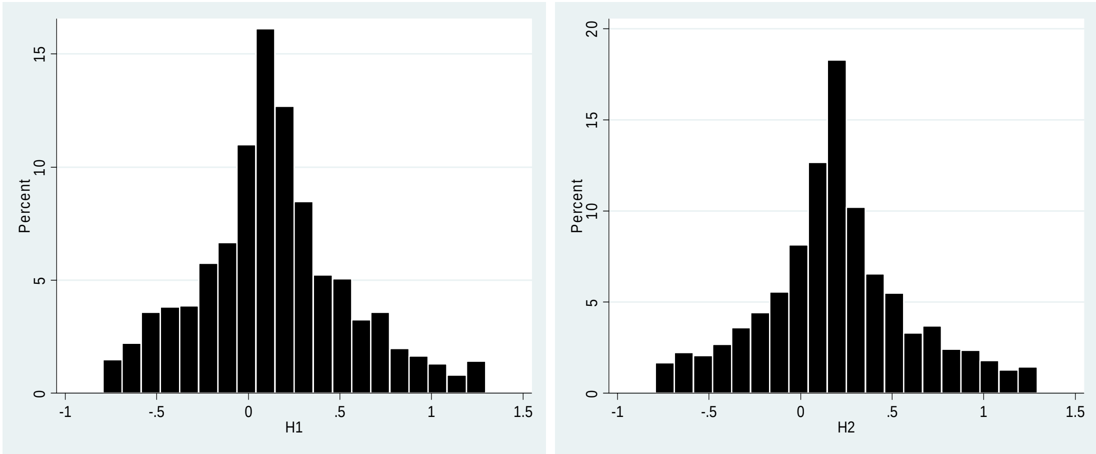
Note: The table presents average coefficients on a country-specific basis and for the whole sample obtained from the estimation of Eq. (1) (excluding the structural and macroeconomic control variables M) using LR. The variables are as follows (ln represents natural logarithm): Dependent variable is rtr, the real total bank revenue, w1 is interest expenses over total deposits, w2 is overheads minus personnel expenses over fixed assets, w3 is personnel expenses over total assets, ea is equity over total assets, la is total loans over total assets, cons is the constant term and H1 is the H-statistic calculated as lnw1+lnw2+lnw3 for each observation in the sample.

Table 3
Average coefficients of Model 2 by country

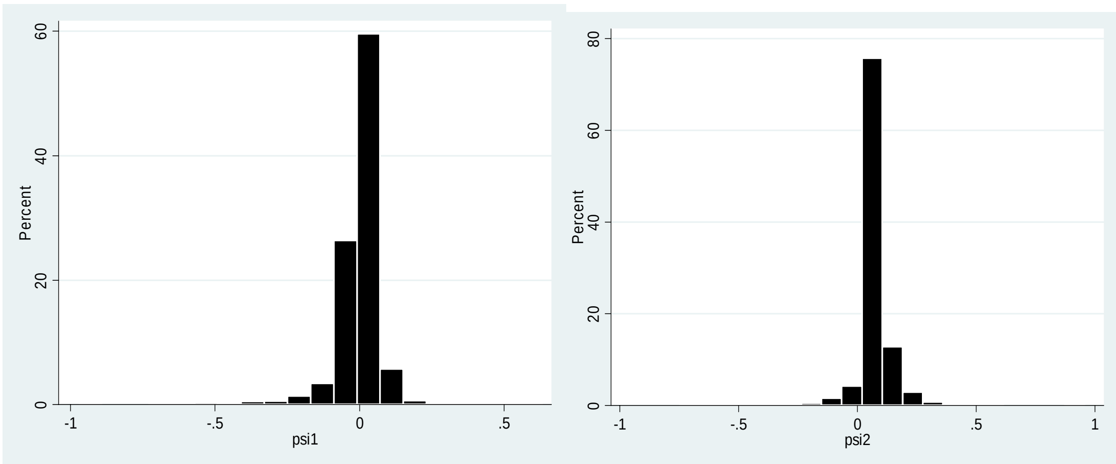
Country	lnw1	lnw2	lnw3	lnea	lnla	ebrd	conc3	state	foreign	inf	gdpcap	cons	H2
Albania	0.783	-0.054	0.153	-1.499	0.102	1.569	3.249	-0.311	0.022	-0.007	0.118	2.752	0.882
Armenia	-0.271	0.445	0.234	-0.246	0.223	0.189	0.023	-0.004	-0.007	0.041	0.153	6.473	0.408
Azerbaijan	0.161	0.238	0.147	-0.699	-0.138	-0.016	-0.093	-0.048	0.008	0.064	0.121	6.920	0.547
Belarus	0.220	0.079	0.117	-0.543	-0.007	-1.610	-0.320	-0.484	-0.020	0.079	0.742	13.404	0.417
Bulgaria	0.342	0.082	-0.210	-0.250	0.116	0.146	0.053	-0.093	0.001	0.047	0.115	5.958	0.214
Croatia	0.229	0.118	-0.287	-0.171	0.227	0.326	0.026	0.159	0.102	0.038	0.085	6.056	0.059
Czech Rep.	0.365	0.092	-0.124	-0.254	0.236	0.114	0.053	0.031	-0.039	0.043	0.090	7.125	0.333
Estonia	0.144	-0.252	-0.030	-0.107	0.019	1.404	0.072	0.025	-0.019	-0.013	0.106	4.764	-0.137
FYROM	-0.631	0.212	0.322	-1.026	-0.251	0.636	0.113	-0.066	0.122	0.023	0.145	3.189	-0.097
Georgia	0.403	0.031	-0.266	-0.309	-0.030	-0.074	0.043	0.123	0.224	0.041	0.046	5.563	0.169
Hungary	0.341	0.033	-0.297	-0.232	-0.003	0.519	0.110	-0.089	0.099	0.038	0.114	5.294	0.078
Kazakhstan	0.035	-0.068	-0.465	-0.101	-0.008	-0.019	-0.032	-0.086	0.115	0.051	0.108	6.070	-0.498
Latvia	0.410	-0.001	-0.427	-0.152	0.237	0.323	-0.038	-0.008	0.056	0.019	0.107	6.552	-0.017
Lithuania	0.225	0.083	-0.213	-0.293	0.093	-0.317	-0.034	-0.049	0.013	0.054	0.110	7.485	0.096
Moldova	0.300	0.030	0.109	-0.382	-0.120	0.838	0.122	-0.047	-0.047	0.029	0.080	5.330	0.438
Poland	0.361	0.130	0.334	-0.387	0.016	-0.373	0.057	-0.091	0.024	0.058	0.114	9.316	0.825
Romania	0.707	0.011	-0.332	-0.540	-0.130	0.631	0.111	0.008	0.052	0.026	0.126	4.693	0.386
Slovakia	0.184	0.352	-1.172	1.364	-0.135	0.408	0.046	0.030	0.025	0.036	0.116	4.282	-0.635
Slovenia	-0.183	0.173	-0.113	-0.489	0.128	0.426	0.062	-0.051	0.010	0.006	0.127	4.965	-0.123
Ukraine	0.487	0.010	-0.151	-0.305	0.216	1.136	-0.115	0.055	0.066	0.028	0.041	5.198	0.346
Average	0.279	0.083	-0.158	-0.293	0.067	0.337	0.020	-0.022	0.005	0.036	0.117	6.147	0.204

Note: The table presents average coefficients on a country-specific basis and for the whole sample obtained from the estimation of Eq. (1) (including the structural and macroeconomic control variables M) using LR. The variables are as follows (ln represents natural logarithm): Dependent variable is rtr, the real total bank revenue, w1 is interest expenses over total deposits, w2 is overheads minus personnel expenses over fixed assets, w3 is personnel expenses over total assets, ea is equity over total assets, la is total loans over total assets, cons is the constant term, ebrd is the index of banking sector reform, conc3 is the 3-bank concentration ratio, state is the asset share of majority state-owned banks, foreign is the asset share of majority foreign-owned banks, inf is the inflation rate, gdpcap is GDP per capita and H2 is the H-statistic calculated as lnw1+lnw2+lnw3 for each observation in the sample.

Figure 1: Frequency distribution of market power estimates of banks and distribution of noise



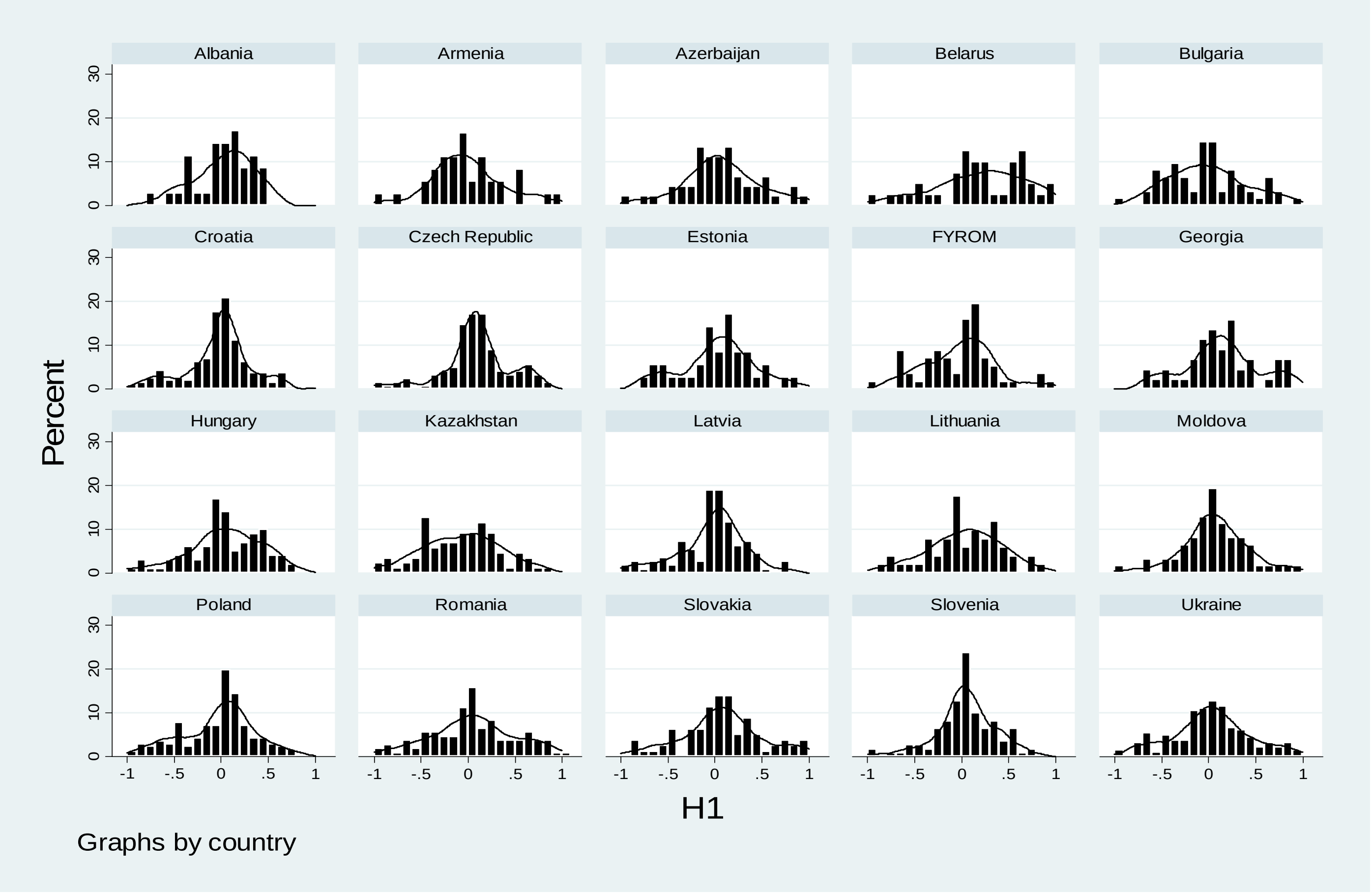
H-statistics obtained from Model 1 H-statistics obtained from Model 2



Noise of Model 1 Noise of Model 2

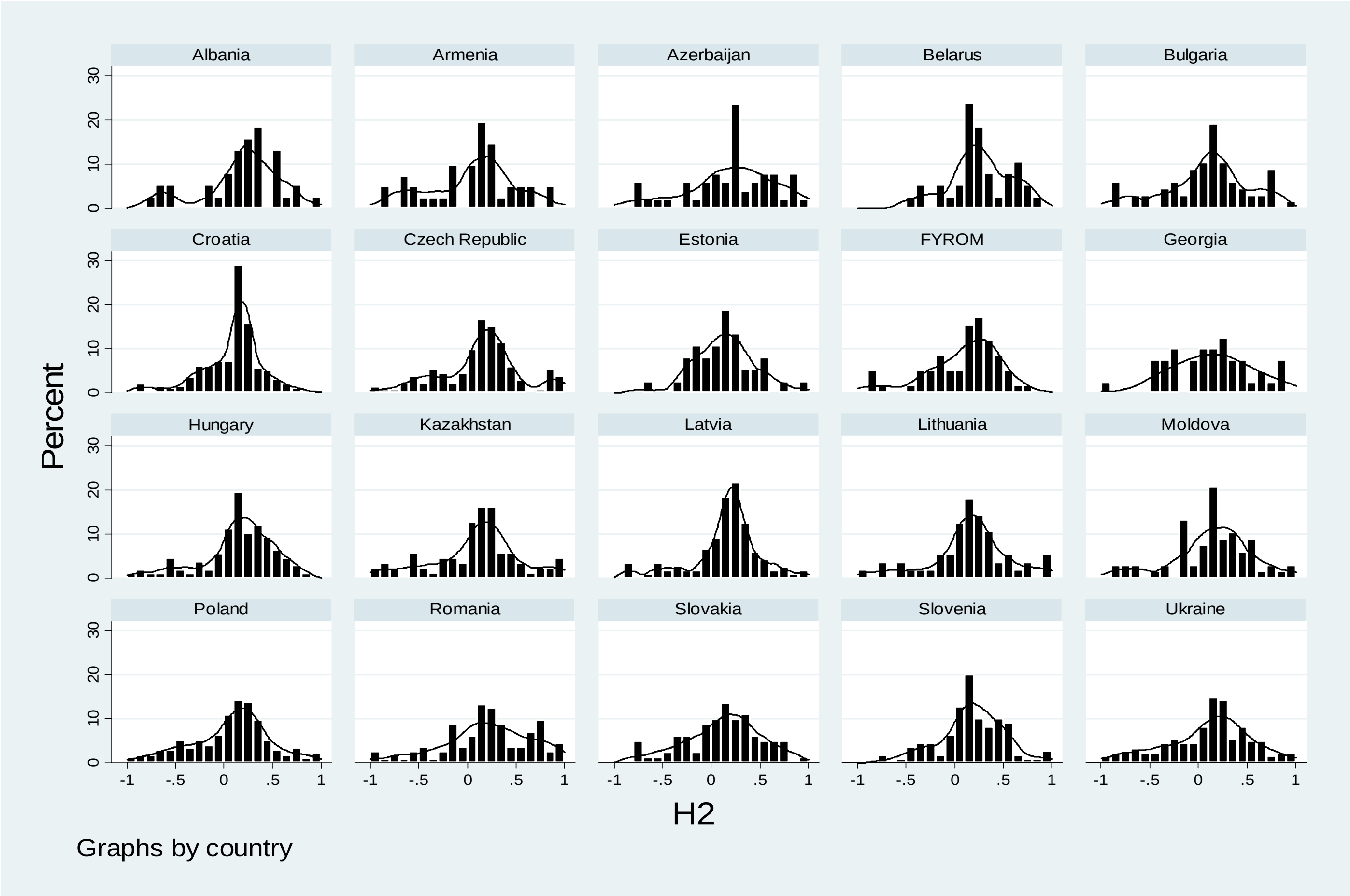
Note: The first two histograms present the distribution of the H-statistic obtained from the estimation of Models 1 (excludes structural and macroeconomic variables) and 2 (includes these variables) with local regression. The other two histograms present the distribution of the noise components generated from the two Models (psi1 and psi2, respectively).

Figure 2a: Market power estimates of banks by country (obtained from the estimation of Model 1)



Note: The histograms present on a country-specific basis the distribution of the H-statistic obtained from the estimation of Model 1 with local regression. For expositional brevity, a kernel density line is added to the graphs.

Figure 2b: Market power estimates of banks by country (obtained from the estimation of Model 2)



Note: The histograms present on a country-specific basis the distribution of the H-statistic obtained from the estimation of Model 2 with local regression. For expositional brevity, a kernel density line is added to the graphs.

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