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Overview

The ***Journal of Banking and Financial Economics*** **(JBFE)** is an open access journal. The submission of manuscripts in free of fee payment. This journal follows a double-blind reviewing procedure.

Aims and Scope

JBFE publishes high quality empirical and theoretical papers spanning all the major research fields in banking and financial economics. The aim of the journal is to provide an outlet for the increasing flow of scholarly research concerning banking, financial institutions and the money and capital markets within which they function. The journal also focuses on interrelations of financial variables, such as prices, interest rates and shares and concentrates on influences of real economic variables on financial ones and vice versa. Macro-financial policy issues, including comparative financial systems, the globalization of financial services, and the impact of these phenomena on economic growth and financial stability, are also within the *JBFE*'s scope of interest. The Journal seeks to promote research that enriches the profession's understanding of the above mentioned as well as to promote the formulation of sound public policies.

Main subjects covered include, e.g.: [1] ***Valuation of assets***: Accounting and financial reporting; Asset pricing; Stochastic models for asset and instrument prices; [2] ***Financial markets and instruments:*** Alternative investments; Commodity and energy markets; Derivatives, stocks and bonds markets; Money markets and instruments; Currency markets; [3] ***Financial institutions, services and regulation***: Banking efficiency; Banking regulation; Bank solvency and capital structure; Credit rating and scoring; Regulation of financial markets and institutions; Systemic risk; [4] ***Corporate finance and governance****:* Behavioral finance; Empirical finance; Financial applications of decision theory or game theory; Financial applications of simulation or numerical methods; Financial forecasting; Financial risk management and analysis; Portfolio optimization and trading.

Special Issues

JBFE welcomes publication of Special Issues, whose aim is to bring together and integrate work on a specific theme; open up a previously under-researched area; or bridge the gap between formerly rather separate research communities, who have been focusing on similar or related topics. Thematic issues are strongly preferred to a group of loosely connected papers.

Proposals of Special Issues should be submitted to at jbfe@wz.uw.edu.pl. All proposals are being reviewed by the Editorial Team on the basis of certain criteria that include e.g.: the novelty, importance and topicality of the theme; whether the papers will form an integrated whole; and the overall 'added value' of a Special Issue.

CONTENTS

Fareword

This special issue of the *Journal of Banking and Financial Economics* focuses on the *Applications of Quantitative Methods in Banking and Finance*. The impact of financial markets and banking sector activities on economic processes has always been discussed in mainstream literature. It became an area of particular interest since the global financial crisis. We tried to encourage submissions related to public finance, corporate finance, corporate governance, financial markets, and macroeconomic processes. This resulted in numerous articles submitted for review. After careful assessment, we have chosen seven papers that we consider the best. However, the overall submissions were of high quality and had interesting implications.

Magdalena Mikołajek-Gocejna attempts to bridge the gap between the theoretical and empirical literature on the stability of ESG companies' systematic risk in the first paper. The author analyses the stability of beta coefficients of 57 companies listed in WIG-ESG and concludes that those coefficients are not stable in the short term.

Another article by Jan Koleśnik and Jacek Nadolski also analyses the financial markets, particularly banking. The authors verify the conceptual model of integrated optimisation of the bank's value and show that it is possible to ensure a compromise between the safety and effectiveness of the bank's operations, which allows for a long-term competitive advantage.

The next, third article by Rafał Zbyrowski concerns the long-term economic processes. The author analyses the long-term price volatility of KGHM shares using a cointegration approach and indicates that those shares are long-term dependent on the quotation of HG copper prices.

The fourth article, written by Marta Idasz-Balina, Rafał Balina, Adam Zając, and Krzysztof Smoleń, analyses the relationship between expenditure on corporate philanthropy and the financial performance of cooperative banks in Poland. The authors use the GMM estimator and data for 2013–2020 to show the positive importance of corporate philanthropy in shaping financial efficiency.

Dorota Żebrowska-Suchodolska and Andrzej Karpio, in the fifth published article, analyse the behaviour of selected segments of the Polish financial market in the initial period of the pandemic and the Russian-Ukrainian war. The research, based on the Warsaw Stock Exchange and selected world stock exchanges, shows the high level of development of the Warsaw Stock Exchange compared to other stock exchanges.

The cross-country analysis is also the subject of the sixth article written by Mateusz Pipień and Abhishek Anand. The research on the procyclicality of loans provided by banks based on data from 13 OECD countries over 16 years indicates that bank-specific variables are more significant as loan supply determinants than macroeconomic variables.

The seventh article, written by Aleksandra Majchrowska and Paweł Strawiński, also links the real economy and financial issues at the regional level. The authors analyse the employment effects of minimum wage increases for different age groups of workers simultaneously in 16 Polish regions in 2006–2020. Their results confirm the differences in employment elasticity for different groups of workers. However, the authors conclude that analyses at the aggregated level might underestimate the employment effects of the minimum wage.

Application of Chow, Cusum and Rolling Window in Testing Stability of Systematic Risk of Companies Listed in WIG-ESG in 2019–2022

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ABSTRACT

The aim of the article is to analyze the stability of beta coefficients of companies listed in WIG-ESG. There are many studies on the stability of companies' systematic risk, but the literature and research lack an analysis of the stability of the beta coefficient for ESG companies.

We examined beta coefficients for 57 companies listed in WIG-ESG, established for sets of daily rates of return between September 3, 2019, to June 6, 2022 (period including COVID-19 crisis and asset price inflation, Russian invasion of Ukraine). We estimate the beta coefficient for the whole as a result of which we obtain the average value of the beta coefficient over the entire analyzed period, and subperiods with fixed length rolling window, resulting in a time series of beta coefficients. To assess beta stability, we used the Chow test with the F statistic, the Cusum test based on generalized fluctuations test framework, and the Wald-Wolfowitz runs test of randomness around the mean for the time series beta coefficients obtained in the rolling window.

The considered tests argue for the instability of the time series of beta coefficients in most of the companies tested: 93% short-term instability cases confirmed by the Chow test, 100% short-term instability cases confirmed by the Wald-Wolfowitz runs test.

The paper is an initial attempt to bridge the gap that presently exists between the theoretical and empirical literature on the stability of ESG companies' systematic risk.

It cannot be ruled out (hypothesis) that the beta coefficient for companies listed in the WIG-ESG index is/will be stable over longer periods of time.

JEL Classification: G11, G12, G13

Keywords: Capital Asset Pricing Model (CAPM), beta coefficient, systematic risk, ESG, environment, social and governance criteria, Cusum Test, Chow Test, rolling window.

1. INTRODUCTION

In recent years, sustainable finance has become one of the most important trends, especially in developed capital markets. Investors, market supervisory authorities and companies, by considering ESG (environmental, social and corporate governance) factors, respond to global challenges that we all face and will face in the coming decades. ESG factors, although they present

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current data, refer primarily to the future, because they show how to effectively manage long-term risk and create value not only for shareholders, but for all stakeholders of the company. The companies that meet the social, environmental, and corporate governance criteria are more aware of the changes taking place in the world, thanks to which they better forecast their future situation, and their operations are more stable and sustainable. In the fifth edition of the GPW survey on the impact of ESG factors on investment decisions, 81% of professional stock market investors in Poland assessed that companies that have implemented the ESG strategy are perceived as entities with lower risk. (GPW, 2019). Moreover, companies with a strong ESG profile are less vulnerable to systematic market shocks and therefore show lower systematic risk (Mikołajek-Gocejna, 2022, pp. 597–615).

Identifying and measuring risk have been of constant interest to both financial theoreticians and practitioners. Various theories have been propounded for pricing of assets considering the risk element. The most common and widely accepted method has been the capital asset pricing model (CAPM) model, which takes into consideration the systematic risk of the asset, measured as the beta coefficient.

The beta coefficient is defined as the ratio of the covariance of the rate of return of the examined financial instrument R_i and the rate of return of the market portfolio R_m to the variance of the rate of return of the market portfolio (Tofallis, 2008, p. 1359):

$$\beta_i = \frac{cov\,(R_i, R_m)}{var(R_m)} = cor(R_i, R_m) \times \sqrt{\frac{var(R_i)}{var(R_m)}}, \tag{1}$$

where:
R_i – measures the rate of return of the financial instrument,
R_m – measures the rate of return of the market portfolio,
$cov\,(R_i, R_m)$ is the covariance between the rates of return.

In general, the calculation of the beta coefficient is based on comparing volatility of the rate of return from shares of a specific company in the adopted unit of time with volatility of the rate of return from the stock exchange portfolio (index) adopted for comparison (Dharmaratne, Harris, 2006, pp. 68–61). Since volatility – in this case, of the rate of return – reflects the risk of their realization, the measurement of the beta coefficient means the measurement and comparison of risks related to the investment in the shares of a given entity and the average, previously defined market portfolio, respectively (this measurement should concern the expected rate of return, practice shows however, that beta is calculated on the basis of historical, i.e. realized rate of return).

The beta coefficient is also an estimator of the parameter of simple linear regression equation proposed by Sharpe (1963). Therefore, the rate of return on shares of the *i*-th company in the *t*-th period can be written as (Elton, Gruber, 1998, p. 154; Jajuga, Jajuga, 1998, p. 63):

$$R_{it} = \alpha_i + \beta_{i\,Sharp} R_{mt} + \varepsilon_{it}, \tag{2}$$

where:
R_{it} – rate of return of shares of the *i*-th company,
R_{mt} – rate of return on an index of the market,
α_i – the free expression of the model, which is a component of the return on shares of the company and independent of the market situation,
$\beta_{i\,Sharp}$ – the direction coefficient constant over time which measures the expected change in R_i depending on the change in R_m,
ε_{it} – is Gaussian noise N $(0, \sigma_i)$ with zero as expected value and standard deviation σ_i,
t – number of observations of the time series.

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In the Capital Asset Pricing Model, there is an additional variable: risk-free rate of return R_F: (Treynor (1961), Sharpe (1964), Lintner (1965a, 1965b), Mossin (1966))

$$R_i = R_F + \beta_{i\,CAMP}(R_m - R_F) + \varepsilon_i. \tag{3}$$

In the equation, the risk-free return R_F can be a deterministic constant or a random variable.

CAPM is the most frequently and most willingly model of estimation of the cost of capital used in practice, due to its easy implication and interpretation.

The beta coefficient is also called stock aggressiveness. Malkiel and Xu (2006) identified this type of risk as the systematic risk, which is undiversifiable.

Possibilities of using beta in the practice of investment processes are closely related not only to the correctness of its estimation, but also its stability over time (Wright, Mason, and Miles 2003). The Sharpe model and Capital Asset Pricing Model assume that beta is stable and predictable over time. (Treynor, 1965, pp. 63–75).

Thus, the main hypothesis of the article is that beta coefficients of ESG companies listed on the Polish capital market are not stable in short time. Despite the problem of beta stability is quite well described in the literature, results of the stability tests carried out over the years by various researchers are ambiguous, inconclusive, and contradictory. Moreover, literature and research lack an analysis of the stability of the beta coefficient for ESG companies. This paper is an initial attempt to bridge the gap that presently exists between the theoretical and empirical literature on the stability of ESG companies' systematic risk.

2. STABILITY OF ESG COMPANIES BETA – LITERATURE REVIEW

An important issue from the point of view of forecasting and the possibility of making investment decisions on the basis is the analysis of beta stability over time and the study of the sensitivity of its assessments to changes in the method of estimating the model and measurement of variables. Beta instability causes low predictive efficiency of the model, as makes it impossible to use the dependencies described by the model in the future. Moreover, inference based on a model with unstable parameters may result in large errors.

2.1. Systematic risk of ESG companies

Literature and research lack an analysis of the beta coefficient stability for ESG companies. Thus, two groups of publications were analyzed. The first covered research on the risk of ESG companies, the second, stability of beta coefficients. It was necessary to combine the two issues and carry out studies on the stability of the systemic risk for ESG companies.

In the literature, there are not many cases of studies analyzing systematic risk of ESG companies or the relationship between ESG factors and company-specific risk (Sassen, Hinze, and Hardeck, 2016; Mikołajek-Gocejna, 2022). Most studies show, that involvement in social and environmental activities leads to improvement in an organization's image, and its credit ratings, as well as lowering the cost of capital (Gangi et al., 2020; Xue et al., 2020), caused largely by a decrease in risks measured appropriately, e.g., by the standard deviation of rates of return or the beta coefficient.

Boutin-Dufresne and Savaria research (2004) showed that corporate social responsibility activities can help diminish the overall business risk of a company, and even improve its long-term risk-adjusted performance.

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Negative correlation between systematic risk and CSR was also confirmed by Jo and Na (2012). Orlitzky and Benjamin (2001) reviewed 18 American cases of studies on the relationship between corporate social performance (CSP) and risk, indicating that integration of ethical factors in corporate management leads to their lower exposure to financial risk. Similar results were obtained by Boutin-Dufresne and Savaria for Canadian firms (2004). Albuquerque et al. (2019) examined the relationship between CSR and firms' systematic risk using a sample of 28578 annual observations of the United States companies over the period 2003–2015 and found that the level of systematic risk is lower for companies with better CSR performance. Similar results were obtained by Shakil (2021), Rehman et al. (2020) and Zhou et al. (2020).

Analysis conducted by Hassan et al. (2021) showed that companies that follow stricter ESG principles are more resilient to systematic market shocks regardless of their country of origin. The authors analyzed 4624 non-financial firms from Africa, Asia, Europe, Latin America, North America, and Oceania over the period 2002–2018. Moreover, Dunna et al. (2018), concluded that high-scoring ESG stocks have lower volatility and betas than lower scoring ESG stocks.

Research conducted by Bouslah, Kryzanowski, and Mzali (2011) showed that not all ESG aspects affect the systematic risk of companies. Employee relations, environment, human rights and corporate governance negatively affect firm risk, but other dimensions (community, diversity and product) do not significantly impact firm risk. Thus, next to the studies that used aggregated ESG measures, there are studies based on individual ESG measures as explanatory variables. For example, Sharfman and Fernando (2008) confirmed the negative correlation between the cost of equity (beta coefficient) and the quality of environmental management in American companies. Zaman et al. (2021) found that eco-innovative companies are less risky. Xue et al. (2020) claimed that involvement in environmental activities can consequently reduce financial risk. Similar results were obtained by Salama et al. (2011). Moreover, Zaman et al. (2021) found a negative relationship between eco-innovation and stock price crash risk. In turn, research conducted by Chen et al. (2020) showed that there is a negative correlation between the dominant role of institutional investors in the shareholding structure of a company and its risk.

2.2. Stability of systematic risk

The problem of beta stability is quite well described in the literature, however, the results of stability tests carried out over the years by various researchers are ambiguous, inconclusive, and contradictory. Most of the analyses were conducted in developed markets, but there are also studies on the stability of systematic risk for companies listed on developing markets. They include both studies on individual stocks as well as portfolios.

Results of empirical work on beta instability can be divided into three groups: those that confirm that beta is stable over time, those that confirm its instability and those that give ambiguous indications (Table 1).

The existence of stability of beta over different phases of the market was confirmed by analyses conducted by Shamsher et al. (1994), Fabozzi and Francis (1977), Fisher and Kamin (1985) Faff (2001), Das (2008), George and Bainy (2012), Harish and Mallikarjunappa (2019).

Several studies documented that beta is time varying because of the influence of micro-economic and macro-economic factors. The time varying nature of beta at the New York Stock Exchange was first discovered by Blume (1971). Instable betas were also confirmed by researches conducted by Sunder (1980), Bos and Newbold (1984) Russel, Impson and Imre (1994), Braun et al. (1995), Brooks et al. (1998), Faff, Hillier, Hillier (2003), Shah, Moonis, (2003), Irala (2007), Sarma and Sarmah (2008), Attya and Eatz (2011), Simon et al. (2012), Mazowina (2013), Celik (2013), Wijethunga and Dayaratne (2015), Ye (2017), Gupta (2020)

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Contradictory results in beta stability were obtained by: Baesel (1974), Levy (1971), Witkowska (2008), Singh (2008), Ray (2010), Deb and Mistra (2011), Terceño et al. (2011), Dubey (2014), Dębski et al. (2011), Ye (2017), Mikołajek-Gocejna (2021).

One of the most widely used methods to estimate beta as a time series process is the Kalman Filter (Kalman, 1960). It has been applied for the estimation of betas and tests for beta constancy in several studies (e. g. Bos, Newbold, 1984; Fisher, Kamin, 1985; Shah, Moonis, 2003). Kalman filters for beta estimation also presented difficulties, due to their failure to deal with the problem of heteroskedasticity (Fisher, Kamin, 1985).

3. METHODOLOGY AND DATA

3.1. Systematic risk estimation and data

In the study, we will estimate the beta coefficient as an estimator of the parameter of simple linear regression equation proposed by Sharpe (1963).

$$R_{it} = \alpha_i + \beta_{i\,Sharp} R_{mt} + \varepsilon_{it}, \quad (4)$$

where t is the index of the moments of time from the period T from which samples of the analyzed rate of returns for the i-th company are derived.

We examined beta coefficients for 57 companies listed in WIG-ESG, established for the sets of daily rates of return between September 3, 2019, to June 6, 2022 (period including COVID-19 crisis and asset price inflation, Russian invasion of Ukraine). To obtain an up-to-date beta rating, the model should be estimated over a relatively short period of time, while maintaining the estimation sample size requirements. Therefore, our studies prefer daily quotations, however we are aware of the limitations of the approach.[1] According to the theoretical assumptions of the Sharp/CAPM model, the market index should cover the broadest spectrum of investment instruments available to investor. Thus, we choose the rate of return from the WIG Index (market index) as the variable explaining the rates of return of individual ESG companies.

We estimate the beta coefficient for:

1) the whole, as a result of which we obtain the average value of the beta coefficient over the entire analysis period,
2) and subperiods with fixed length rolling windows, resulting in a time series of beta coefficients.

In the study covering the whole period, we used the beta coefficient estimation by the OLS regression of the Sharp equation (4), which ensures that estimators are unbiased (or at least asymptotic, unbiased and consistent when the variable R_m is random):

$$\begin{bmatrix} \alpha_i \\ \beta_{i\,Sharp} \end{bmatrix} = [R_m{}' R_m]^{-1} R_m{}' R_i, \quad (5)$$

where:

R_i – (n x 1) vector of daily return on assets i,

R_m – (n x 2) matrix of daily return on a market portfolio proxy with 1 in the first column (for intercept).

[1] The use of daily returns avoids the dilemma of how to estimate them that accompanies longer intervals. In addition, aggregating daily returns to e.g., monthly returns causes a loss of important information. An important argument for the use of high-frequency data is also the possibility of obtaining a relatively long sample for a short period of time (i.e., many observations, which gives relatively low standard errors)

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This method is the simplest computationally, although it is numerically less efficient than the one used of definition (1) and efficient recursive algorithms for calculating moments. Due to the purpose of the research, we prioritize the ease of calculations over their efficiency.

In the rolling regression (Zivot and Wang, 2006, pp. 342–349), period T is divided into sub-periods:

1) containing the same number of 20 observations,
2) which we shift in the time domain by one observation (rolling window) from the beginning to the end of the period T,
3) beta coefficient $\beta_{i\ Sharp}(t)$ estimated for the data from a given subperiod (window) is assigned to the end of t of the subperiod:

$$\begin{bmatrix} \alpha_i(t) \\ \beta_{i\ Sharp}(t) \end{bmatrix} = \left[R_m(t)'\, R_m(t)\right]^{-1} R_m(t)' R_i(t), \tag{6}$$

where:

$R_i(t)$ is an (20 x 1) vector of daily return on assets i in which the first element is $R_{i\,t-19}$ and the last is $R_{i\,t}$,

$R_m(t)$ is an (20 x 1) matrix of daily return on a market portfolio proxy in which the first row is the vector $(1, R_{m\,t-19})$ and the last is $(1, R_{m\,t})$.

As a result of the procedure, we obtain a time series of estimated beta coefficients.

The 20-day length of the time window is dictated by the length of the series (686 days), the daily data frequency that corresponds to the average length of month, and by the desire to obtain a given degree of data smoothing, and the number of regressions required (667 for each of the 57 companies). Assigning the result of the beta parameter estimation to the end of the interval, results in no beta assigned to the initial 19 days period.

In the estimation, we assume that the random regression component ε_i is normally distributed.

3.2. Stability testing

The issue of beta stability can be treated as a problem of invariance of their estimates, and it applies both to its stability over time, as well as to no sensitivity to changes in the method and frequency of measurement of variables and methods of model estimation (Tarczyński et al., 2013, p. 71).

To assess beta stability, we used:

1) Chow test (Chow, 1960), with the F statistic,
2) Cusum test (Ploberger and Kramer, 1992), based on the generalized fluctuations test framework,
3) Wald-Wolfowitz runs test of randomness around the mean for the time series beta coefficients obtained in rolling windows.

In the Chow test period T with daily data is divided into two parts T_1 and T_2 with a shifting time of division from the 20th day from the beginning to the 20th day before the end of period T. Thus, the division point covers all possible dates for dividing the series into two disjoint parts with a minimum number of 20 observations in each part. The test compares OLS residuals estimated (just like (5)) from models estimated separately in T_1 and T_2 subsamples with OLS residuals estimated for the whole series[2]:

[2] Here and further designations in the equation adapted to the designations of the variables in the article.

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$$F_i(t) = \frac{\hat{u}_i^T \hat{u}_i - \hat{e}_i^T(t)\hat{e}_i(t)}{\hat{e}_i^T(t)\hat{e}_i(t) / (n-2k)}, \tag{7}$$

where:
$t = 20 \ldots (n-19)$ is the division point of period T,
$\hat{u}_i$ – OLS residuals from the model, where parameters are fitted for all observations,
$\hat{e}_i(t)$ – OLS residuals from the full model, where coefficients in subsamples $1 \le T_1 \le t$ and $t+1 \le T_2 \le n$ are estimated separately,
n – number of observations,
k – number of regression coefficients ($k=2$),
$F_i(t)$ – has an asymptotic χ^2 distribution with k degrees of freedom ($F_i(t)/k$ has a F distribution with k and $n-2k$ degrees of freedom).

The Cusum test (Ploberger and Kramer, 1992) is based on the recursive residuals. We estimate a simple OLS model (just like (5)) for sub-periods from the first observation to the end of the sub-period. The end of the sub-period varies from the third observation ($k+1$, where $k=2$) to the one before last one ($n-1$). Based on the obtained estimators for data up to the moment $t-1$, we predict the value of the rate of return R_{it} of the financial instrument for the moment t with an error:

$$\hat{u}_{it} = R_{it} - [\alpha_i(t-1) + \beta_{i\,Sharp}(t-1)R_{mt}], \tag{8}$$

where:
$\alpha_i(t-1)$, $\beta_{i\,Sharp}(t-1)$ – OLS estimate (6) based on all observations up to $t-1$ of assets i,
$t = (1+k) \ldots n$, (for $k=2$, $t=3 \ldots n$)

The variance of predictor is $\sigma^2\left(1 + [1R_{mt}](R_m(t-1)'R_m(t-1))^{-1}\begin{bmatrix}1\\R_{mt}\end{bmatrix}\right)$, where: $R_m(t-1)$ is a ($t-1$ x 2) matrix of monthly return on a market portfolio proxy in which the first row is the vector (1, R_{m1}) and the last is (1, $R_{m(t-1)}$), and σ^2 is the variance of disturbance. After scaling the $\hat{u}_{it}$ errors, we get recursive residuals:

$$w_{it} = \frac{\hat{u}_{it}}{\sqrt{\left(1 + [1R_{mt}]R_m(t-1)'R_m(t-1)\right)^{-1}\begin{bmatrix}1\\R_{mt}\end{bmatrix}}}, \tag{9}$$

with the zero expected value and constant variance σ^2 (homoscedasticity)[3].

We cumulatively sum the standardized recursive residuals obtaining:

$$W_{i\tau} = \frac{1}{\tilde{\sigma}\sqrt{\eta}} \sum_{t=3}^{2+\tau} w_{it}, \tag{10}$$

where:
$\eta = n-2$ is the number of recursive residuals,
$\tau = 1 \ldots \eta$ is the index of cumulative sums of recursive residuals,
$\tilde{\sigma} = \frac{1}{n-2}\sum_{j=3}^{n}(w_{it} - \overline{w_\iota})^2$ is the variance estimate of w_{it}.

[3] OLS residuals are not homoscedastic, even if the variance of the disturbance is constant.

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When the residuals are Gaussian noise, their cumulative sum takes the form of the Standard Brownian Motion (Wiener Process) (Latała, 2011). It allows to establish the limit $\pm b(\tau)$ beyond which fluctuations indicate changes in the parameters of the regression function (instability):

$$b(\tau)=\lambda(1+2\tau), \tag{11}$$

where:
λ determines the confidence level.

Stability of beta coefficients obtained by the moving window method can be understood from the statistical point of view as randomness of their deviations from mean value in the period 2019–2022. With beta stability, deviations from the mean are followed by a rapid return to it (white noise). When deviations cluster we are dealing with instability. For example: in the case of a linear upward trend, initially there is a group of negative deviations from the mean, then positive ones. For a polynomial trend (from quadratic upwards) and a sinusoidal trend, negative and positive deviations from the mean occur alternately. For a stochastic trend (random straying, unit root), a group of positive (or negative) and negative (or positive) deviations from the mean (depending on the observed realization of the process) should be expected successively.

The Wald-Wolfowitz runs test of randomness was used to test the randomness of deviations from the mean. The Wald-Wolfowitz runs test is a nonparametric test (distribution-free). There is no need to make restrictive assumptions concerning the specific distribution.

The Wald-Wolfowitz runs test statistic (Z) compares the realized number R of series of positive and negative deviations from the mean in a time-ordered series with the expected number of series at random deviations from the mean $E(R)$:

$$Z=\frac{R-E(R)}{\sqrt{VAR(R)}}, \tag{12}$$

where:

$E(R)=\dfrac{2n_1 n_2}{n_1+n_2}+1,$

$VAR(R)=\dfrac{2n_1 n_2(2n_1 n_2-n_1-n_2)}{(n_1+n_2)^2(n_1+n_2-1)}+1,$

n_1, n_2 – the number of positive and negative deviations from the mean.

The Z statistic has a standardized normal distribution for a large sample ($n_1>10$, $n_2>10$).

3.3. Stages of research

The research included six stages:

1) a priori selection of the significance level of the tests $\alpha=0.05$ (probability of the first type error),
2) estimation of the beta coefficient for the data for the entire period from a given company using the linear OLS regressions,
3) regression stability testing (including beta coefficient) using Chow and Cusum tests,
4) estimation of the time series of beta coefficients in subperiods using the rolling window method (OLS),
5) the Wald-Wolfowitz runs test of randomness around the mean for the time series beta coefficients obtained in rolling windows (ad. 4).

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We conducted the research in the *R* statistical program environment using the test implementation:

1) the Chow test in F Statistics (Fstats) and Cusum test in Empirical Fluctuation Processes (efp), from library: Testing, Monitoring, and Dating Structural Changes (Strucchange) (Zeileis et al., 2002),
2) Wald-Wolfowitz Runs Test (runs.test) from library: Testing Randomness in R (Caeiro, Mateus, 2022),

The data set includes daily rates of return on the market index, rates of return on the index of the analyzed asset for given moments of time (panel data) from September 5, 2019, to June 6, 2022 (observations from 686 periods – days). Simple rates of return were used. Data provider: Warsaw Stock Exchange

4. RESULTS AND DISCUSSION

4.1. Beta stability in Chow and Cusum tests

The estimated beta coefficients shows that investments in ESG companies listed on the Polish capital market were on average perceived as less risky (beta < 1 for all 57 companies) than in the diversified market portfolio, even if we consider a COVID-19 period, inflation, and war crisis.

Results of the Chow stability tests are presented in the graphs in Figure 1 (57 panels), Cusum in Figure 2 (57 panels). Values of the F test statistics are presented in Table 2.

In 53 out of 57 companies (Table 2), the value of the F Chow statistic was high and very unlikely (right tail of the distribution), so at the significance level of $\alpha = 0.05$, it is reasonable to reject the null hypothesis (regression stability) in favor of the alternative hypothesis about instability of the parameters of the regression function. For ABS, GTC, Mabion and Forte (FTE) the Chow test did not allow to reject the hypothesis about the stability of the beta coefficient at the significance level of $\alpha = 0.05$. The mean p-value for the companies was 0.153 over a variation range of 0.053 (Mabion) to 0.233 (ABS). The obtained value of the F statistic allows the adoption of the null hypothesis about the stability of the beta coefficient for the companies.

The Cusum test showed different results. Rejection of the null hypothesis of the stability of the regression coefficients at the significance level $\alpha = 0.05$ as a support of the alternative hypothesis of the instability of the coefficients is justified only in the case of GTN (p-value = 0.003) and PKN Orlen (p-value = 0.018). For the rest of companies, the average p-value was 0.507 with a volatility range from 0.054 for Alior Bank to 0.996 for Kernel. For 55 out of 57 companies the obtained value of the F statistic allows the adoption of the null hypothesis about the stability of the beta coefficient.

Only for 4 companies (ABS, GTC, Mabion and Forte) were there no grounds to reject the null hypothesis of stability according to both tests. In the case of 2 companies (GTN and PKN Orlen) both tests require the rejection of the null hypothesis in favor of the alternative hypothesis of instability of beta parameters. For 51 out of 57 companies Chow and Cusum tests showed divergent results. It may indicate a low power of the Cusum test (low probability of rejecting a false null hypothesis) or too short series of data from periods of hypothetical beta stability (before 2020). Thus, we decided to test beta stability using the rolling window regression method.

4.2. Rolling-window regression method and analysis of beta stationarity

Beta time series charts for individual companies estimated by the rolling window regression method and the average beta level are presented in Figure 3 (57 panels) and Table 4.

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For all of the analyzed companies we reject the hypothesis of beta stationarity over time. The p-value was on average $6.76 * 10^{-95}$ with a range from $3.01 * 10^{-141}$ (PKO) to $3.85 * 10^{-93}$ (MAB).

The outstandingly low risk of making an error of the first type is due to the small number of runs in the studied series. The expected value of the number of series with random deviations from the mean is 331 on average. Meanwhile, the observed (actual) number of series is 36 on average, which is an order of magnitude smaller than the expected number with randomness. In other words, beta values return to the mean too rarely and deviate from it for too long. This can indicate the presence of a deterministic trend and/or autocorrelation without a unit root and/or a stochastic trend (unit root).

CONCLUSIONS

Beta coefficients for the rates of return of most of the 57 ESG companies in the years 2019–2022 are not stable in short term.

The following statistical evidence supports instability:

1. 93% (53 companies out of 57) short-term instability cases confirmed by the Chow test.
2. 100% short-term instability cases confirmed by the Wald-Wolfowitz runs test of randomness around the mean, at the significance level $\alpha = 0.05$, the beta coefficient is unstable in the short term of 2019–2022.

The Cusum test showed different results as the only. Rejection of the null hypothesis of the regression coefficients stability at the significance level $\alpha = 0.05$ as a support of the alternative hypothesis of the instability of the coefficients was justified only in two of the analyzed ESG companies.

Of course, it cannot be ruled out (hypothesis) that the beta coefficient for companies listed in the WIG-ESG index is/will be stable over longer periods of time. Narrowing down the study to the years 2019–2022 was because the index itself has been listed since 2019.

The considered tests argue for the instability of the time series of beta coefficients in most of the companies tested.

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Table 1
Beta stability in academic research

Stability of beta	Author	Market	Period of analysis	Number of stocks	Method of testing
Stable beta	Fabozzi and Francis (1977)	NYSE	January 1966 to December 1971	700	Dummy as variable
	Fisher and Kamin (1985)	NYSE	1926–1979	All listed	cross-sectional regression, variance analysis, Kalman Filter
	Faff (2001)	Australian Stock Exchange	1974 to 1995 (monthly return)	24 industry portfolio	Dummy variable, Regression analysis
	Sromon, Das (2008)	Indian Stock Exchange NSE Nifty	February 1999 to September 2007	39 stocks	Time as variable Dummy as variable
	George and Bainy (2012)	Indian Stock Exchange BSE100 Index	1996–2009	169 stocks	Time as variable Dummy as variable
	Harish and Mallikarjunappa (2019)	Indian Stock Exchange S&P BSE Sensex companies	2000–2014	30 stocks	Chow test, multiple breaking point test, CUSUM test
Instable beta	Blume (1971)	NYSE	January 1926 to June 1968	All listed	Regression analysis Analysis of beta correlations
	Sunder (1980)	NYSE	1926–1975	127 stocks	Variance analysis
	Bos, Newbold (1984)				Kalman Filter
	Russell, Impson and Imre (1994)	NYSE	500, 250, 200, 125, and 100 trading days	2,497 stocks	Variance analysis
	Braun et al. (1995)	NYSE	July 1926–December 1990	CRSP NYSE	Rolling windows, GARCH
	Brooks et al. (1998)	Singapore Stock Exchange	1986 to 1993	247 stocks	OLS regression estimates
	Faff, Hillier, Hillier (2003)	UK	1st January, 1969 to 30th April, 1998	32 industry sector portfolios	GARCH model, rolling window
	Shah, Moonis (2003)	India's Bombay Stock Exchange	1st May 1996 to 30th March 2000	50 stocks	GARCH, Kalman Filter
	Irala (2007)	Indian Stock Market BSE	April 1994 to March2006	660 stocks	statistical significance of the CAPM model (t-Student test)
	Sarma and Sarmah (2008)	Indian Stock Market BSE	December 2001 to November2006	5 stocks	Chow test
	Razvan et al. (2009)	Bukarest Stock Exchange	20th January to 20th July 2009	10 stocks	statistical significance of the CAPM model(t-Student test)
Instable beta	Javid & Ahmad (2011)	Pakistan Stock Market Karachi Stock Exchange	1993–2007	50 stocks	Dummy as variable
	Mazowina (2013)	Zimbabwean Stock Market	February 2009 to 31 December 2012	66 stocks	Chow test

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Table 1 – continued

Stability of beta	Author	Market	Period of analysis	Number of stocks	Method of testing
Instable beta	Celik (2013)	Turkish Stock Market	03.01.2005–31.12.2009	Istanbul Stock Exchange (ISE) sector indices and the ISE-30 All Share Index	Rolling windows
	Simon et al. (2012)	Brazilian stock exchange (BM&FBOVESPA)	2002 and 2011	All stocks	Analysis of beta correlations
	Wijethunga A.W.G. Dayaratne D.A. (2015)	Colombo stock exchange	2005–2013	26 stocks	Rolling windows
	Ye (2017)	China's Stock Exchange Shanghai Stock Exchange and the Shenzhen Stock Exchange	Before 2008	208 stocks	Dummy as a variable
	Gupta (2020)	Bombay Stock Exchange	January 2006 to January 2018	11 sectors	Chow Test Dummy Variable
	Levy 1971	NYSE	1962–1970	500 stocks	Correlation analysis
	Baesel (1974)	NYSE	January 1950 to 1967	160 stocks	Chow-test
	Witkowska (2008)	Warsaw Stock Exchange	2000–2006	8 stocks	t-Student test
Contractionary results	Singh (2008)	Bombay Stock Exchange BSE	1991–2002	158 stocks	Regression analysis
	Ray (2010)	Bombay Stock Exchange BSE100	January 1996 to December 2009	100 stocks	Time as variable Dummy as variable Chow test
	Deb and Mistra (2011)	Indian Stock Market BSE	1996 to 2010	158 stocks	Dummy as variable
	Terceño et al. (2011)	Hong Kong Stock Exchange	01.01.2005 and 06.31.2009	All stocks	OLS regression estimates
	Dubey (2014)	National Stock Exchange of India	June 15, 2001 to March 31, 2010 t	25 stocks	OLS regression estimates, wavelet filters
	Dębski et al. (2016)	Warsaw Stock Exchange	2005–2013	134 stocks	Chow test
	Dębski et al. (2017)	Warsaw, Frankfurt and Paris Stock Exchange	2005–2015	37 stocks 28 stocks 36 stocks	t-Student test, Chow test
	Ye (2017)	Shenzhen Stock Exchange which	January 2008 to December 2013	208 stocks	t-Student test dummy variable
	Mikołajek-Gocejna (2021)	emerging markets	2005–2021	25 emerging markets indexes	Chow-test Cusum- test Rolling-window

Source: Own collaboration.

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Table 2
Chow and Cusum Test

Walor	Chow.date	Chow.sup.F	Chow.sup.F.pvalue	Cusum.S	Cusum.S.pvalue
IIB	382	15.08185528	0.019973285	0.647821148	0.327845666
DNP	286	30.52889011	1.08E-05	0.491109827	0.646765335
LTS	368	49.32294455	5.43E-10	0.791890783	0.145420383
MIL	378	113.9468673	0	0.734618166	0.204967266
ING	283	77.83149039	1.11E-16	0.534963667	0.548012718
OPL	359	23.97115944	0.00029592	0.409997695	0.824710486
MBK	402	73.54352641	9.90E-16	0.596196782	0.420722389
PGE	365	46.80376457	2.10E-09	0.861671606	0.092397623
CCC	363	88.51940363	0	0.835298751	0.110174712
ABS	129	9.077899211	0.233223383	0.320920321	0.9488163
KRU	129	28.0469548	3.84E-05	0.763552274	0.172902915
ALR	387	155.275043	0	0.9380663	0.053787773
EAT	287	23.49410408	0.000374523	0.708004418	0.238267184
PLW	360	46.23915774	2.84E-09	0.582102012	0.448527549
KTY	438	53.41596002	5.99E-11	0.378141994	0.882660864
BHW	379	114.6822427	0	0.450062631	0.739790413
JSW	381	48.63765734	7.85E-10	0.654193277	0.317402503
GTC	276	9.238391044	0.220097067	0.472763041	0.688630113
CAR	379	39.67107573	9.29E-08	0.534859982	0.548240886
ATT	378	49.23252259	5.70E-10	0.733330974	0.206492187
EUR	377	40.61439883	5.64E-08	0.461204888	0.714829501
BDX	367	22.50915189	0.000607508	0.467051358	0.701606453
ENG	425	19.91837428	0.002127018	0.36816124	0.898235543
KER	608	128.8762717	0	0.232403379	0.965952905
ENA	377	31.76535268	5.74E-06	0.628207876	0.361415019
TPE	359	31.42702312	6.84E-06	0.602723896	0.408187535
FMF	381	39.34008645	1.11E-07	0.511410835	0.600601617
CMR	544	41.60377911	3.34E-08	0.334056038	0.939119414
LCC	377	50.27374761	3.26E-10	0.420610376	0.803214462
WPL	378	55.69859382	1.74E-11	0.454188632	0.730590552
ECH	357	49.21092465	5.77E-10	0.344339182	0.929026721
GPW	316	36.13063626	5.96E-07	0.357433279	0.913294154

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Table 2 – continued

Walor	Chow.date	Chow.sup.F	Chow.sup.F.pvalue	Cusum.S	Cusum.S.pvalue
PKP	373	62.34990402	4.68E-13	0.68056555	0.276606314
VRG	130	16.25462029	0.011777334	0.626268015	0.364851298
CIE	386	32.05733983	4.94E-06	0.48603112	0.658363247
BFT	292	43.38643443	1.30E-08	0.457931241	0.72219986
MAB	367	12.8237853	0.053278843	0.514295356	0.59409052
AMC	371	51.24419302	1.93E-10	0.797223443	0.140659091
FTE	378	11.14458497	0.106267508	0.551025104	0.51311024
LVC	462	27.41579518	5.29E-05	0.500232976	0.625962749
LWB	463	38.86331504	1.42E-07	0.756345023	0.180500265
BRS	352	17.89436681	0.005531515	0.54891527	0.517641872
STP	289	33.46658137	2.38E-06	0.65642274	0.313802547
PXM	358	44.69984055	6.45E-09	0.584578514	0.443570835
GNB	391	5.76518121	2.89E-11	0.430552857	0.782337045
CIG	373	16.75431427	0.009373255	0.599674445	0.414016321
TRK	368	25.21049507	0.000159893	0.566419866	0.480587001
GTN	396	27.7096287	4.56E-05	1.268167143	0.003028214
PKO	368	193.8597071	0	0.648530667	0.32667157
PZU	373	115.5760748	0	0.673814359	0.286678652
PKN	368	82.93981804	0	1.075136042	0.018119942
CDR	349	33.33283174	2.56E-06	0.362188842	0.90684615
LPP	379	91.49276578	0	0.647212353	0.328855338
SPL	373	103.3113881	0	0.641553246	0.338340222
KGH	368	83.82726214	0	0.659523834	0.308841615
CPS	347	35.68615614	7.52E-07	0.44308895	0.755198423
PGN	377	32.00106162	5.09E-06	0.85915987	0.093981604

Source: own estimation.

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Table 3
Linear trend regression

Company	a	a.se	a.t	a.pvalue	b	b.se	b.t	b.pvalue	betats.se	betats.R2	betats.df	betats.F	DW	pv_of_DW
IIB	–0.10934	0.039239	–2.78645	0.0055	0.000773	0.000102	7.591129	1.08E-13	0.506131	0.079744	665	57.62525	0.135391	3.45E-129
DNP	–0.22538	0.038838	–5.80316	1.01E-08	0.00126	0.000101	12.50971	2.14E-32	0.50096	0.190498	665	156.4929	0.142529	3.21E-128
LTS	–0.53456	0.039991	–13.3672	2.88E-36	0.002805	0.000104	27.04442	3.18E-109	0.515826	0.523776	665	731.4004	0.110361	1.30E-132
MIL	–0.61177	0.063329	–9.66015	9.51E-21	0.003704	0.000164	22.54674	4.72E-84	0.816862	0.433249	665	508.3553	0.080081	8.15E-137
ING	–0.5202	0.036921	–14.0898	1.22E-39	0.002388	9.58E-05	24.93442	2.15E-97	0.476226	0.483184	665	621.7251	0.085091	4.08E-136
OPL	–0.36444	0.024154	–15.0882	1.87E-44	0.001569	6.27E-05	25.03753	5.67E-98	0.311554	0.485245	665	626.8779	0.210043	3.02E-119
MBK	–0.43055	0.059367	–7.25233	1.14E-12	0.002994	0.000154	19.44058	5.34E-67	0.765749	0.362377	665	377.936	0.083031	2.11E-136
PGE	–0.65123	0.053985	–12.0632	1.92E-30	0.003359	0.00014	23.99133	4.11E-92	0.696327	0.463962	665	575.584	0.128772	4.33E-130
CCC	–1.11246	0.064932	–17.1328	8.90E-55	0.004509	0.000168	26.77214	1.07E-107	0.837532	0.518725	665	716.7475	0.069055	2.31E-138
ABS	–0.23277	0.028858	–8.06603	3.39E-15	0.000832	7.49E-05	11.11845	1.86E-26	0.372232	0.156755	665	123.62	0.180573	4.01E-123
KRU	–0.56067	0.047961	–11.6901	7.64E-29	0.002366	0.000124	19.02112	9.65E-65	0.618633	0.352359	665	361.8029	0.179437	2.84E-123
ALR	–0.81117	0.057181	–14.1862	4.24E-40	0.004168	0.000148	28.09997	3.96E-115	0.737552	0.542832	665	789.6083	0.078982	5.72E-137
EAT	–0.44618	0.039715	–11.2346	6.18E-27	0.002164	0.000103	21.00701	1.55E-75	0.512267	0.398894	665	441.2944	0.141319	2.20E-128
PLW	–0.23091	0.043952	–5.25359	2.01E-07	0.001608	0.000114	14.1014	1.07E-39	0.566922	0.23019	665	198.8495	0.140928	1.95E-128
KTY	–0.37574	0.036678	–10.2442	5.75E-23	0.001764	9.51E-05	18.54246	3.49E-62	0.473099	0.340816	665	343.8229	0.124189	1.03E-130
BHW	–0.439	0.038992	–11.2588	4.91E-27	0.00227	0.000101	22.44798	1.67E-83	0.502938	0.431095	665	503.9119	0.081999	1.51E-136
JSW	–0.85591	0.098975	–8.64777	3.92E-17	0.004553	0.000257	17.73435	6.51E-58	1.276644	0.321087	665	314.5073	0.085206	4.24E-136
GTC	–0.11969	0.036165	–3.30943	0.000985	0.000739	9.38E-05	7.879995	1.34E-14	0.466479	0.085401	665	62.09431	0.07637	2.46E-137
CAR	–0.35157	0.041364	–8.49942	1.25E-16	0.001759	0.000107	16.39514	5.47E-51	0.533544	0.287857	665	268.8006	0.127661	3.06E-130
ATT	–0.49452	0.056158	–8.80581	1.12E-17	0.002243	0.000146	15.40004	5.43E-46	0.724366	0.262881	665	237.1612	0.096747	1.71E-134
EUR	–0.12467	0.041921	–2.97386	0.003047	0.001648	0.000109	15.15995	8.30E-45	0.540721	0.256837	665	229.8241	0.117675	1.32E-131
BDX	–0.50655	0.039252	–12.9052	3.66E-34	0.001714	0.000102	16.82985	3.26E-53	0.506293	0.298704	665	283.244	0.120805	3.53E-131
ENG	–0.14249	0.018359	–7.76142	3.18E-14	0.000711	4.76E-05	14.93518	1.05E-43	0.236805	0.251176	665	223.0597	0.173244	4.26E-124
KER	–0.55356	0.047807	–11.579	2.25E-28	0.002423	0.000124	19.53706	1.61E-67	0.616645	0.364668	665	381.6967	0.126997	2.48E-130
ENA	–0.42754	0.048882	–8.74644	1.80E-17	0.002521	0.000127	19.88542	2.08E-69	0.630513	0.372896	665	395.4299	0.109046	8.56E-133
TPE	–0.41522	0.068442	–6.06673	2.19E-09	0.002383	0.000178	13.42528	1.55E-36	0.882804	0.21324	665	180.2382	0.093357	5.78E-135
FMF	–0.29596	0.057144	–5.17916	2.96E-07	0.00258	0.000148	17.40339	3.50E-56	0.737075	0.31293	665	302.8781	0.108318	6.80E-133
CMR	–0.44025	0.038053	–11.5692	2.48E-28	0.001823	9.87E-05	18.47179	8.29E-62	0.490838	0.339102	665	341.2069	0.12469	1.20E-130
LCC	–0.23598	0.046138	–5.11457	4.12E-07	0.001223	0.00012	10.22093	7.08E-23	0.595123	0.135766	665	104.4673	0.081938	1.48E-136

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Table 3 – continued

Company	a	a.se	a.t	a.pvalue	b	b.se	b.t	b.pvalue	betats.se	betats.R2	betats.df	betats.F	DW	pv_of_DW
WPL	–0.57783	0.043125	–13.399	2.05E-36	0.00245	0.000112	21.89962	1.84E-80	0.556254	0.419008	665	479.5935	0.121088	3.86E-131
ECH	–0.23806	0.034685	–6.86346	1.54E-11	0.001299	9.00E-05	14.4402	2.60E-41	0.447389	0.238712	665	208.5193	0.104903	2.30E-133
GPW	–0.35757	0.024844	–14.3925	4.40E-41	0.001215	6.44E-05	18.84996	7.97E-64	0.320456	0.348244	665	355.3211	0.149209	2.56E-127
PKP	–0.38135	0.046959	–8.12093	2.25E-15	0.002668	0.000122	21.90694	1.68E-80	0.60571	0.41917	665	479.9141	0.101601	8.03E-134
VRG	–0.25261	0.032935	–7.67011	6.13E-14	0.001151	8.54E-05	13.46807	9.86E-37	0.424815	0.214309	665	181.3888	0.148726	2.21E-127
CIE	–0.52295	0.046032	–11.3605	1.86E-27	0.002109	0.000119	17.65967	1.60E-57	0.593754	0.31925	665	311.8638	0.115442	6.50E-132
BFT	–0.24151	0.038776	–6.22829	8.36E-10	0.001276	0.000101	12.68881	3.42E-33	0.500155	0.194921	665	161.006	0.121956	5.08E-131
MAB	–1.06978	0.096004	–11.1431	1.47E-26	0.004043	0.000249	16.23379	3.61E-50	1.238328	0.283819	665	263.5359	0.16929	1.27E-124
AMC	–0.47292	0.036745	–12.8702	5.26E-34	0.001827	9.53E-05	19.17004	1.53E-65	0.473962	0.355926	665	367.4904	0.13	6.37E-130
FTE	–0.21564	0.057772	–3.73259	0.000206	0.001419	0.00015	9.466455	4.93E-20	0.745177	0.118754	665	89.61378	0.105587	2.86E-133
LVC	–0.48436	0.042387	–11.427	9.79E-28	0.001795	0.00011	16.32939	1.18E-50	0.546735	0.286212	665	266.649	0.141167	2.10E-128
LWB	–0.61397	0.072098	–8.51577	1.10E-16	0.003455	0.000187	18.47396	8.07E-62	0.929971	0.339155	665	341.2873	0.062795	3.03E-139
BRS	–0.23199	0.026468	–8.76469	1.55E-17	0.00121	6.87E-05	17.62468	2.45E-57	0.341406	0.318389	665	310.6292	0.247096	1.84E-114
STP	–0.40903	0.053849	–7.59578	1.04E-13	0.00272	0.00014	19.47433	3.51E-67	0.694585	0.363179	665	379.2497	0.120484	3.19E-131
PXM	–0.39289	0.063313	–6.20561	9.58E-10	0.002432	0.000164	14.80953	4.28E-43	0.816648	0.248012	665	219.3222	0.090427	2.26E-135
GNB	–1.05458	0.082906	–12.7201	2.48E-33	0.004577	0.000215	21.28337	4.67E-77	1.069378	0.405178	665	452.9817	0.141582	2.39E-128
CIG	–0.27678	0.05756	–4.80856	1.88E-06	0.0014	0.000149	9.376604	1.05E-19	0.742444	0.116773	665	87.92071	0.159514	6.24E-126
TRK	–0.50191	0.051612	–9.72467	5.47E-21	0.00203	0.000134	15.16011	8.29E-45	0.665722	0.256841	665	229.829	0.159826	6.87E-126
GTN	–0.85955	0.057786	–14.8747	2.06E-43	0.003045	0.00015	20.3124	9.88E-72	0.745361	0.382884	665	412.5934	0.153256	8.99E-127
PKO	–0.59322	0.044165	–13.4321	1.44E-36	0.003524	0.000115	30.76182	6.43E-130	0.569665	0.587287	665	946.2896	0.060767	1.57E-139
PZU	–0.64264	0.02658	–24.1776	3.73E-93	0.002849	6.89E-05	41.32759	7.33E-186	0.342844	0.71976	665	1707.97	0.096994	1.85E-134
PKN	–0.54446	0.039955	–13.6268	1.81E-37	0.002964	0.000104	28.5984	6.54E-118	0.515367	0.551545	665	817.8687	0.094282	7.77E-135
CDR	–0.2245	0.051361	–4.37095	1.44E-05	0.001865	0.000133	13.99659	3.35E-39	0.662494	0.227557	665	195.9045	0.123868	9.27E-131
LPP	–0.56185	0.055261	–10.1672	1.14E-22	0.003434	0.000143	23.95832	6.29E-92	0.712791	0.463277	665	574.0013	0.088847	1.36E-135
SPL	–0.60266	0.04691	–12.8471	6.69E-34	0.003295	0.000122	27.08268	1.94E-109	0.605076	0.524481	665	733.4715	0.072525	7.11E-138
KGH	–0.83565	0.05662	–14.7588	7.54E-43	0.003914	0.000147	26.64717	5.35E-107	0.730327	0.516389	665	710.0716	0.065246	6.72E-139
CPS	–0.41877	0.033395	–12.5398	1.57E-32	0.001524	8.66E-05	17.59572	3.47E-57	0.430751	0.317675	665	309.6093	0.101721	8.34E-134
PGN	–0.34212	0.030527	–11.2072	8.01E-27	0.001952	7.92E-05	24.64875	8.57E-96	0.393756	0.477432	665	607.561	0.165295	3.71E-125

Source: Own estimation.

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Table 4
Run tests

Company	R	n1	n2	E(R)	VAR(R)	z-value	p-value
IIB	49	348	319	333.8695652	165.8697866	–22.1188	2.08E-108
DNP	46	296	371	330.2833583	162.3096797	–22.3141	2.70E-110
LTS	27	310	357	332.844078	164.8478198	–23.8209	2.03E-125
MIL	12	312	355	333.113943	165.1164523	–24.9899	7.88E-138
ING	22	322	345	334.1034483	166.103309	–24.2164	1.50E-129
OPL	52	317	350	333.6836582	165.6842834	–21.8837	3.71E-106
MBK	32	301	366	331.3328336	163.3475197	–23.4206	2.64E-121
PGE	30	361	306	332.2323838	164.2397293	–23.5832	5.74E-123
CCC	14	326	341	334.3313343	166.3310017	–24.8378	3.50E-136
ABS	48	368	299	330.9310345	162.9497845	–22.1643	7.60E-109
KRU	54	302	365	331.5247376	163.5376538	–21.7017	1.98E-104
ALR	22	302	365	331.5247376	163.5376538	–24.204	2.02E-129
EAT	40	336	331	334.4812594	166.4808845	–22.8231	2.70E-115
PLW	56	334	333	334.4992504	166.498875	–21.5833	2.58E-103
KTY	23	325	342	334.2833583	166.2830534	–24.1397	9.58E-129
BHW	27	289	378	328.5622189	160.6147823	–23.7949	3.77E-125
JSW	35	331	336	334.4812594	166.4808845	–23.2107	3.55E-119
GTC	56	331	336	334.4812594	166.4808845	–21.5831	2.59E-103
CAR	40	306	361	332.2323838	164.2397293	–22.8029	4.30E-115
ATT	28	309	358	332.7001499	164.7046386	–23.7421	1.32E-124
EUR	48	296	371	330.2833583	162.3096797	–22.1571	8.91E-109
BDX	41	326	341	334.3313343	166.3310017	–22.7443	1.64E-114
ENG	46	359	308	332.5502249	164.5555577	–22.338	1.58E-110
KER	38	231	436	302.9970015	136.4867746	–22.6827	6.63E-114
ENA	25	322	345	334.1034483	166.103309	–23.9836	4.12E-127
TPE	46	341	326	334.3313343	166.3310017	–22.3566	1.04E-110
FMF	36	326	341	334.3313343	166.3310017	–23.132	2.21E-118
CMR	48	290	377	328.826087	160.8740499	–22.1409	1.28E-108
LCC	48	306	361	332.2323838	164.2397293	–22.1786	5.53E-109
WPL	34	322	345	334.1034483	166.103309	–23.2853	6.25E-120
ECH	39	308	359	332.5502249	164.5555577	–22.8837	6.75E-116
GPW	48	375	292	329.3358321	161.3754997	–22.1466	1.13E-108
PKP	34	314	353	333.3598201	165.3613967	–23.2796	7.13E-120

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Table 4 – continued

Company	R	n1	n2	E(R)	VAR(R)	z-value	p-value
VRG	45	341	326	334.3313343	166.3310017	–22.4341	1.83E-111
CIE	37	297	370	330.5052474	162.528833	–23.0224	2.78E-117
BFT	45	358	309	332.7001499	164.7046386	–22.4175	2.66E-111
MAB	68	272	395	323.1589205	155.3516684	–20.4716	3.85E-93
AMC	45	317	350	333.6836582	165.6842834	–22.4275	2.12E-111
FTE	56	366	301	331.3328336	163.3475197	–21.5428	6.19E-103
LVC	51	309	358	332.7001499	164.7046386	–21.95	8.66E-107
LWB	33	254	413	315.5487256	148.0876157	–23.2185	2.96E-119
BRS	38	279	388	325.5937031	157.7124301	–22.9006	4.59E-116
STP	47	300	367	331.1349325	163.1515597	–22.2448	1.27E-109
PXM	15	306	361	332.2323838	164.2397293	–24.7536	2.84E-135
GNB	37	262	405	319.1709145	151.5233633	–22.9231	2.74E-116
CIG	51	350	317	333.6836582	165.6842834	–21.9614	6.74E-107
TRK	34	353	314	333.3598201	165.3613967	–23.2796	7.13E-120
GTN	34	326	341	334.3313343	166.3310017	–23.287	6.00E-120
PKO	8	319	348	333.8695652	165.8697866	–25.3023	3.01E-141
PZU	10	305	362	332.0644678	164.072999	–25.1434	1.67E-139
PKN	14	315	352	333.4737631	165.4749691	–24.8353	3.73E-136
CDR	36	324	343	334.2293853	166.2291199	–23.1311	2.25E-118
LPP	10	297	370	330.5052474	162.528833	–25.1403	1.81E-139
SPL	16	330	337	334.4632684	166.462895	–24.6832	1.62E-134
KGH	8	301	366	331.3328336	163.3475197	–25.2984	3.32E-141
CPS	40	329	338	334.4392804	166.4389105	–22.8228	2.72E-115
PGN	34	312	355	333.113943	165.1164523	–23.2778	7.45E-120
Mean	36.07018	316.9123	350.0877	331.3998527	163.4544162	–23.1002	6.76E-95
Min	8	231	292	302.9970015	136.4867746	–25.3023	3E-141
Max	68	375	436	334.4992504	166.498875	–20.4716	3.85E-93

Source: Own estimation

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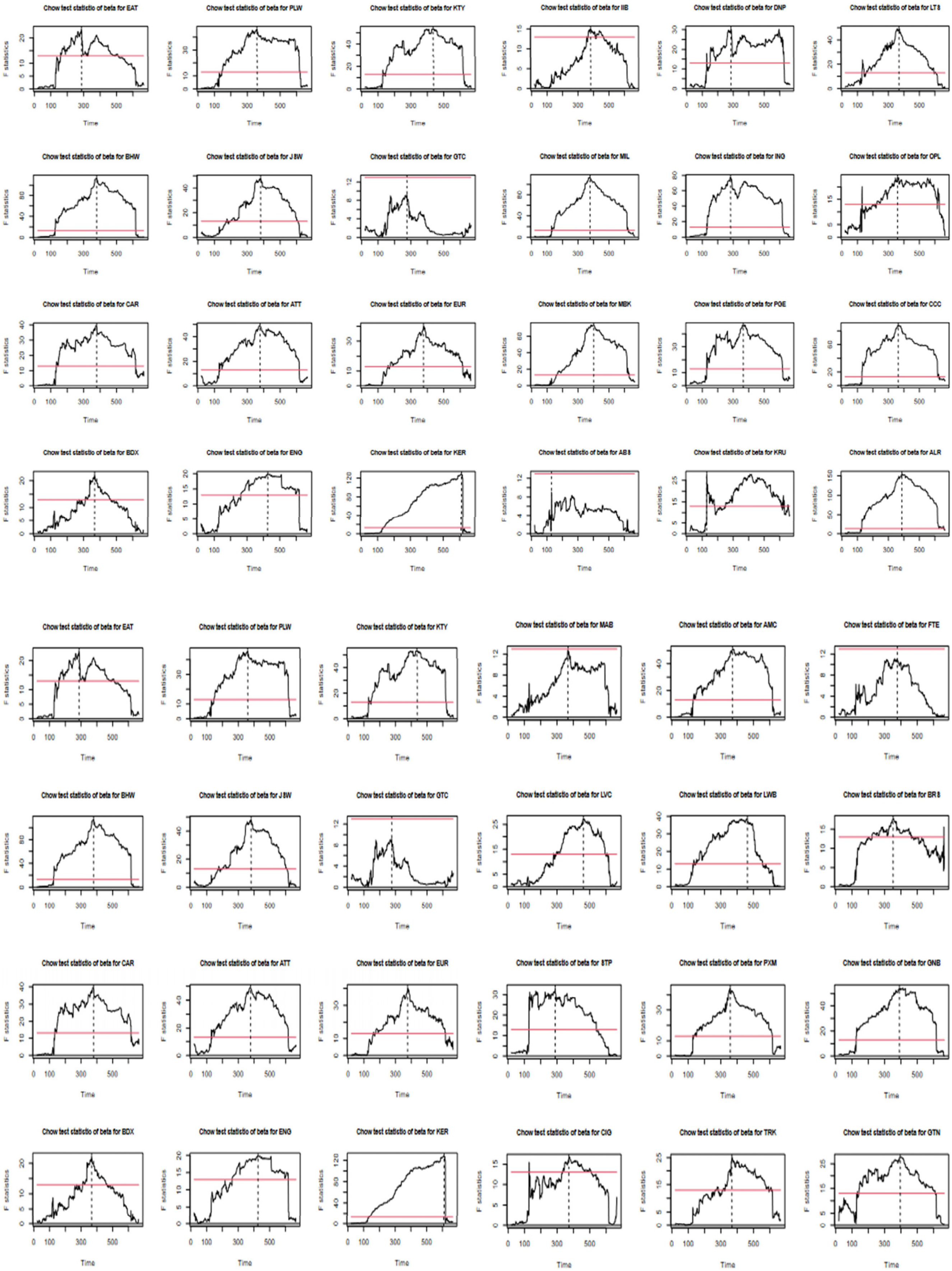
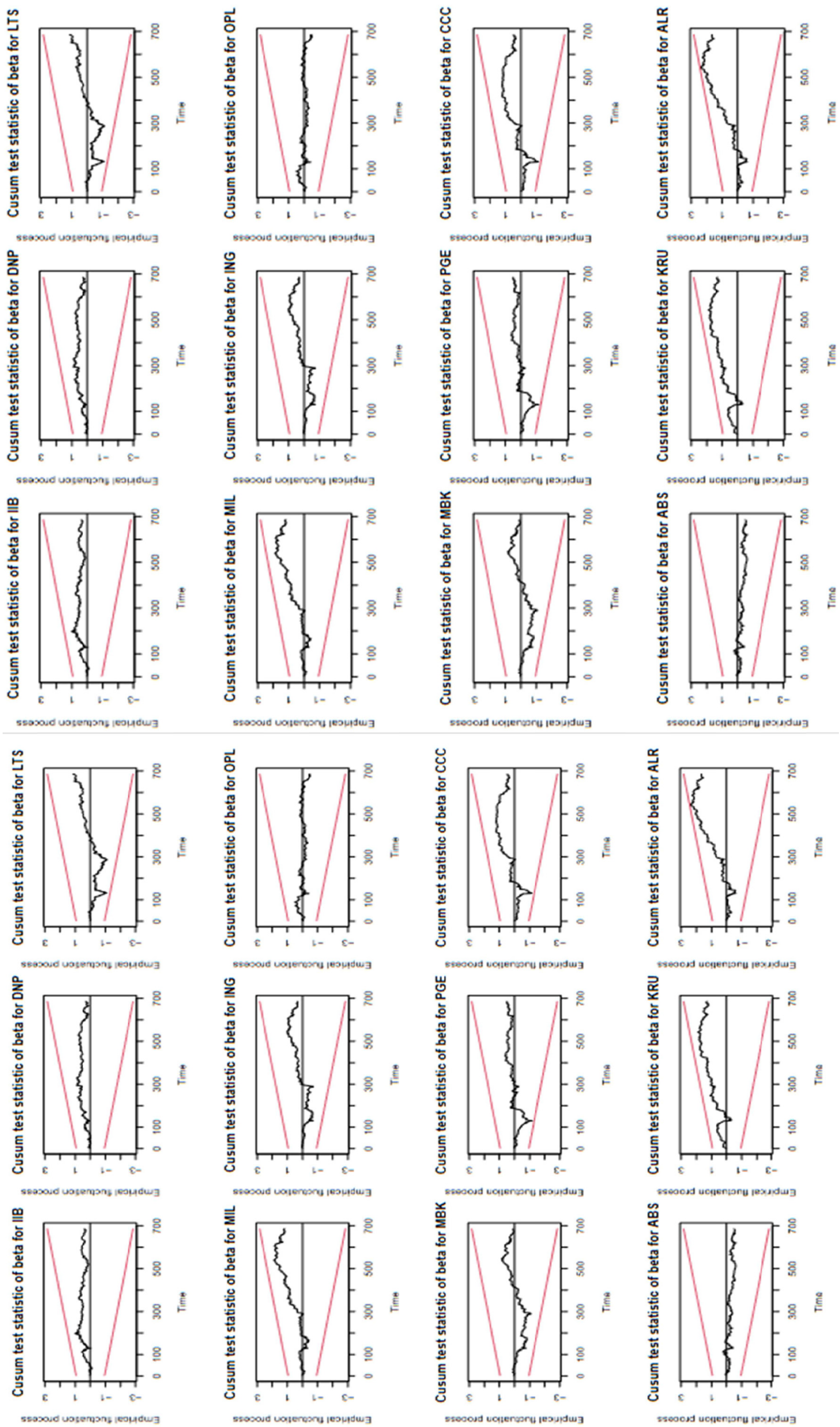
Figure 1
Chow test



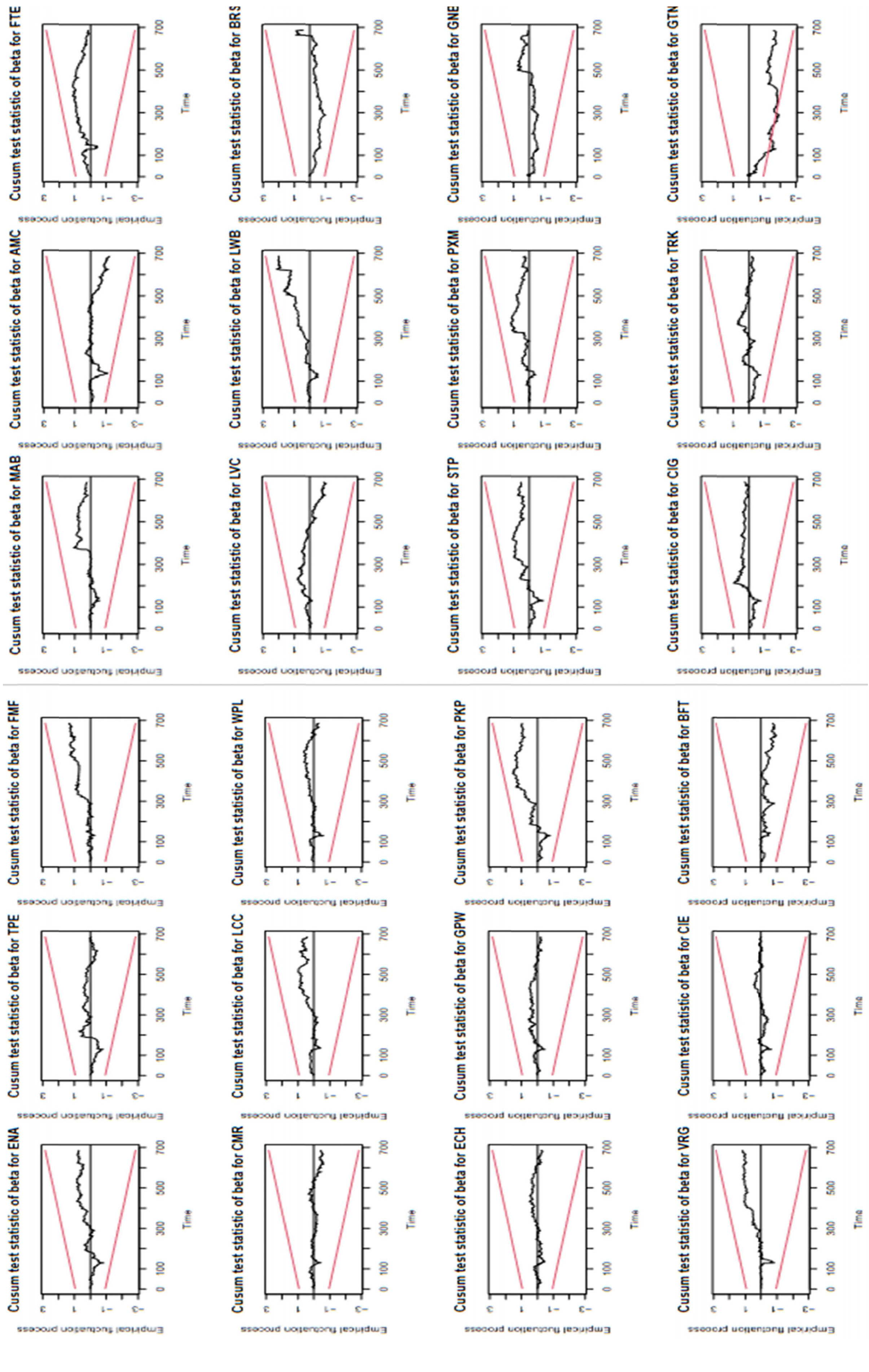
DOI: 10.7172/2353-6845.jbfe.2023.2.1

Figure 2
Cusum test



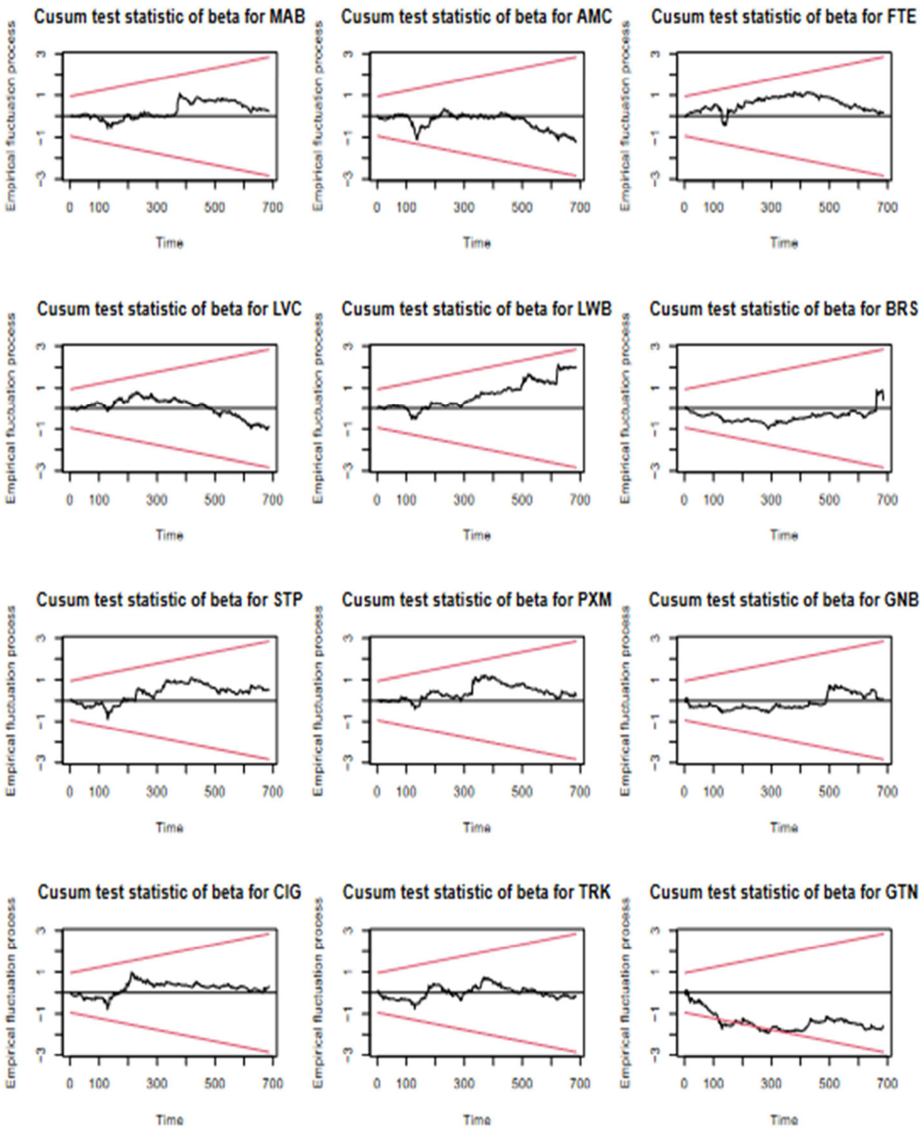
DOI: 10.7172/2353-6845.jbfe.2023.2.1

Figure 2 – continued



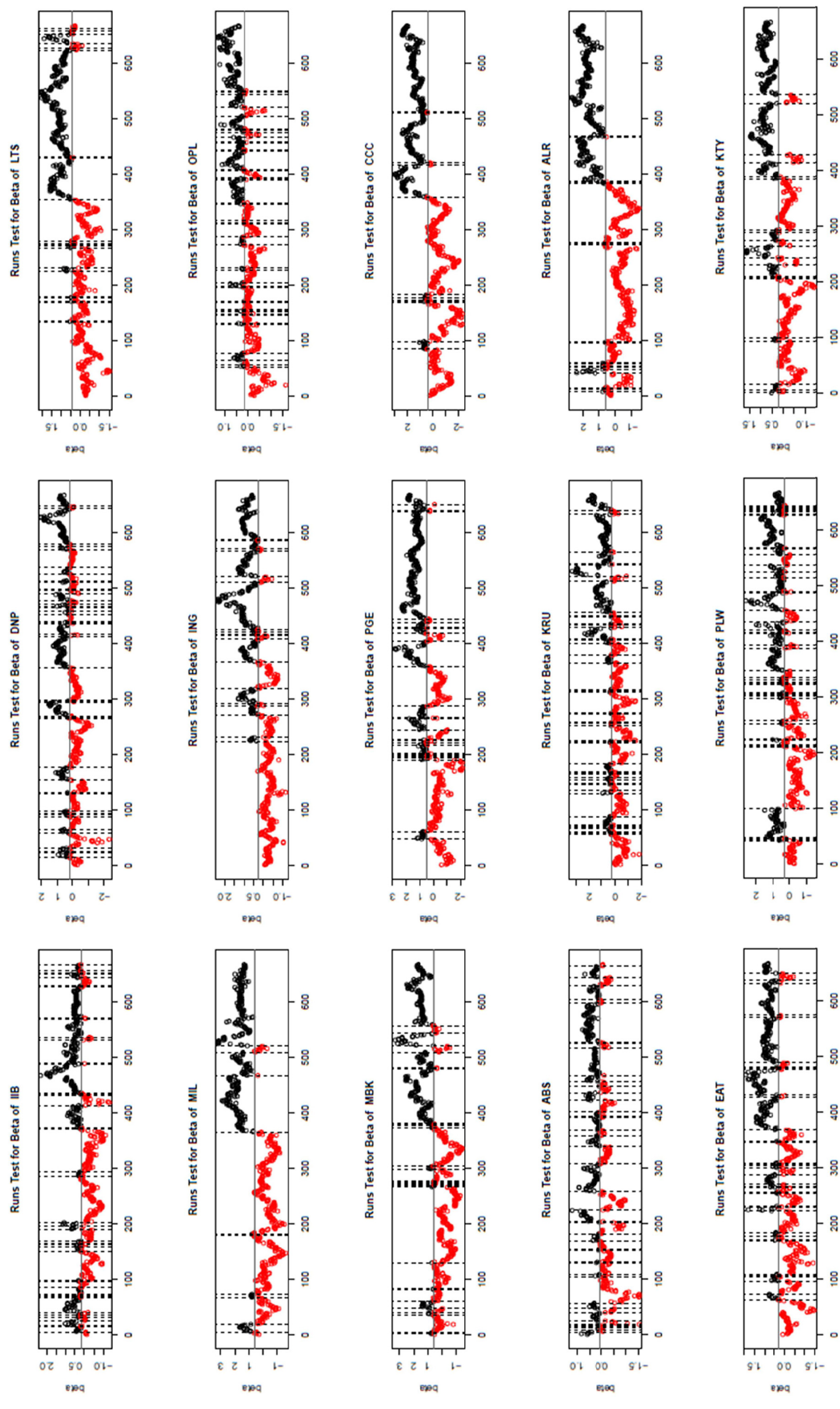
DOI: 10.7172/2353-6845.jbfe.2023.2.1

Figure 2 – continued



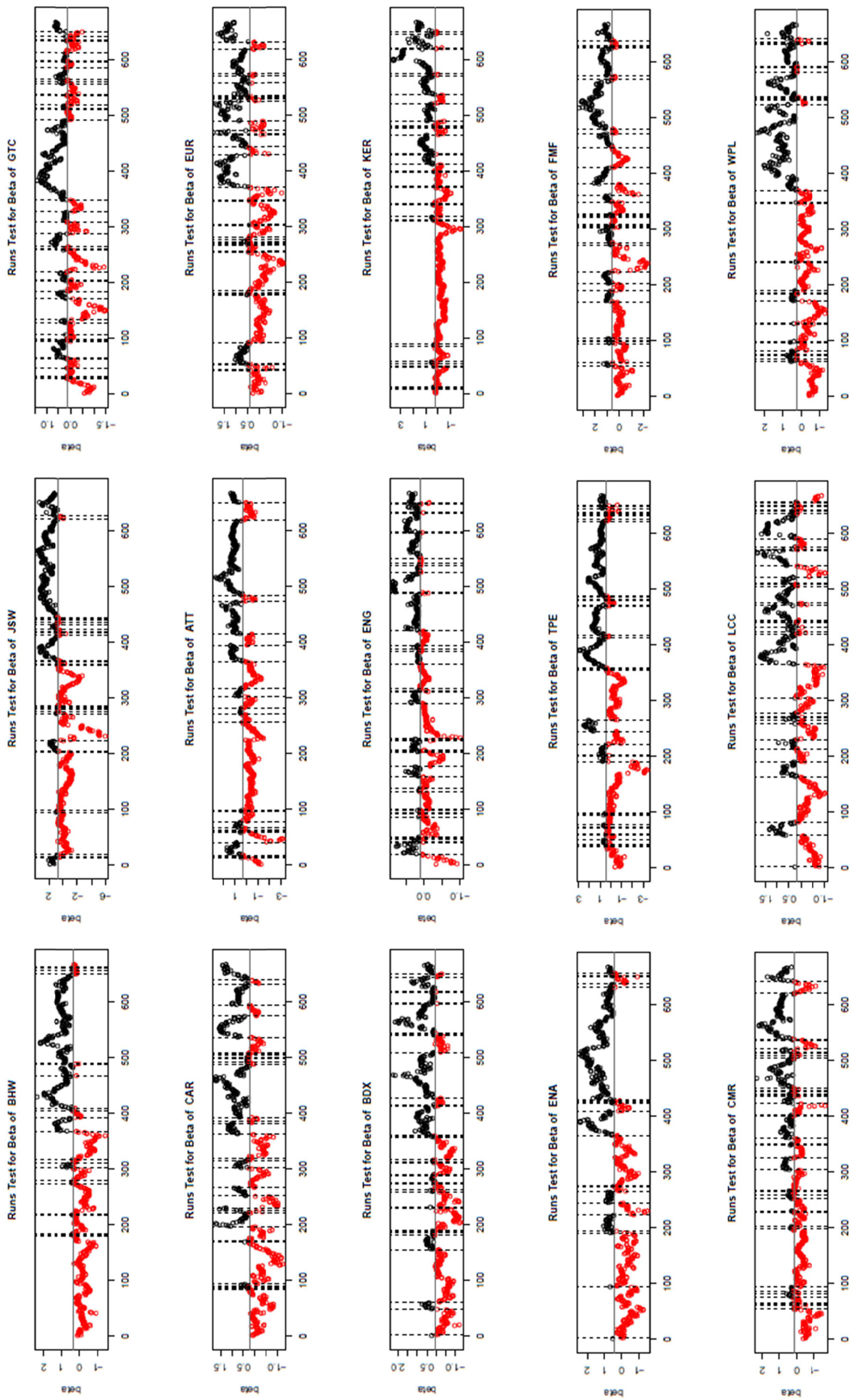
DOI: 10.7172/2353-6845.jbfe.2023.2.1

Figure 3
Rolling window run test



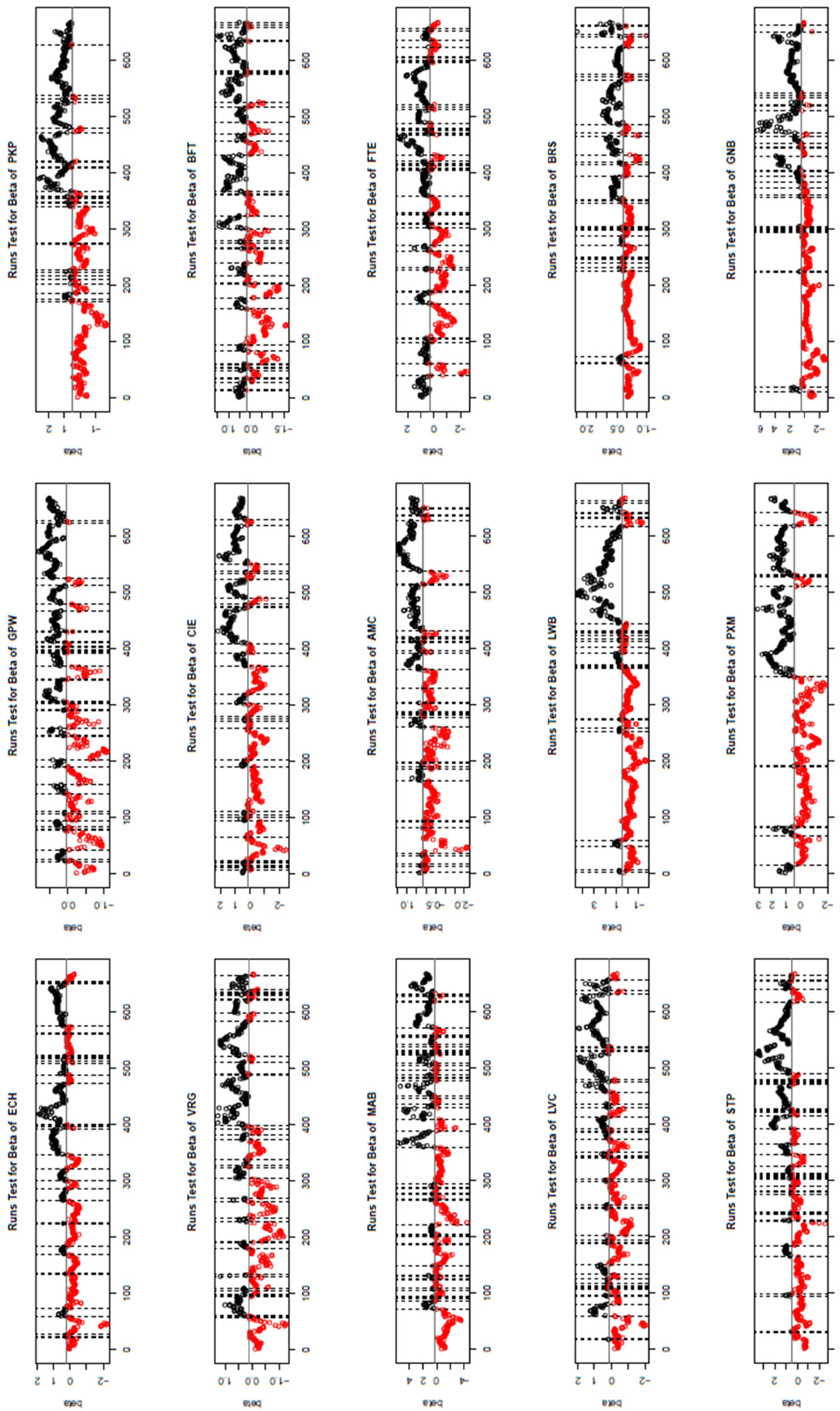
DOI: 10.7172/2353-6845.jbfe.2023.2.1

Figure 3 – continued



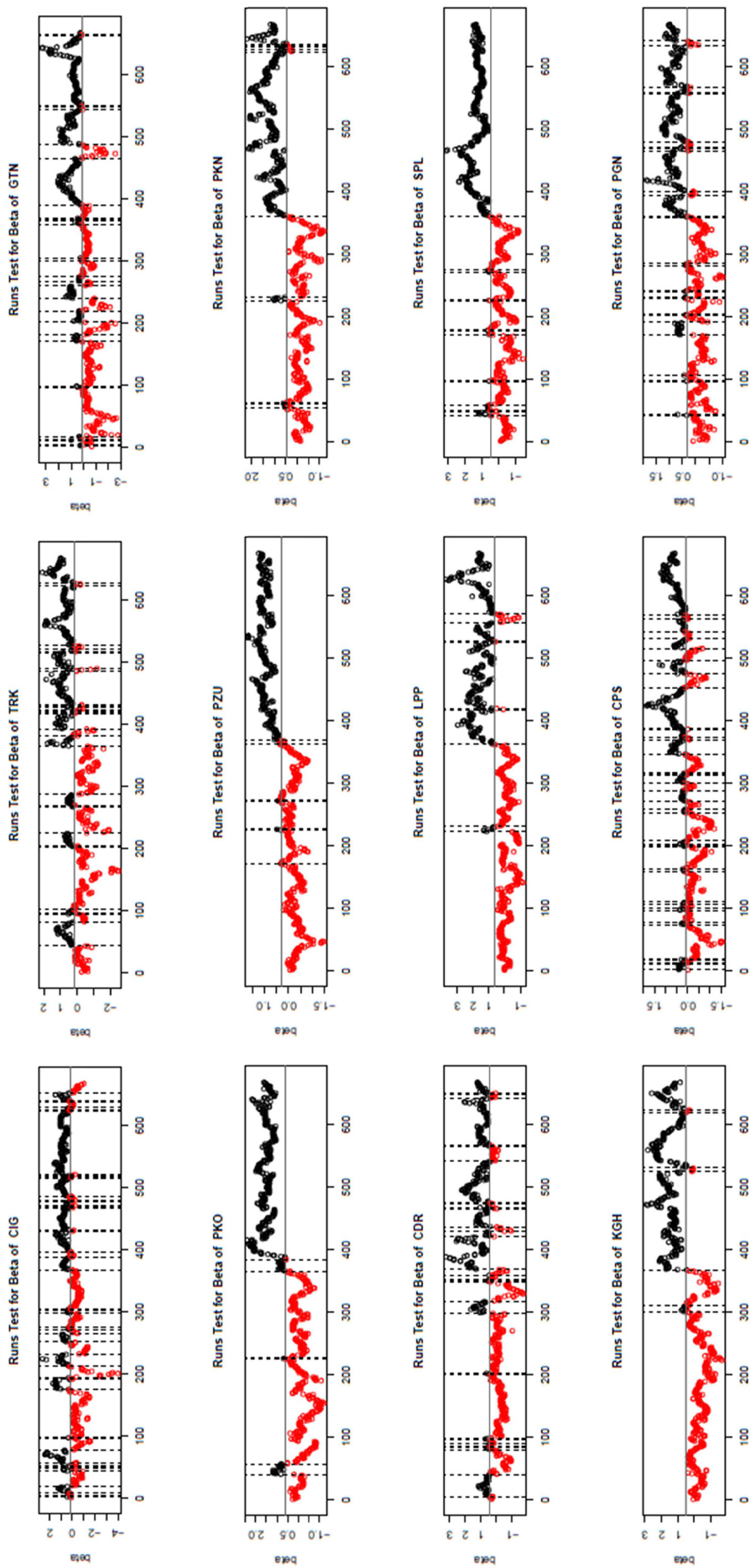
DOI: 10.7172/2353-6845.jbfe.2023.2.1

Figure 3 – continued



DOI: 10.7172/2353-6845.jbfe.2023.2.1

Figure 3 – continued



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Optimization of Banks' Value – Practical Challenges

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ABSTRACT

The aim of the article is to verify the conceptual model of integrated optimization of a bank's value, which enables the integration of the risk management process with business processes while maintaining compromise between the safety (stability) of a bank's operations and striving to maximize its value.

The model is an attempt at a comprehensive solution to such dilemmas as shaping a bank's value ex ante, not ex post. Verification of the model has shown that the model works in accordance with the adopted assumptions and leads to the achievement of the basic goal for which it was constructed. In practice, it means the possibility of ensuring a compromise between the safety and effectiveness of a bank's operations, which, in the context of ongoing changes in its environment, allows for a long-term competitive advantage.

JEL classification: C61, G21, G32

Keywords: Bank management model, Integrated value optimization, Stochastic simulations, Decision support systems.

1. INTRODUCTION

The progressing globalisation and integration of financial markets as well as the recently observed tendency to regulate the phenomena have a number of consequences determining activities of banks (Basel Committee on Banking Supervision, 2011). As a result, the quantitative and qualitative changes taking place in the banking sector, and above all the growing competition on the financial services market, put a new light on the problem of managing the effectiveness of banks (Committee on the Global Financial System, 2018). The issue gained special importance

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as a result of the 2008+ financial crisis, the real effects of which and the adopted preventive solutions limited the effectiveness of banks (Oino, 2018). The construction of the new structure of financial supervision is a consequence of the imperfection of institutional supervision in many countries, which did not provide sufficient protection of financial stability to domestic markets in the conditions of crisis (World Bank, 2020). The crisis also revealed the need for greater coordination of supervisory activities at the supranational and global level in order to effectively counter arbitrage between individual countries (The de Larosière Group, 2009) and the contagion effect (Koleśnik, 2021). At the same time, the growing needs of the bank's environment, significant changes in the operating conditions and the dynamic situation on the financial markets make it necessary to formulate new operating strategies and develop new ways of managing banks (Andrle et al., 2017). The observed phenomena on the global financial market pose both an opportunity for banks and a threat to their development. Advantages are manifested in the possibility of diversifying activities in the context of customers and their segments, product groups, area of operation, levels and types of risks taken. On the other hand, the crisis, in addition to the imperfections of the global financial system, also revealed the insufficient quality of bank management methods used (Holland, 2010; Laurens, 2012). The vast majority of banks have so far been assessed through the prism of their financial results, disregarding the scale and level of risk taken. The effects of the practice are still felt by financial markets. In the light of it, the postulate to integrate the bank's business goals with the goals of risk management systems (Nishiguchi et al., 1998) should be considered correct, which enables the implementation of the concept of sustainable development on a microeconomic basis.

The above-mentioned phenomena taking place in the bank's environment have a number of implications that have become a challenge for each bank, in particular in terms of management (Härle et al., 2016; Jain et al., 2021). The phenomenon of globalisation, in a broad sense, exerts a significant pressure on the improvement of operational efficiency as a key element of competitive advantage (Balling et al., 2001). Even if some banks have adopted a market niche (specialization) strategy, they usually reach the growth limit set by the size of a given market segment after some time. Then the problem of strategic directions for further development arises. Of course, they can only stick to their own specialization, which is facilitated by the dynamically changing environment and the related possibility to meet new needs. However, in the longer term, this may prove insufficient due to the strong competition in the banking sector, including cross-border banks. In such a situation, a more advantageous solution is the evolution towards a universal bank with an offer for all market segments (Mergaerts and Vennet, 2016). An additional advantage of the solution is the possibility of a more flexible diversification of activities and the associated risks. However, the problem in this case is how to expand the bank's business profile. It can usually be achieved in two ways: mergers/acquisitions of other banks or through reorganization (changes) on your own. The first method, however, requires sufficient capital resources, which is usually very difficult in practice in the case of a specialist bank. The second, in turn, requires, above all, appropriately qualified personnel and technological innovation, which is also indirectly related to capital resources (Zaleska and Kondraciuk, 2019). Regardless of the chosen development strategy, the basic problem ultimately comes down to the issue of operational efficiency (Paula, 2002). It determines, in the long term, a bank's ability to develop its activities in the broad sense (territorial or segment expansion). An important issue in this context is also the effectiveness of competitors (the entire banking sector). Significantly lower efficiency than the average in the sector will have an impact on weakening the competitive ability of a given bank. Therefore, regardless of the profile of its business, each bank should take into account the issue of efficiency when designing all processes. In the current reality, it is important because technological progress constantly creates new opportunities and tools to increase the effectiveness of the bank's operations (Weigelt and Sakar, 2012; Le and Ngo, 2020). It applies to the sale of banking products process as well as to other aspects of the bank's operation, including, in particular, the management process.

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In response to the described challenges facing the banking sector, the authors proposed a conceptual model of integrated optimization of bank's value (Koleśnik and Nadolski, 2021), which answers the question: how can a bank survive and develop in the conditions of strong competition, globalisation, financial markets integration (putting pressure on operational efficiency), while maximizing its value, as well as meeting the expectations of all stakeholder groups, including banking supervision? The purpose of the article is to verify the correct operation of this model and whether, as a result of its application, the bank can optimize its value (optimization should be understood as maximizing the value at an acceptable level of risk).

Due to the fact that the model is conceptual in nature, verification of the model will be carried out on the basis of a simulation method using pseudo-random numbers. High requirements regarding the detail of input data and the limited availability of relevant data limit the possibility of verification by authors to only one bank.

2. DESCRIPTION OF DATA

A simulation of the model's operation involved data of a specific bank from an EU Member State, hereinafter referred to as Bank X. The data was transformed in order to prevent identification. For the purposes of verification of model's operation, the Bank's initial balance sheet is presented in Table A.1 (Appendix A, Table 1). The parameters of the selected bank, including size, type and scale of operations and balance sheet structure, were similar to the average value of analogous parameters of a typical European bank, which is not identified as systemically important nor is it subject to direct supervision by the ECB within the banking union.

The balance sheet of Bank X does not show all the balance sheet items, which could actually occur and is presented in a simplified form, mainly to limit the duration of the simulations. This does not mean, however, that the results obtained will not be representative. From the point of view of the goal and structure of the model, neither the degree of balance sheet complexity nor the number of balance sheet items is important. The aspects affect only the duration of individual simulations but they have no impact on the final results obtained in different simulations, and thus obtained in the whole test. The simplification will also allow us to highlight proper effects of model operation without dispersing them excessively.

For the model to operate correctly, it is necessary to specify all the transactions carried out by the bank, including respective levels of income and risk. In order to verify the model, specification of parameters of the transactions (Table A.2) whose level was estimated on the basis of 7-year time series and characteristics of the transactions effected by Bank X were used.

The necessary input parameters of the model, which determine the target structure of transaction/balance sheet, are: ROE and total capital ratio (rate of return and risk appetite/bank security level). For the purposes of the simulations, it was assumed that the parameters are: 8.00% and 13.00%, respectively. Moreover, in order to determine the income value (regarded as the financial result of core activity) on the basis of the minimum ROE value, it is necessary to define the share of the remaining bank activity in the result obtained from core activity, as well as the rate of income tax. It was assumed that they are: 79.11% and 21.11%, respectively (values estimated on the basis of Bank X data).

Simplification made for the purposes of the simulations is based on a constant algorithm of off-balance sheet items, which is compatible with the observations made so far. On the basis of Bank X data, it was estimated that the degree of utilisation of off-balance sheet items was as follows: Z01 – 80%; Z02 – 70%; Z03 – 95% and Z04 – 25%. Furthermore, participation of the off-balance sheet items in creating the risk rate of individual off-balance sheet items is shown in Table A.3.

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3. MODEL DESCRIPTION

Verification covers only the strategic and management modules, which have a direct impact on stimulating the activities focusing the bank value optimization. Operational module derives from the other two modules and is designed to stimulate the desired behaviour of sales structures in the process of product sale (shaping the bank's balance sheet structure) and to implement a mechanism of participation in the effects of wrong decisions. The level of profits is directly related to the level of dynamic margin, which is the main tool of internal demand transmission and a stimulator of the desired structure of products/transactions sale. Therefore, in view of the goal of the model, its verification will be limited to the abovementioned modules, while its operation effectiveness will be tested by means of such measures and ratios as:

- ROE,
- ROA,
- total capital ratio,
- income rate,
- income to risk ratio.

The said measures, calculated to simulate the final outcome of the model's operation will be analysed, i.e. they will be compared with the measures calculated for the initial and target level.

In order to ensure reliability of model operation, including accuracy of final results, as well as to minimize fault tolerance, the number of simulations was set at 1000, while the number of steps in a single simulation – at 100 (N = 100). The course of a single step of model operation simulation can be related to one working day of bank activity and it includes: generating optimizing and related transactions (each time 10 transactions are generated – n = 10), generating other transactions and adjustments, and solving the optimization problem.

The model verification process assumes that different kinds of transactions will be effected; they are supposed to simulate events occurring during normal operation of every bank. It requires that the transactions are each time recorded in the balance sheet (balance sheet valuation). For the purposes of the simulation process, they are shown in the amount of the concluded transaction (without balance sheet valuation), which can only slightly affect the results obtained. Moreover, we can assume that simulated operations resulting from normal bank operation effectively eliminate the adopted simplification.

The adopted simulation mechanism of bank operation envisages the following types of transactions (arbitrary terminology):

- optimization transactions,
- transactions related to optimization transactions,
- random transactions,
- transactions related to random transactions,
- manual transactions,
- transactions of financial result revaluation,
- transactions related to transactions of financial result revaluation.

Optimization transactions are designed to be a set of transactions occurring on the market, which are possible to be made. They can be regarded as representing the current market demand. The transactions, however, are only a set of available transactions, out of which transactions of the most desired (in terms of bank value maximization) parameters are selected. The assumption adopted for the type of transactions is the following question: which transaction should be selected if we could choose only one from a quite numerous set? Each choice of a transaction of the most advantageous parameters should help the bank obtain the target transaction/balance sheet structure, at the same time eliminating unwanted transactions. Naturally, in real circumstances, it is difficult to imagine such a situation, but this approach is connected with the need of verification of model operation. In other words, whether the adopted criteria for selecting a transaction and

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thereby shaping the main factor motivating the sales structures (benefits) are appropriate and whether they will contribute to an increase of the bank value.

An assumption was made that optimization transactions are the ones which can be made by sales structures, i.e. receivables from the non-financial and public sector (A041, A042, A043, A05), debt securities (A06), securities (A07) and obligations to the non-financial and public sector (P031, P032, P032 and P04). The types of transactions are subject to the dynamic margin mechanism and they are the main factor affecting the transaction/balance sheet structure.

The process of generating optimization transactions is based on the mechanism of pseudo-random numbers, which involves random selection of a number from uniform distribution, which is a basis for creating the transaction amount, income rate and risk. The parameters are created by means of the quantile function of normal distribution, assuming that a higher transaction carries greater risk but has a lower income rate. The assumptions are based on empirical observations, which indicate that a higher transaction amount usually carries greater risk (resulting also from higher concentration, liquidity or credit risk). A reverse trend can be observed in the case of income rate, where higher transaction amounts are usually connected with lower margins as compared with transactions involving low amounts. Large amounts are usually given preference by a possibility to negotiate interest rates, which results in reduced transaction income. Conversion of random numbers from uniform distribution into the parameters of optimization transaction is based on the assumptions presented in Table A.4.

The process of generating optimization transaction each time involves a random selection of 100 numbers for every type of transaction. In effect, the optimum transaction is each time selected from a set of 1000 new transactions (10 types of optimization transactions, and 100 transactions per each type). At the same time, in order to prevent a set from reducing its number, the transaction amounts expressed in negative numbers, which can be created as a result of random selection mechanism, are converted into amounts expressed in positive numbers.

When an optimization transaction is effected, adjustments to the balance sheet items are required. It is due to the fact that crediting actually involves indication of balance sheet sources of financing of such a transaction (change in the balance sheet structure on the assets side). On the other hand, when a deposit is made, the balance sheet total increases and its structure is changed (change on the side of assets and liabilities). In the verification model, the type of transactions which introduce necessary adjustments after an optimization transaction is made are transactions connected with optimization transactions. The main assumption for the type of transaction is the fact that each optimization transaction, when effected, is settled by the amounts due from the financial sector (A03) and debt securities (A06). However, some limitations were applied with regard to the balance sheet items and the very optimization transaction. The limitations are supposed to reflect the actual conditions in which the bank operates. An optimization transaction is controlled to check whether it will or will not cause the limits established for a given type of transaction (the limits constitute limiting conditions for the bank value optimization) and the target risk rate. It is also checked whether it is possible to settle the optimization transaction by means of the abovementioned balance sheet items (control of acceptable limits). If the two conditions are met, the selected optimization transaction may be effected (otherwise, it is returned and the selection process is repeated). If the transaction is made, there arises a problem of indicating its sources of financing (active transactions) or its allocation (passive transactions). The adopted solution is presented in Table A.5.

The above model of related transactions typology should be interpreted in the following way: in the case of an active transaction, if it can be financed from items A03 and A06 (difference between their balance sheet total and the minimum level resulting from the limit), the A03 financing variant is adopted (for $A03 > 0$ and $A06 = 0$; $A03 > 0$ and $A06 < 0$; $A03 = 0$ and $A06 < 0$), the A06 variant (for $A03 = 0$ and $A06 > 0$; $A03 < 0$ and $A06 > 0$; $A03 < 0$ and $A06 = 0$) or the A03 and A06 variants – in proportion to the surplus/deficit (for $A03 > 0$ and $A06 > 0$; $A03 = 0$ and

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A06 = 0; A03 < 0 and A06 < 0). If the effected optimization transaction involves a purchase of debt securities (A06), such a transaction is fully financed from item A03. On the other hand, in the case of passive transactions, the algorithm is analogous to the presented one, except that the upper limit for A03 and A06 is examined (whether the limits of maximum engagement will be exceeded or not after the transaction is made).

In result of the presented algorithm, related transactions (one or two) are created. They have opposite signs (+/–) to the optimization transaction (total of the optimization transaction and the related transactions equals zero) if the optimization transaction refers to the assets. If the optimization transaction refers to the liabilities, they have the same sign. This is supposed to reflect the real mechanism of recording operations adopted by the bank. Moreover, in accordance with accounting principles, the algorithm ensures that the total of assets and liabilities will be consistent. On the other hand, the income and risk rates for related transactions are generated by means of pseudo-random numbers of uniform distribution using the quantile function of normal distribution. The difference in respect of the parameters for optimization transactions lies in the independent character of the random number for the income and risk rates. In principle, this should reflect varying conditions which often occur on the market, as well as the need to perform regulatory and hedging transactions in the dynamically changing reality. Average income and risk rates necessary to generate their random values for items A03 and A06 are compliant with Table A.2, while standard deviation for the income rate is 0.05%, as well as A03 – 0.10% and A06 – 0.00% for the risk rate.

So far, the transactions discussed in the article have reflected only a simulation of making transactions which are desired from the point of view of bank value maximization and the transactions resulting therefrom (the need to adjust the balance sheet and/or the balance sheet total). However, normal bank operations involve a series of other transactions, which are only a consequence of decisions made by the clients and booked economic events. For a simulation to better reflect the actual conditions of bank operation, such operations must be generated and shown in the balance sheet. The transaction category is arbitrarily referred to as random transactions and it is supposed to make the simulation more dynamic. Every day, in the bank there occur operations which are not necessarily desired from the bank's point of view but the character of services offered by the bank offers essentially unlimited opportunities for the clients to use the funds they deposited. Furthermore, there are a number of ongoing processes in the bank, including management of assets and liabilities, which also affect the bank's balance sheet structure. In order to reflect the operations, a mechanism of random generation of transaction parameters was applied to all balance sheet items, analogously to the case of related transactions connected with optimization transactions (independently generated amount, profit and risk rate). Generated transaction amounts can have both negative and positive values. They cannot, however, cause a negative balance (then the transaction amount equals 0). The parameters applied to generate random transactions amounts are shown in Table A.6.

Average values used for generating income and risk rates for different types of balance sheet transactions each time derive from the closing balance sheet (balance sheet which includes all the transactions effected within one simulation event). The standard deviation was established at 0.05%. It is worth noting – in Table A.6 – that an average greater than zero means that a balance sheet item is likely to rise in a long-term perspective.

In order to book random transactions, application of related transactions is required, like in the case of optimization transactions. Related transactions are supposed to finance the effected random transactions and to make the balance sheet total consistent. Also here, the effected transactions are settled from items A03 and A06, but it is done collectively (as is the case in reality). Settlement of individual balance sheet items is usually done at the end of the business day, not after each effected transaction. The amount to be settled is calculated as a difference between the total of active and the total of passive random transactions. If there is a surplus, items A03 and A06 are

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diminished in proportion to the current balance by the obtained value (decline in the balance sheet total value). Otherwise, they are increased (rise in the balance sheet total value). Related transactions are booked simultaneously with random transactions, which ensure compatibility of the balance sheet total. On the other hand, the parameters necessary to generate income and risk rates are analogous to those in random transactions (the same source of data).

Manual transactions (arbitrary term) are supposed to reflect settlement transactions performed at the end of every business day. It was assumed that only items A01, A03 and A06 are subject to such operations. They consist solely in settling the balance between the items and clearing them in proportion to the structure of their target levels in order to ensure the lowest possible mismatch between the balance sheet structure and the target structure, as well as to eliminate the effects of all the previous transactions. The income and risk ratios are generated analogously to those in the previous transaction type.

All the previous transactions cause changes in the balance sheet structure, balance of individual balance sheet items and their average weighed income and risk rates. It naturally affects the value of the financial result, which has not been discussed herein yet, but undoubtedly, it is changed too. According to the adopted solution, financial results depend on the average weighed income rate (all the previously described transactions considered). The product of this rate and the value of balance sheet total represents the result obtained from banking operations. The net financial result is obtained after an adjustment for the ratio of the other bank activity to the banking operations and for the income tax rate. The difference between the obtained value and the financial result achieved so far has to be recorded in item P10.

Inclusion of the transaction which involves revaluation of the financial result will make the balance sheet total inconsistent. It is therefore necessary to make another adjustment (transaction related to the transaction of financial result revaluation). It is carried out in exactly the same way as manual transactions are made, the only difference being that the amount to be recorded (increase/decrease in net financial result) is divided only between two items, i.e. A03 and A06.

Record of the transaction types described above is supposed to reflect their influence on the balance sheet structure, income and risk rates of individual balance sheet items. The parameters obtained in result of subsequent transactions being made are subject to continuous change. For this reason, it is each time necessary to solve the optimization problem since the input parameters, which considerably affect the final effect of model operation, are changed. The objective function, formulated on the basis of a simplified balance sheet and assumptions described above, can be presented in the following way:

$$F(x_i, y_j, z_k) = B \times \left(\sum_{i \in A} x_i D_i + \sum_{j \in P} y_j D_j + \sum_{k \in Z} z_k D_k\right), \tag{1}$$

where:
$F()$ – objective function,
B – value of balance sheet total,
D_i, D_j, D_k – income rate for a given transaction/item/product
x_i, y_j, z_k – share of assets (A), liabilities (P) and off-balance item (Z) in balance sheet total,
$A = \{$A01, A02, A03, A041, A042, A043, A05, A06, A07, A08, A09$\}$,
$P = \{$P01, P02, P031, P032, P033, P04, P05, P06, P07, P08, P09, P10$\}$,
$Z = \{$Z01, Z02, Z03, Z04$\}$.

The objective function is maximized using the following limiting conditions:

$$B \times \left(\sum_{i \in A} x_i R_i + \sum_{j \in P} y_j R_j + \sum_{k \in Z} z_k R_k\right) \le WR \tag{2}$$

$$\sum_{i \in A} x_i = \sum_{j \in P} y_j = 1 \tag{3}$$

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$$\sum_{k \in Z} z_k \leq 0.20 \tag{4}$$

$$L_d \leq x_i, y_j, z_k \leq L_g \tag{5}$$

$$x_{A041} + x_{A042} + x_{A043} \leq 0.75 \tag{6}$$

$$y_{P031} + y_{P032} + y_{P033} \leq 0.85, \tag{7}$$

where:
R_i, R_j, R_k – risk rate for a given transaction/item/product,
WR – admissible value of risk calculated on the basis of the target total capital ratio and current own funds,
L_d – bottom limit of share of a given balance sheet item in the balance sheet total,
L_g – upper limit of share of a given balance sheet item in the balance sheet total.

Assumptions (4), (6) and (7) were adopted by the authors arbitrarily for the purpose of verifying the operation of the model and are intended to reflect, respectively:
- the permissible share of off-balance sheet items in the carrying amount is 20%,
- credit exposures to the non-financial sector at a level not higher than 75% of the carrying amount,
- financing the bank's operations with deposits from the non-financial sector at a level not higher than 85%.

Upper and bottom limits (L_d and L_g) for individual balance sheet items are presented in Table A.7.

The income and risk rates used for calculating the value of objective function and boundary conditions each time derive from transaction parameters specification (Table A.2), not from the balance shown after the transaction is effected (current). Otherwise, the model would cause greater exposure of the bank to the economic cycle (the structure would be maximized on the basis of current parameters of profit and risk, not on the parameters estimated in long time series).

It can also be noted that Table A.7 does not include all the balance sheet items which are subject to limitations. It is due to the fact that the missing balance sheet items are limited on the basis of the current structure, being exogenous parameters for the optimization problem. This approach is based on the assumption that some balance sheet items can be shaped by the motivational system, while others are only a consequence of specific events and phenomena occurring in the bank and its environment. In practice, the calculated share of an item which is not subject to limitations constitutes a boundary condition for the goal function. In other words, x_i, y_i equal the calculated share of a given item in the balance sheet total.

The balance sheet total (current balance sheet total) also constitutes exogenous data for the optimization problem, as well as the share of off-balance sheet items, where a constant algorithm of its creation was adopted.

Microsoft Excel including Solver software was used to solve optimization problems. The obtained results are presented in Appendices B and C hereto.

4. SIMULATION RESULTS

On the basis of the obtained results (Table B.1 – Appendix B, Table 1), we can conclude that the model causes growth of the balance sheet total by an average of 2.26% (average growth amount 6,766.58 thousand monetary units /K MU/) at the median which differs from the average only by 26.60 K MU. The level of standard deviation, which is only 0.58% in respect of the average, is relatively low. Analysis of a frequency diagram (Figure C.1 – Appendix C, Figure 1)

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indicates that in the case of 50% of performed simulations, the final balance sheet total oscillated between 305,181.40 and 307,519.40 K MU.

The off-balance sheet items, which increased by an average of 20.77%, showed greater growth dynamics (with an average growth amount of 9,480.31 K MU). The difference between the median and the average was only 4.84 K MU. As was the case of the balance sheet total, the level of standard deviation, which was 1.27% in respect of the average, is relatively low. In the case of approx. 80% of performed simulations, the off-balance sheet total oscillated between 54,192.41 and 56,006.12 K MU.

With regard to the balance sheet structure, we can therefore conclude that the model stimulates growth of the bank book value. At the same time, results of individual simulations show considerable concentration, which means that the model operation is stable while the obtained results do not indicate substantial dispersion.

The net profit recorded average growth of 16.14% (amount of 326.59 K MU), which – with regard to an increase in bank value – is a desirable outcome. The difference between the median and the average remained at quite a low level of –0.02 K MU while the level of standard deviation in respect of the average was only 0.76%. **It means that each time the model causes growth of bank value while the generated financial result is characterised by very low dispersion**. It is demonstrated in the frequency diagram (over 95% of simulations generated net financial result with the range between 2,316.65 and 2,384.15 K MU).

Similar conclusions can be drawn in respect of income, from which the financial result derives (due to the assumed stable relative level of profit encumbered with the result from non-banking activity and with the income tax). Differences refer only to absolute values while the relative ones remained at a similar level.

The average value of risk expressed as an amount reached the level of 15,406.80 K MU and it was higher that the initial value by 6.47% (936.80 K MU). It should be noted that the recorded increase in risk value was lower than in the case of income (income rose by an average of 1,917.18 K MU). It is quite significant if we consider that the basic goal of the model was to increase the bank value exactly, among others, by the selection an algorithm which gives preference to the transactions of desired income/risk ratio. In other words, the model causes a change in transactions carrying very high risk into transactions carrying lower risk while maintaining appropriate profitability level. The difference between the median and the average remained at quite a low level of 12.02 K MU while the level of standard deviation in respect of the average was only 1.63%. In the case of approx. 90% of the performed simulations, the amount of risk generated by the bank oscillated between 15,001.28 and 15,826.18 K MU, which indicates – like in previous cases – low dispersion of simulations results.

In the case of income and risk value, the observed phenomena are reflected in the ratio between the values. As a result of the model's operation, the income/risk ratio followed the expected values recording an average increase by 8.53% (7.27 percentage points). **In practice, it would mean an improvement in bank operation profitability and an increase of bank value, i.e. realization of the abovementioned basic goal of the model.** The absolute difference between the median and the average was –0.04 p.p., at standard deviation to the average ratio of 1.06%. It indicates a big concentration of the obtained results, which is also confirmed by the frequency diagram (over 90% of observations oscillate between 91.03% and 94.32%). Stability of the model's operation may be additionally confirmed by the figures illustrating the course of different simulations (Figures C.12–17). In the case of income/risk ratio, but also other values, a curve illustrating the course of this ratio during the simulation can be plotted. The important fact is that each simulation course has a quite characteristic and similar shape, while individual results (of different steps, as well as of any other simulation) show a considerable concentration.

Although income and risk rates were analogous to the income and risk values, effects of the model's operation are more spectacular in this case. In other words, they better illustrate

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the mechanism of the model's operation. The income rate increased on average by 11.56% (an increase of 0.22 p.p.) while the risk rate rose only by 2.81% (an increase of 0.06 p.p.). **It means that the transaction selection mechanism gives preference to low-risk and high-profitability rate transactions, which was one of the fundamental assumptions for the model.** The medians for the two ratios do not differ from the average values, which were: 2.13% and 2.31% for the income and risk rates, respectively. On the other hand, the standard deviation in relations to the average income rate showed a much lower value (0.60%) than for the risk rate (1.52%), which results from the applied random selection mechanism. Nonetheless, stability of the model's operation is confirmed by frequency diagrams (Figures C.7–8), which demonstrate that over 90% of observations for the income rate oscillate between 2.1111% and 2.1564%, while for the risk rate – between 2.2479% and 2.3655%. The figure plotting the risk rate value at individual steps of the simulation indicates even greater dispersion of the obtained results. Nevertheless, different data series allow us to plot the characteristic shape of the curve. In the case of income rate, the shape of the curve is markedly easier to plot, and the results of individual simulations are characterised by a pronounced concentration.

The most important synthetic bank security ratio is the total capital ratio. As a result of the model application, its average value from all the performed simulations was 12.74%, which means a decrease by 5.29% (0.71 p.p.) in respect of the initial value. The fall in the value of total capital ratio below the target value (for the purposes of the simulation its value was adopted at 13.00%) results exclusively from the lack of boundary conditions in respect of risk (lack of risk limits). The step was made only because it was necessary to obtain unambiguous results of the model's operation with regard to the bank financial effectiveness while limiting the number of steps in individual simulations at the same time. It should be noted, however, that in no case did it reach a value below the required regulatory minimum, which – for Bank X – is 9.25% (own funds requirements – 8%, capital conservation buffer – 1.25%, countercyclical capital buffer – 0.00%, systemically important institution buffers – 0.00%, systemic risk buffers – 0.00%). The difference between the median and the average was only 0.01 p.p., while the standard deviation level in relations to the average, which was 1.63%, should be regarded as relatively low. Over 90% of the recorded results oscillate between 12.38% and 13.11%, while the diagram plotting the course of individual steps in each simulation confirms stability of the results generated by the model.

Return on equity (ROE) – a very important ratio for bank investors – oscillated around 9.86%, which is higher than the initial value by 15.90% (1.35 p.p.). It means that an increase in profitability of bank operations was markedly higher than the decrease of the abovementioned total capital ratio, which is in turn an important ratio for other bank stakeholders. We can therefore conclude that the extreme increase in benefits was higher than the extreme increase of potential risk costs. There was no difference between the median and the average while the standard deviation level in relation to the average should be regarded as relatively low (0.76%). Results of over 90% of performed simulations oscillated between 9.74% and 10.00%.

Another reliable ratio illustrating an improvement in bank operation profitability is return on assets (ROA), which recorded an average increase by 13.58% (average increase by 0.09 p.p.) with no difference between the median and the average. The standard deviation level was relatively low with an average level of 0.005%, and in respect of the average – only 0.65%. An analysis of the frequency diagram (Figure C.11) indicates that in the case of over 90% of performed simulations the final ROA value oscillated between 0.7581% and 0.7760%. The figure illustrating the value at different steps of each simulation has a very specific shape and is characterised by a considerable concentration of results, similarly as the figure for the ROE ratio.

Correctness of the model's operation and accurate simulation of the conditions of bank operation can be evaluated on the basis of identified correlation relationships between profitability and risk ratios. Table B.2 shows Pearson's correlation between profitability/income ratio, total capital ratio and risk ratio observed for every simulation. As can be seen, total capital ratio,

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representing the bank security level, in each case shows a negative correlation with profitability/income ratio. As per J. Guilford's classification, the strength of the relations can be described as very high. The return on equity ratio demonstrates the strongest relationship with total capital ratio, which means that a change in the bank security level will undoubtedly be reflected in the return on capital level. The relationship, however, is in inverse proportion. Differences in the levels of correlation between total capital ratio and different profitability/income ratio are due to methodological aspects of ratio determination (different reference base – the higher the base, the weaker the relationship). Regardless of the differences, the observed correlations between income/profits and risk show the expected trend, this means that both the model assumptions and the mechanism of random and selecting transactions were designed correctly.

On the other hand, the risk ratio has a high (ROE) and very high (ROA, profitability) positive correlation with the profitability/income ratios. The phenomenon is compatible with the previously mentioned statement with regard to profitability and risk. It is worth noting that the strongest relationship occurs with the ROA ratio and only slightly less strong – with income rate. It means that the ratios may serve as perfect predictors (stimulants) of the effects of the model's operation. In contrast, correlation with the ROE ratio shows a lesser but still quite strong relationship with the risk rate. In other words, the risk rate affects the value of ROE, which is a key measure for bank investors, to a lesser extent but it significantly determines other profitability ratios.

The correlation relationships shown in the figures (Figures C.18–23) confirm the conclusions presented above. Very strong correlation relationships between ROE and total capital ratio, ROA and risk rate as well as between the income and risk rates can be observed. In the case of the other relationships, greater dispersion of individual observations can be noted. However, relatively small ranges between extreme values for individual variables must be emphasized.

At the basis of the model structure there was an underlying assumption that the model mechanism would cause a reconstruction of the initial balance sheet structure tending to a balance sheet structure which is an optimum in terms of bank value maximization in the long-term perspective. Therefore, apart from the financial effectiveness and risk rates, the mismatch between the final balance sheet structure and the optimum structure is also a measure of the model's operation effectiveness. The mismatch was measured by dispersion measures, namely the range and standard deviation. When calculating the range, differences between the initial or final structure and the target structure are determined in the first place, and then the maximum and minimum mismatch values are differentiated.

As can be observed on the basis of the obtained results (Table B.3), the model causes a decrease in the range of assets in respect of the initial structure by an average of 80.90% (average decrease of range by 19.91 p.p.) with the median differing from the average only by 0.01 p.p. The standard deviation level, which is 19.24% in respect of the average, should be regarded as relatively significant. An analysis of a frequency diagram (Figure C.24) indicates that in the case of over 90% of simulations performed, the final range value oscillated between 3.09% and 6.10%, which – in comparison with the initial value of 24.61% – is a very good result. In practice, it means that the model significantly affects the reduction of the existing mismatch; maintaining high operation stability at the same time.

In the case of liabilities, where the range decreased only by 5.35% (a decrease by 0.41 p.p.), with no difference between the median and the average, the model showed a definitely lower level of mismatch reduction. A lower value was also observed in the case of the coefficient of variability, which totalled 8.33% and was twice lower than in the case of assets. In respect of over 90% of the performed simulations, the final range of liabilities oscillated between 6.25% and 8.32%, however, approx. 26% of the observations exceeded the initial value. The situation was caused in the first place by the transaction selection mechanism, which first selected active transactions and only when the engagement limits were finished, it effected passive transactions. In the context of reducing the number of steps in individual simulations, it resulted in a lower number of passive

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transactions being made, and thereby a smaller impact on the structure of liabilities, which was regulated by random transactions to a greater extent (hence the cases of exceeding the initial range value).

In the case of off-balance sheet items, which – due to the simulation model assumptions – derived from the effected active transactions, a similar (to the case of assets) model's operation was observed.

The range for the off-balance sheet items recorded a medium decrease of 48.89% (change by 0.62 p.p.), with no difference between the median and the average. Like in the case of assets, standard deviation level, which is 20.61% in respect of the average, can be described as relatively significant. An analysis of a frequency diagram (Figure C.26) indicates that in the case of over 90% of the performed simulations, the final range value of the off-balance sheet items oscillated between 0.42% and 0.87%, and it did not exceed the initial value. In practice, it means that the model has a significant impact on the reduction of the existing mismatch in respect of the off-balance items.

As can be observed, the model causes a decrease in the standard deviation of the final assets structure in respect of the target structure by an average of 82.15% (average deviation reduction by 5.11 p.p.), with the median differing from the average only by 0.01 p.p. The level of coefficient of variability, which is 20.20%, should be assessed as relatively significant. An analysis of frequency diagram indicates that in the case of over 90% of the performed simulations, the final value of standard deviation of the mismatch oscillated between 0.74% and 1.49%, which – in comparison with the initial value of 6.23% – is a very good result. **In practice, it means that the model has a significant impact on the reduction of mismatch in respect of the target structure, maintaining high operation stability at the same time.**

The model showed a decisively lower level of mismatch reduction in the case of liabilities, where standard deviation decreased only by 3.33% (a decrease by 0.06 p.p.), with no difference between the median and the average. A lower value was also observed in the case of the coefficient of variability, which totalled 6.59% and was three times lower than in the case of assets. In the case of over 90% of the performed simulations, the final range of liabilities oscillated between 1.54% and 1.92%, while over 30% of the observations exceeded the initial value. The reason for the situation is analogous to the one mentioned in the liabilities range analysis.

The standard deviation for the off-balance sheet items recorded a medium decrease of 55.17% (a change by 0.33 p.p.), with no difference between the median and the average. Like in the case of range measure, the level of variability coefficient, which is 20.51%, can be assessed as relatively significant. An analysis of a frequency diagram indicates that in the case of over 90% of the performed simulations, the final value of standard deviation of off-balance sheet items oscillated between 0.18% and 0.36%, and it did not exceed the initial value. **It confirms the thesis that the model has a significant impact on the reduction of the existing mismatch in respect of the off-balance sheet items.**

Essentially, the model should reduce the final balance sheet mismatch in respect of the target structure. On the other hand, the target structure should not exceed the engagement/concentration limits determined by the bank. It is therefore necessary to examine whether the final balance sheet structure is consistent with the limits set by the bank in order to ensure that they are not exceeded by the model's operation.

For the purposes, a mismatch measure calculated as a total of absolute differences (occurring only when limits are exceeded) between the initial or final structure and the (upper and lower) limits for the assets, liabilities and off-balance sheet items was used.

The mismatch between the balance sheet structure and the limits set by the bank in result of the model application was an average of 0.44 p.p. with no such mismatch recorded for the initial balance sheet. The difference between the median and the average was only –0.04 p.p. On the other hand, the standard deviation to the average ratio, which was 66.28%, should be

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assessed as particularly significant. An analysis of a frequency diagram indicates that in the case of 95% of performed simulations, the final value of limit mismatch does not exceed the value of 0.99%, while all observations exceeded the initial value. The indicated effect is caused by an implemented mechanism generating random transactions, which cause a slight overrun of the limits set by the bank when the effected optimization transactions exhausted them. In terms of the model's operation, it is acceptable, especially that the observed overruns were not higher than 2 p.p. and the transaction selection mechanism did not continue effecting transactions with limit overrun once the limit was exceeded.

The performed simulations indicate that the model's application and implemented transaction selection mechanism (a parameter controlling the level of dynamic margin) enable an increase of the bank value (financial result) and better financial effectiveness of the bank. Naturally, it should be remembered that the effected transactions had a random character. Nonetheless, the transaction selection algorithm worked correctly, selecting transactions which reduced the difference between the initial and the target balance sheet structure, at the same time improving financial ratios of the bank. **In practice, it means that regardless of real transaction parameters (other than random ones), the model will give preference to the transactions of expected characteristics and will stimulate the bank sales structures to perform expected operations. The model operational stability, proved by quite low values of dispersion measures, should also be positively assessed. It means that – despite the different transaction parameters and random selection mechanism generating random operations, in each case the model ensures an increase in bank value and improvement of bank operations effectiveness, while the values achieved during individual simulations are similar. The simulation model also reflects the real conditions of bank operation in a way which should be positively assessed in the light of our analysis of correlation between profitability/income and risk.**

5. CONCLUSIONS

The conducted verification of the conceptual model of integrated optimization of a bank's value has shown that the model works in accordance with the assumptions adopted by the authors and leads to the achievement of the basic goal for which it was constructed (Koleśnik and Nadolski, 2021). The model was designed to stimulate the bank's sales structures to conclude transactions that consequently maximize its value. The model was designed for it and only it. A separate issue is whether concluding such transactions will be possible (demand and supply). A simulation is intended to demonstrate only whether the model works correctly and achieves the intended goal within certain constraints. The applied boundary conditions in the defined optimization problem, as well as their process location, not only reflect regulatory prudential requirements, but also increase flexibility of the model in the context of its implementation in banks with different organizational and competence solutions. In the algorithmic layer, they are an element that implements a compromise between the conflicting goals of various groups of bank stakeholders. It is also worth noting that with their help it is possible to adapt the model to possible future prudential requirements without the need to change the structure of the model.

The simulations carried out showed that the model increases profitability of the balance sheet while preferring transactions with a lower risk rate, which in turn leads to optimization and an increase in the value of the bank in the long term. It also effectively reduces the mismatch between the structure of the initial balance sheet and the target structure, while maintaining the limits applicable in the bank (the small excesses found were caused only by random transactions, which were exogenous to the model). The obtained results confirmed stability of its operation

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(small dispersion scale of observations of key financial efficiency and security indicators) and the correctness of the transaction selection mechanism, which is a key element of the entire model.

In conclusion, the model proposed by the authors is a successful attempt to integrate the risk management process and the bank's business activity based on the current solutions used in the banking activity and their creative synthesis enabling the automation of the bank management process. The inclusion of risk in the process of making decisions of an operational, managerial and, above all, strategic nature as an important criterion (equally important as the profitability criterion) allows for the optimal allocation of capital from the point of view of the sustainable nature of the business, and above all from the point of view of building a sustainable competitive advantage.

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APPENDICES

Appendix A – INPUT DATA

Table 1
Initial balance sheet of Bank X

Item	Assets	Balance	Profitability rate (%)	Risk rate (%)
A01	Cash	4,143.93	–1.02	0.00
A02	Cash at Central Bank	0.00	0.00	0.00
A03	Receivables from financial institutions	80,275.09	0.89	1.62
A04	Receivables from non-financial sector	125,642.53	6.64	7.23
A041	Corporate	23,167.18	3.94	9.06
A042	Small and medium enterprises	76,140.40	6.99	6.48
A043	Retail	26,334.95	8.02	7.77
A05	Receivables from governmental and self-governmental institutions	20,784.17	0.97	1.47
A06	Debt securities	53,668.91	1.38	0.00
A07	Securities	0.00	0.00	0.00
A08	Intangible and legal assets and tangible fixed assets	13,062.19	0.00	8.26
A09	Other assets	2,021.98	0.00	2.16
	TOTAL	**299,598.80**	**3.32**	**3.94**
	Liabilities			
P01	Liabilities payable to the Central Bank	0.00	0.00	0.00
P02	Liabilities payable to financial institutions	3,139.87	–0.40	0.70
P03	Liabilities payable to non-financial sector	244,383.16	0.90	0.65
P031	Corporate	12,071.59	0.72	1.27
P032	Small and medium enterprises	58,764.27	0.71	1.07
P033	Retail	173,547.30	0.98	0.46
P04	Liabilities payable to governmental and self-governmental institutions	16,985.62	0.54	1.77
P05	Liabilities due to financial instruments	0.00	0.00	0.00
P06	Provisions	266.00	0.00	0.00
P07	Subordinated liabilities	780.00	–0.46	0.00
P08	Share capital	23,782.51	0.00	0.00
P09	Reserves and other liabilities	8,238.72	0.00	0.00
P10	Net profit (loss)	2,022.92	0.00	0.00
	TOTAL	**299,598.80**	**0.76**	**0.63**
	Off-balance sheet items			
Z01	Low risk items e.g. undrawn credit facilities (unconditionally cancellable)	14,426.56	0.10	0.00
Z02	Medium/low risk items e.g. undrawn credit facilities with an original maturity of up to and including one year	21,642.42	0.20	1.01
Z03	Medium risk items e.g. undrawn credit facilities with an original maturity of more than one year	2,334.18	0.20	2.73
Z04	Other items	7,234.23	0.50	6.61
	TOTAL	**45,637.39**	**0.22**	**1.67**

Source: Authors' calculations based on Bank X data.

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Table 2
Specification of transaction parameters (%)

Item	Name	Profitability rate	Risk rate	Transaction Time
A01	Cash	–1.02	0.00	4.17
A02	Cash at Central Bank	–0.57	0.00	33.33
A03	Receivables from financial institutions	0.90	1.61	66.67
A041	Corporate	3.98	9.05	100.00
A042	Small and medium enterprises	7.06	6.44	66.67
A043	Retail	8.10	7.75	33.33
A05	Receivables from governmental and self-governmental institutions	0.99	1.45	66.67
A06	Debt securities	1.48	0.00	8.33
A07	Securities	5.57	9.04	8.33
A08	Intangible and legal assets and tangible fixed assets	0.00	8.06	10.42
A09	Other assets	0.00	2.06	8.33
P01	Liabilities payable to the Central Bank	–0.30	0.00	4.17
P02	Liabilities payable to financial institutions	–0.40	0.67	50.00
P031	Corporate	0.82	1.22	4.17
P032	Small and medium enterprises	0.91	1.05	3.13
P033	Retail	1.06	0.45	2.08
P04	Liabilities payable to governmental and self-governmental institutions	0.74	1.75	6.25
P05	Liabilities due to financial instruments	–1.73	0.46	66.67
P06	Provisions	0.00	0.00	8.33
P07	Subordinated liabilities	–0.46	0.00	66.67
P08	Share capital	0.00	0.00	4.17
P09	Reserves and other liabilities	0.00	0.00	4.17
P10	Net profit (loss)	0.00	0.00	12.50
Z01	Low risk items	0.10	0.00	———
Z02	Medium/low risk items	0.20	1.01	———
Z03	Medium risk items	0.20	2.73	———
Z04	Other items	0.50	6.61	———

Source: Authors' calculations based on Bank X data.

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Table 3
Effect of balance sheet items on the risk rate of off-balance sheet items (%)

Balance sheet items	Off-balance sheet items			
	Z01	Z02	Z03	Z04
A03	40.00	40.00	20.00	0.00
A041	13.33	20.00	26.67	5.00
A042	16.67	25.00	25.00	2.00
A043	66.67	16.67	16.67	0.00
A05	40.00	30.00	30.00	0.00

Source: Authors' calculations.

Table 4
Conversion parameters for optimization transactions

Transaction type	Parameter	Value		
		Amount	Profitability rate (%)	Risk rate (%)
A041	Average	1,000	3.98	9.05
	Standard deviation	500	1.00	0.80
A042	Average	250	7.06	6.44
	Standard deviation	100	1.30	0.25
A043	Average	50	8.10	7.75
	Standard deviation	15	0.50	0.20
A05	Average	300	0.99	1.45
	Standard deviation	50	0.10	0.05
A06	Average	500	1.48	0.00
	Standard deviation	200	0.05	0.00
A07	Average	150	5.57	9.04
	Standard deviation	50	1.00	1.00
P031	Average	600	0.82	1.22
	Standard deviation	200	0.10	0.20
P032	Average	200	0.91	1.05
	Standard deviation	50	0.15	0.25
P033	Average	100	1.06	0.45
	Standard deviation	35	0.20	0.04
P04	Average	200	0.74	1.75
	Standard deviation	70	0.10	0.50

Source: Authors' calculations.

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Table 5
Classification of transactions related to optimization transactions

	A06 > 0	**A06 = 0**	**A06 < 0**
A03 > 0	proportionally	A03	A03
A03 = 0	A06	proportionally	A03
A03 < 0	A06	A06	proportionally

Source: Authors' calculations.

Table 6
Parameters of random transactions generation

Transaction type	**Average**	**Standard deviation**
A01	0.00	100.00
A02	0.00	0.00
A03	0.00	50.00
A041	0.00	250.00
A042	0.00	80.00
A043	0.00	50.00
A05	0.00	150.00
A06	0.00	300.00
A07	0.00	0.00
A08	5.00	35.00
A09	2.50	20.00
P01	0.00	0.00
P02	–2.50	5.00
P031	0.00	100.00
P032	0.00	80.00
P033	0.00	40.00
P04	5.00	30.00
P05	0.00	0.00
P06	0.00	2.00
P07	–1.00	4.00
P08	0.50	0.20
P09	4.00	30.00
P10	0.00	0.00
FW	2.00	0.50

Source: Authors' calculations.

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Table 7
Engagement limits for balance sheet items (%)

Item	Name	Bottom limit	Upper limit
A01	Cash	1.00	2.00
A02	Cash at Central Bank	0.00	0.00
A03	Receivables from financial institutions	10.00	30.00
A041	Corporate	5.00	10.00
A042	Small and medium enterprises	15.00	30.00
A043	Retail	5.00	35.00
A05	Receivables from governmental and self-governmental institutions	5.00	15.00
A06	Debt securities	10.00	20.00
A07	Securities	0.00	5.00
P01	Liabilities payable to the Central Bank	0.00	0.00
P02	Liabilities payable to financial institutions	0.00	15.00
P031	Corporate	0.00	15.00
P032	Small and medium enterprises	15.00	30.00
P033	Retail	30.00	60.00
P04	Liabilities payable to governmental and self-governmental institutions	5.00	20.00
P05	Liabilities payable to the Central Bank	0.00	10.00

Source: Authors' calculations.

APPENDIX B – RESULTS OF THE MODEL VERIFICATION: TABLES

Table 1
Results of performed simulations

Specification	Opening balance	Closing balance		
		Average	Median	Standard deviation
Balance sheet total	299,598.80	306,365.38	306,391.98	1,779.46
Total off-balance sheet items	45,637.39	55,117.70	55,122.54	700.70
Net profit	2,022.92	2,349.51	2,349.49	17.82
Income (amount)	12,339.50	14,256.68	14,256.63	108.12
Risk (amount)	14,470.00	15,406.80	15,418.82	250.65
Income/risk ratio	85.28%	92.55%	92.51%	0.9796%
Income rate (D)	1.91%	2.13%	2.13%	0.0128%
Risk rate (R)	2.24%	2.31%	2.31%	0.0350%
Total capital ratio	13.45%	12.74%	12.72%	0.2078%
ROE	8.51%	9.86%	9.86%	0.0748%
ROA	0.68%	0.77%	0.77%	0.0050%

Source: Authors' calculations.

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Table 2
Correlations between selected profitability and risk ratios

	Total capital ratio (TCR)	Risk rate (R)
ROE	–0.854147049	0.662126973
ROA	–0.763037971	0.879222567
Income rate (D)	–0.706471799	0.847636647

Source: Authors' calculations.

Table 3
Balance sheet structure mismatch (%)

Specification	Opening balance	Closing balance		
		Average	Median	Standard deviation
Range ASSETS	24.61	4.70	4.69	0.90
Range LIABILITIES	7.70	7.29	7.29	0.61
Range OFF-BALANCE	1.27	0.65	0.65	0.13
Standard deviation ASSETS	6.23	1.11	1.10	0.22
Standard deviation LIABILITIES	1.80	1.74	1.73	0.11
Standard deviation OFF-BALANCE	0.60	0.27	0.27	0.05
Mismatch between limits	0.00	0.44	0.40	0.29

Source: Authors' calculations.

APPENDIX C – THE MODEL VERIFICATION RESULTS: DIAGRAMS

Empirical distributions of the results of performed simulations

Figure 1

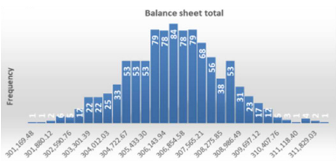


Figure 2

Figure 3

Figure 4

Figure 5

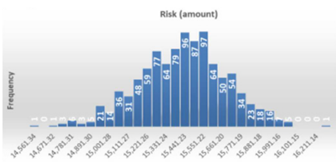


Figure 6

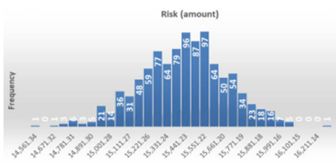
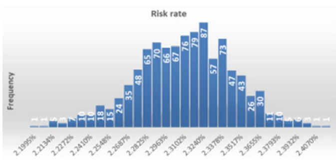


Figure 7

Figure 8



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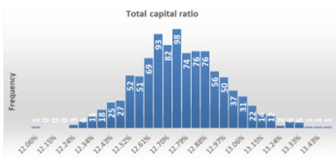

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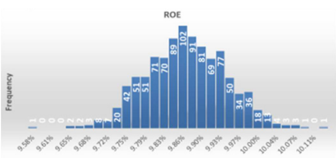

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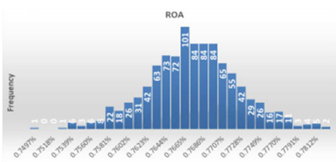

Source: Authors' elaboration.

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Selected profitability and risk ratios in the course of individual simulations

Figure 12

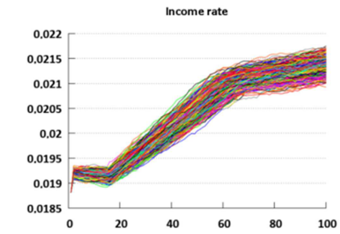


Figure 13

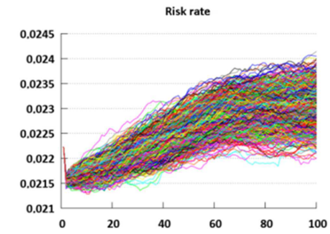


Figure 14

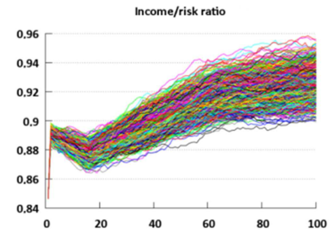


Figure 15

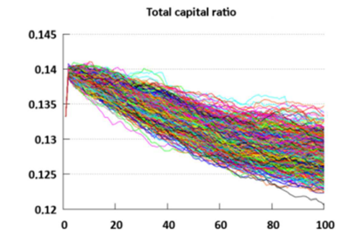


Figure 16

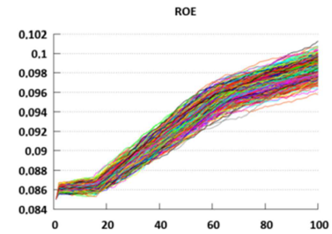
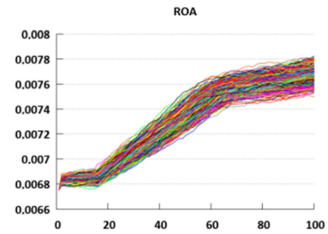


Figure 17



Source: Authors' elaboration.

Scatter plots of correlations between selected profitability and risk ratios

Figure 18

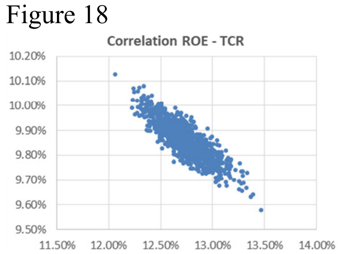


Figure 19

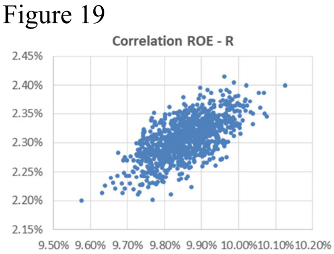


Figure 20

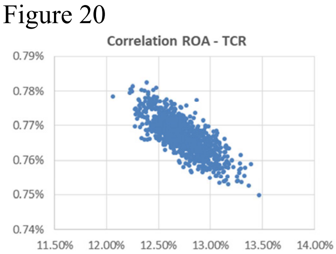


Figure 21

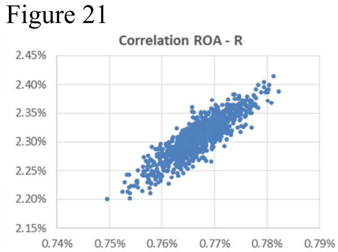


Figure 22

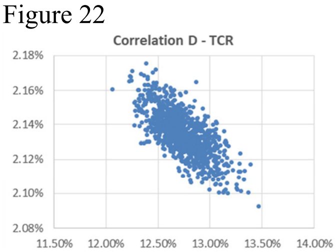
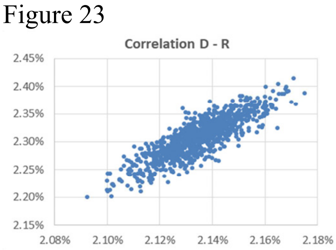


Figure 23



Source: Authors' elaboration.

Histograms of the results of model impact on the balance sheet structure

Figure 24

Figure 25

Figure 26

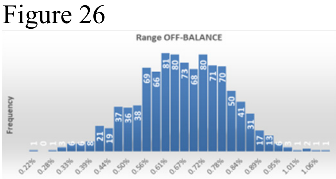


Figure 27

Figure 28

Figure 29

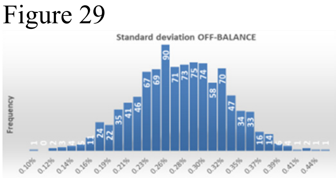
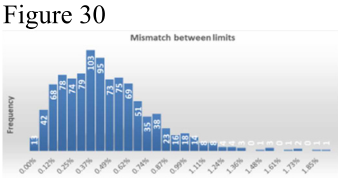


Figure 30



Source: Authors' elaboration.

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Long-term relationship of KGHM share prices and the market value of high grade copper

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ABSTRACT

The aim of the article is to try to explain the long-term price volatility of KGHM shares. Therefore the paper presents the relationship between KGHM stock prices and High Grade copper prices. The empirical part of the paper uses econometric cointegration analysis. Based on the estimated models, the thesis of the existence of a long-run relationship between the studied variables was confirmed. Within the framework of econometric analyses, Johansen and Engle-Granger procedures and the Granger test of causality were applied. The study was conducted using monthly data covering quotes from August 2012 to April 2021. In the end, both modeling procedures used led the researcher to convergent conclusions. Moreover very similar values of long-run equilibrium parameter estimates were obtained for both methods. Thus, on the basis, it is necessary to confirm the main hypothesis formulated at the beginning of the study, that is, in the case of a company operating within the KGHM Polska Miedź mining industry, the quotation of its shares is long-term dependent on the quotation of HG copper prices.

JEL classification: G17, C01, C5, G1

Keywords: causality, KGHM, Johansen procedure, Engle-Granger procedure, cointegration.

1. INTRODUCTION

The main aim of the article is to try to explain the long-term equilibrium of KGHM share prices listed on the Warsaw Stock Exchange (WSE). At the outset of the analysis carried out, it was assumed that the quotation of KGHM shares may depend on the quotation of copper, an important raw material exploited by the company. KGHM is unique among top copper producers in generating majority of its revenues from copper ore mining and copper processing. Therefore the relationship between prices of copper and share prices might be strong. It is worth noting that other top copper producers (BHP, Rio Tinto, Glencore) are more diversified mining companies. In the case of higher diversification the company's revenues might be dependant on prices of various commodities. The main hypothesis of the conducted study boils down to the statement that, in the case of KGHM Polska Miedź, a company operating within the mining industry, the quotation of its shares depends on the quotation of HG (Hight Grade) copper prices in the long term. Moreover, the course of the time series of both $KGHM_t$ and HG_t variables seen in Figure 1 (included in the methodology and data section) allows us to expect that

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the series are characterized by common long-term trends. Due to the statistical properties of the time series of KGHM stock quotes and copper prices, it is insufficient to use the quantitative correlation coefficient or classical regression analysis in their analysis. In the case of time series analysis, ignoring the issue of stationarity leads to the risk of the appearance of so-called sham regressions (Granger and Newbold, 1974). Econometric analysis for non-stationary time series requires the use of non-classical techniques (Majsterek, 2014; Domanski and Pruska, 2000). Hence, realization of the stated aim of the study requires a recourse to methods of analyzing financial time series and stochastic processes, which are the object of interest of the so-called new econometrics. In the indicated area, it becomes possible to identify common stochastic trends, long-term equilibrium relations and mechanisms of short-term adjustments, the knowledge of which is economically important in the context of the formulated objective of the study.

2. LITERATURE REVIEW

In the scientific literature, there have been attempts to analyze price quotations of copper and KGHM shares (Fijałkowski, 2011) for daily data of the Warsaw Stock Exchange from September 15, 2010 to March 15, 2011. The cited example covered a short period of time and was based on the estimation of a classical regression model. There are many valuable publications regarding the relationship between the stock and commodity markets. In most cases, researchers deal with links between stock and commodity markets from a volatility perspective (Creti et al., 2013; Kang et al., 2016; Vardar et al., 2018). Especially, a study about prices of shares of Southern Copper Corporation should be mentioned here (Zevallos and Carpio, 2015). It is worth noting that the cited studies concern stock and commodity markets volatility. It is an important objective of literature review to identify the gap in the existing literature. Therefore the article concerns a cointegration analysis between variables from the stock and commodity markets. The author developed a hypothesis about long terms relationship existing between $KGHM_t$ and HG_t. In this case the cointegration analysis let the author identify a common long-term trend despite of short-term volatility. Financial time series are often characterized by non-stationarity and require the adoption of procedures coinciding with the so-called new econometrics (Charemza and Deadman, 1997). For the reason, it seems interesting to carry out a study that is subject to the rigors of non-stationary data analysis and, at the same time, covers a sufficiently wide time period so that conclusions can be generalized to the long-term horizon. It should be added that cointegration analysis is more common in macroeconomic studies (Karp et al., 2013; Kusideł, 2000), as well as in the analysis of stock market index correlations (Assidenou, 2011). An example of this type of analysis for financial data is the study of the relationship between stock market indices, exchange rates (Gędek, 2013; Buszkowska, 2014). In the publications there is a typical problem of volatility clustering for stock market data, where the so-called effect "volatility clustering" is one of the accentuated properties of financial time series. A cointegration analysis for non-aggregated series is a far less common phenomenon, although also found in scientific literature (Baghestani, 1991). A review of the publications leads to the conclusion that it is more common for such models to be used in the analysis of commodity prices quoted on commodity markets. It is especially true for energy commodities such as liquid fuels and crude oil (Gędek, 2017; Socha, 2013; Erdos, 2012; Asche et al., 2003). Within the precious metals market, publications provide many interesting analyses of long-term relationships conducted from an investment perspective. Hence, they focus on the analysis of the cointegration of gold (Mamcarz, 2018), silver (Kasprzak-Czelej, 2016) and their relationship with the aforementioned energy commodities (Sindhu, 2013; Kasprzak-Czelej, 2018). In the English-language literature, researchers point to the existence of a long-term relationship between the prices of the raw materials (Lucey and Tully, 2006; Pierdzioch et al., 2015). It is worth noting that the researchers' results may not be fully comparable, due to

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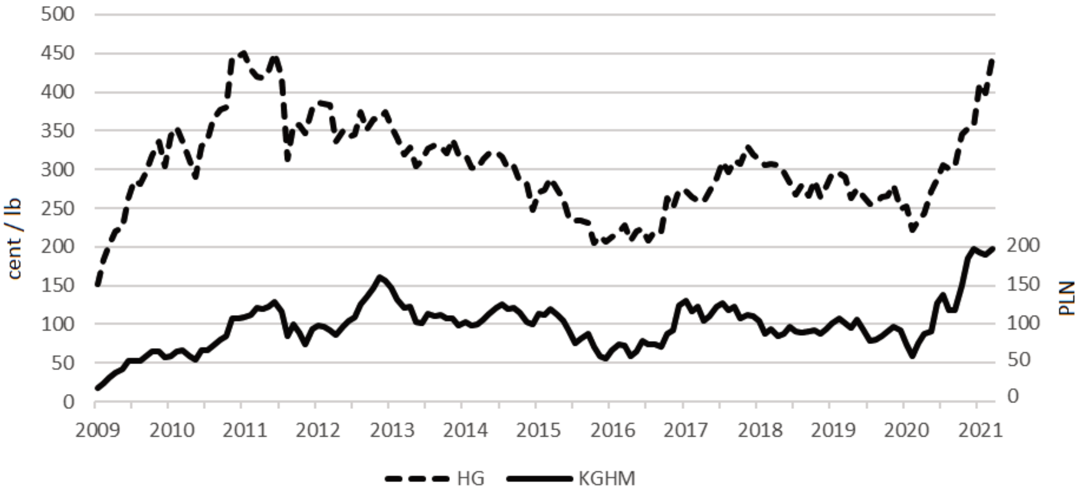
different methodology and reference to different countries and research periods (Mamcarz, 2018). In addition, the result of the analysis may be an apparent regression phenomenon if the property of stationarity and cointegration is ignored (Demidova-Menzel and Heidorn, 2007; Sindhu, 2013). The problem does not occur in the analysis of return series due to their zero degree of integration.

3. METHODOLOGY AND DATA

To the author's best knowledge and the literature review conducted, in this case cointegration modelling might provide an estimation of the long-term equilibrium of variables mentioned. The study was conducted using monthly data covering quotes from August 2012 to April 2021. The data used in the empirical part of the article has a monthly frequency. Designations of the variables specific to the time series under study are as follows:
$KGHM_t$ – monthly quotation of KGHM shares during period *t* on the Warsaw Stock Exchange
HG_t – monthly quotation of copper type HG in period *t* (COMEX exchange, cent per lb).

Figure 1
Time series of KGHM quotations and HG copper prices at levels



Source: own elaboration.

The initial modeling step when analyzing time-series data is to determine the degree of integration. In this case, popular tests such as ADF (Augmented Dickey-Fuller test), PP (Phillips-Perron test) or KPSS (Kwiatkowski-Phillips-Schmidt-Shin test) can be used. The Augmented Dickey-Fuller (ADF) test was based on LS (the Least Squares method) estimation of the auxiliary regression, the basic form of which in the version without a constant and deterministic trend can be written as (Maddala, 2006):

$$\Delta y_t = \delta y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-1} + \varepsilon_t, \qquad \text{Eq. (3.1)}$$

where: δ and γ – parameters estimated by the Least Squares method, k – number of lags, Δy_{t-1} – value of the series of first differences of the studied phenomenon in period $t-1$, ε_t – random component.

The null hypothesis of the ADF test assumes non-stationarity of the y_t series caused by the presence of a unit root. While, the alternative hypothesis of the ADF test assumes stationarity of the time series under study.

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The analysis also referred to the Granger causality test. Causality in the Granger sense is written as $x \rightarrow y$ and involves testing whether the variable x can be removed from that part of the VAR (Vector Autoregression) model (Sims, 1980), which describes y (Granger, 1969). Dependency testing, in the Granger sense, allows one to determine whether the highlighted variable is the cause of changes in the value of the other variable. The approach assumes that variable x is the cause of changes in variable y if the values of variable y can be more accurately predicted by using past values of variable x than without taking them into account. Thus, the Granger test was developed based on a system of equations:

$$y_t = c_1 + \sum_{i=1}^{k} \beta_{1i} y_{t-1} + \sum_{j=1}^{k} \beta_{1j} x_{t-j} + u_{1t}$$

$$x_t = c_2 + \sum_{i=1}^{k} \beta_{2i} x_{t-1} + \sum_{j=1}^{k} \beta_{2j} y_{t-j} + u_{2t}, \qquad \text{Eq. (3.2)}$$

where: y_t – values of variable y in the current period, x_t – values of variable x in the current period, β – structural parameters of the model, u_t – random component of the model.

An interesting issue in this case is causality testing for non-stationary or cointegrated variables. Quite an easy way to deal with this is to use the procedure proposed by Toda and Yamamoto (1995). The procedure is based on the VAR model despite of non-stationarity or cointegration variables. In addition, the cointegration Engle-Granger procedure allows the determination of a single cointegration vector that is consistent with economic theory. It uses a regression equation $\ln y_t = \beta \cdot \ln x_t + c + \xi_t$, which describes the equilibrium relationship between the x and y variables (Charemza and Deadman, 1997). The procedure can be used for variables integrated to the first degree (Enders, 1995). Preliminary estimation of the equilibrium relationship $\ln y_t = \beta \cdot \ln x_t + c + \xi_t$ by the least squares method is used to learn the value of the residual component. Only diagnosing the stationarity of the residual component ξ_t determines the presence of a cointegrating vector. Such a vector describes the long-term relationship between variables. In addition, based on Granger's representation theorem for cointegrated variables, it is possible to estimate a short-run ECM (error correction model) of the form: $\Delta \ln y_t = \alpha \cdot \Delta \ln x_t - \gamma \cdot \xi_{t-1} + c$.

The critical discussion of the chosen method should be included here. The Engle-Granger procedure limits the researcher to the possibility of knowing one cointegrating vector the most (Gajda, 2004). Moreover, it will only be the cointegrating vector that corresponds to the a priori accepted economic theory (Welfe, 1998), which was the basis for the division of the variables under study into exogenous and endogenous.

The Engle-Granger method is an interesting tool for studying time series cointegration, but it has several shortcomings:

1. The procedure requires a two-stage estimation. First, the residuals resulting from the long-run equilibrium relationship are generated, and then the same residuals become the basis for the second regression. Thus, if the regression of the long-run relationship has been subject to an error then this error will be transferred to the second stage of the procedure.
2. Estimating the long-run equilibrium relationship requires making a theoretical assumption about which of the variables under study is the dependent variable and which is the independent variable. Particularly in the case of a limited sample size, this can lead to the paradox that for a pair of specific variables, the long-run relationship will be confirmed or not depending on the scenario of defining the dependent and independent variable. According to asymptotic theory, the problem will not occur with sufficiently large samples, although the practice of economic research often struggles with the phenomenon of limited sample size.

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3. Testing for cointegration seems clear in the case of a pair of variables. In this case, the long-run equilibrium relationship can be unambiguously determined or the absence of cointegration can be found. In a situation where the number of explanatory variables is greater than 1, the whole analysis gets complicated. If the model contains »*n*« variables, there can be from 0 to $(n-1)$ linearly independent cointegrating vectors. Meanwhile, the Engle-Granger procedure does not even offer a test of the number of such vectors, much less determines them.

The properties of the Engle-Granger procedure and its limitations make it a good idea to verify the cointegration test also on the basis of the Johansen procedure (Johansen, 1988). It is an approach commonly used when modeling multiple cointegrated variables. Jahansen's approach, together with VAR and Vector Error Correction Model (VECM) models, allows cointegration to be analyzed multidimensionally. As mentioned earlier, while Engle-Granger's approach is based on the assumptions of economic theory and allows the determination of a maximum of one cointegrating vector, Johansen's procedure allows the testing of a total number of such vectors. In addition, in the Johansen approach it is possible to determine all cointegrating vectors. The Johansen approach uses vector autoregressive VAR models.

A VAR model without boundary conditions and deterministic components is written in a compact form using the equation (Sims, 1980):

$$Y_t = \sum_{i=1}^{k} A_i Y_{t-1} + E_t, \qquad \text{Eq. (3.3)}$$

where: Y_t – is a vector containing n variables of the model; A_i – is a matrix of parameters with lagged variables; E_t – is a vector of random components.

Using the so-called cointegrating transformation, the VAR model can be brought to the form of the VECM model (Brooks, 2012):

$$\Delta Y_t = \sum_{i=1}^{k} \Gamma_i \Delta Y_{t-i} + \Pi Y_{t-k} + U_t, \qquad \text{Eq. (3.4)}$$

where: $\sum_{i=1}^{k} \Gamma_i \Delta Y_{t-i}$ – is the short-run part of the model (deviations from the long-run path); ΠY_{t-k} – is the long-term part of the model containing long-term trends; U_t – is a vector of random components.

If there are n variables in the vector Y_t and the row of the matrix Π is equal to $0 < r < n$ then the matrix Π can be decomposed as follows:

$$\Pi = \alpha \beta^T, \qquad \text{Eq. (3.5)}$$

where: α – adjustment matrix, β – matrix of long-term trends.

The matrix β consists of cointegrating vectors that, when normalized, can be interpreted as long-term parameters. Assuming that the vector Y_t consists of n variables then the matrix Π will be of dimension $(n \times n)$, so its row can be equal to at most n. Based on the Granger's Representation Theorem (Engle and Granger, 1987; Johansen, 1991) it can be assumed that:

1. If the row of the matrix Π is equal to n then all of the studied variables included in Y_t are stationary. Then it is advisable to estimate the traditional VAR model for the levels of the variables. In addition, if the row of the matrix Π is equal 0 then the traditional VAR model for

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variable increments is recommended. All components of the matrix Π are equal 0 therefore the representation in the form of error correction does not exist.

2. If the row of the matrix Π is equal to $0 < r < n$ then there exists a representation Π such that $\Pi = \alpha\beta^T$ within which the matrices α and β have dimension $(n \times r)$. Then the matrix β is a cointegration matrix that has the property that $\beta Y_t \sim I(0)$, when $Y_t \sim I(0)$. In this case, a VECM model is built, which is a transformation of the VAR model through the use of a cointegrating transformation. Within a VAR model with n variables, there can be at most $r = n - 1$ cointegrating vectors.

An important practical issue in this case is determining the appropriate number of delays for VAR or VECM models. One possibility in this regard seems to be the use of lag length tests (Kusideł, 2000), which, however, do not always lead to clear conclusions. Often, a more effective technique for determining lags k is to assess stability of the estimates of the adjustment coefficients included in the matrix of α. Then, the smallest possible value of lag k is taken, which provides similar values for the coefficients of the matrix α. For VAR models, the lag k should be adequate for the length of the response (i.e., adjustment) to the deviation from the long-run trajectory according to the principle of the error correction mechanism (ECM). Theoretically, the adjustments are assumed to occur over a relatively short period. However, lags k should be large enough to eliminate possible autocorrelation of the model's random components. Thus, in general, long-term trends should not depend on lags k in the model, but already short-term adjustments will depend on k. In the study presented here, lags of 1 to 2 were assumed for the VECM model. It should be noted that any reduction in the dimension of the estimated model has a beneficial effect on reducing the number of its parameters required for estimation. The critical discussion of the Johansen approach focuses on the application of the VAR / VECM model. In this case every variable is specified as endogenous and some researchers point out the non-theoretical approach.

4. RESULTS

4.1. Evaluation of the stationarity of the studied time series

In verifying the stationarity of the time series, it turned out that both series are non-stationary in levels, but have stationary first differences. The property should be described as their integration on first differences. Thus, the application of the differential filter made it possible to transform both non-stationary processes into stationary series. Table 1 contains the results of ADF tests for the levels and differences of the KGHM and HG copper quotation series.

Table 1
Stationarity test of KGHM and HG copper quotations on levels and increments

	KGHM		HG	
	t-Statistic	**Probability**	**t-Statistic**	**Probability**
ADF level	–1.653772	0.4517	–1.224534	0.6616
Test (ADF level) critical values 5% level	–2.889474		–2.889753	
ADF 1st difference	–8.035865	0.0000	–5.415778	0.0000
Test (ADF 1st difference) critical values 5% level	–2.889474		–2.889753	

Source: own elaboration based on test checks and critical values obtained in Eviews program.

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It is worth noting that although Table 1 contains the critical values of the unit root tests for a significance level of 0.05, analogous conclusions were obtained even for a much lower significance level of 0.01. Both time series are integrated on the first differences.

4.2. Testing the causality of KGHM and HG copper quotes in the Granger sense

Granger tests include levels of the variables in the VAR model according to the Toda and Yamamoto procedure (1995) with the Wald test statistic. In this case the Wald test is asymptotically chi-square distributed. The procedure is based on the VAR model despite of variables non-stationarity or cointegration. The null hypothesis of the Granger causality test assumes the absence of causality, in which case all β parameters are equal to zero. The alternative hypothesis, on the other hand, assumes that if causality is present then some β_i or β_j parameters are different from zero.

Table 2
Granger causality test for variables $KGHM_t$ and HG_t

Null Hypothesis:	Obs	Chi-sq	Prob.
KGHM does not Granger Cause HG	102	4.8009	0.1870
HG does not Granger Cause KGHM		12.8149	0.0051

Source: own elaboration based on a causality test conducted in Eviews software.

The results of the test contained in Table 2 indicate the existence of significant causality in the Granger sense of the changes in the levels of HG copper quotations relative to the changes in the levels of KGHM stock quotations. It is indicated by the relatively high value of the F statistic = 12.8149 and the low value of the empirical significance level (Probability = 0.0051), as a consequence of which the null hypothesis of causality in the Granger sense test should be rejected. The opposite relationship is not substantiated, that is, changes in the KGHM's stock price are not the cause of changes in the HG copper price. In this case, there is no basis for rejecting the null hypothesis of the Granger test performed (Probability = 0.1870).

4.3. Study of the long-term relationship of KGHM and HG copper quotations

Table 3
Long-term relationship model for natural logarithms of $KGHM_t$ and HG_t variables

Included observations: 106				
Variable	Coefficient	Std. Error	t-Statistic	Prob.
LN(HG)	1.367883	0.080279	17.03912	0.0000
C	–3.100294	0.454077	–6.827685	0.0000
R-squared	0.736263	Mean dependent var		4.633575
Adjusted R-squared	0.733727	S.D. dependent var		0.260688
S.E. of regression	0.134520	Akaike info criterion		–1.155527
Sum squared resid	1.881933	Schwarz criterion		–1.105274
Log likelihood	63.24293	F-statistic		290.3317

Source: own development in Eviews software.

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The model in Table 3 shows a correct fit to the empirical data. The adjusted coefficient of determination in this case was more than 0.73. It means that about 73 percent of thevolatility of KGHM stock prices was described by the volatility of HG copper prices. In the estimated model, both the long-run relationship parameter and the constant appear to be significantly different from zero, but it should be remembered that the indications of Student's t-test or F-test can be misleading due to the non-stationarity of the time series under study at levels. The preliminary form of the long-run relationship equation should be written as:

$$\ln(KGHM_t) = 1.367883 * \ln(HG_t) - 3.100294. \qquad \text{Eq. (4.1)}$$

Thus, potentially, a 1 percent increase in the HG copper quotation results in an average increase in the KGHM stock price quotation of 1.367883 percent over the long term, assuming ceteris paribus. Whether the indicated relationship is true can only be determined by testing for stationarity at the levels of the residual component implied by the estimated model.

Table 4
Stationarity test of the residual component from long-run relationship model at levels

	Residuals	
	t-Statistic	**Probability**
ADF level	–3.905480	0.0028
Test (ADF level) critical values 5% level	–2.889200	

Source: own elaboration based on test check and critical value obtained in Eviews program.

Table 4 confirms the stationarity of the tested residual component at levels because the value of the ADF test (–3.905480) is less than the critical value for the significance level of 0.05. Thus, the modeled time series is characterized by a long-run relationship called cointegration. Moreover, according to Granger's representation theorem, there should be a short-run relationship in the case with an error correction mechanism.

As can be seen in Table 5 for the studied set of variables, the short-run relationship with the error-correction mechanism is also true. The ECM model has a moderate fit to the empirical data (i.e., the adjusted coefficient of determination was 0.429852). It should be noted, however, that this is a model estimated for incremental variables, where this type of model does not usually

Table 5
ECM short-run relationship model for logarithm increments of $KGHM_t$ and HG_t variables

Included observations: 105 after adjusting endpoints				
Variable	**Coefficient**	**Std. Error**	**t-Statistic**	**Prob.**
D(LN(HG))	1.231513	0.144465	8.524654	0.0000
Z(–1)	–0.235451	0.062487	–3.768016	0.0003
R-squared	0.435334	Mean dependent var		0.006065
Adjusted R-squared	0.429852	S.D. dependent var		0.112544
S.E. of regression	0.084980	Akaike info criterion		–2.073933
Sum squared resid	0.743828	Schwarz criterion		–2.023381
Log likelihood	110.8815	F-statistic		79.40865

Source: own development in Eviews software.

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have high coefficients of fit to empirical data. In this case, about 43 percent of the variation in the incremental natural logarithm of KGHM's stock price was explained by the model built, namely the variation in the incremental natural logarithm of HG copper prices and the one measurement period lagged residual component derived from the long-run relationship. The coefficient of short-run elasticity in the case was 1.231513. On the other hand, the adjustment parameter has a correct value less than zero and at the same time less than unity in absolute value. Thus, the studied system has the ability to return to equilibrium. The short-run model with the error correction mechanism in the final version was estimated without a constant value. Hence, the short-run relationship was written as:

$$\Delta\ln(KGHM_t) = 1.231513 * \Delta\ln(HG_t) - 0.235451 * z_{t-1}, \qquad \text{Eq. (4.2)}$$

where: z_{t-1} – the value of the residual component in period $t-1$ from the long-run relationship model.

Based on the estimated model, it can be concluded that a 1 percentage point increase in the growth rate of the HG copper quotation results in an average increase in the growth rate of the KGHM stock price of 1.231513 percentage points, in the short term, assuming ceteris paribus. In addition, the adjustment parameter of –0.235451 suggests that more than 23% of the previous period's imbalance will be removed in the current period as a result of the ECM's error correction mechanism.

Table 6 contains results of the Johansen cointegration test, which confirms the conclusions of the Engle-Granger procedure. Thus, for the studied variables $KGHM_t$ and HG_t there is a long-run relationship described by a single cointegrating vector. The Johansen cointegration test proceeds sequentially that is: in the first stage, the value of the trace test (Trace Statistic) is 15.79454, which is greater than the critical value (15.49471). This means that the number of cointegration vectors is greater than 0. In stage two, the trace test (Trace Statistic) is 2.657169 which is less than the critical value (3.841466). Hence, the number of cointegration vectors equals one.

Table 6
Johansen test of the number of cointegration vectors for the variables $\ln(KGHM_t)$ and $\ln(HG_t)$

Series: LN(KGHM) LN(HG)
Lags interval (in first differences): 1 to 2
Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.119748	15.79454	15.49471	0.0451
At most 1	0.025468	2.657169	3.841466	0.1031

Trace test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
** MacKinnon-Haug-Michelis (1999) p-values

Source: own development in Eviews software.

Thus, the results of the Johansen test indicate the existence of a long-term relationship between KGHM stock prices and HG copper prices. Hence, it can be assumed that the Engle-Granger procedure carried out earlier for the study also led to significant conclusions. The next step in the analysis will be to see whether the two cointegration test procedures yield similar estimates of the coefficient of long-run elasticity.

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Based on the VECM model and the $\Pi = \alpha\beta^T$ matrix, a cointegration vector was estimated to describe the long-term relationship between the $KGHM_t$ and HG_t variables. Its parameter after normalization is included in Table 7.

Table 7
Long-term relationship between $KGHM_t$ and HG_t variables based on the Johansen procedure

Normalized cointegrating coefficients (standard error in parentheses)	
Ln(*KGHM*)	**Ln(*HG*)**
1.000000	–1.318959
	(0.21062)

Source: own compilation based on calculations in Eviews software.

The analytical form of the cointegrating relationship can be expressed by the equation (Syczewska, 1999; Majsterek, 2014):

$$\ln(KGHM_t) = 1.318959 * \ln(HG_t). \quad \text{Eq. (4.3)}$$

Based on the estimated equation, it can be concluded that a 1 percent increase in the quoted price of HG copper results in an average increase of 1.318959 percent in KGHM's stock price over the long term, assuming ceteris paribus (Gajda, 2004; Borkowski et al., 2007). The impact of HG copper quotations on KGHM stock prices should be considered significant in the long term.

At the same time, recall that the estimate of the coefficient of the long-run relationship obtained from the two-stage estimation according to the Engle-Granger procedure was 1.367883. Thus, not only the signs, but also the values of the obtained parameters of long-run elasticity turned out to be similar (1.318959 and 1.367883). A slightly higher parameter value was recorded for the Engle-Granger procedure relative to the Johansen procedure. Referring to the theoretical assumptions of the study and the literature review, it shoud be noted that the long-term equlibrium of KGHM share prices and the HG copper is a unique case of the relationship dependence between the stock and commodity markets. Most publications cited in the literature review section concern the study of volatility (Creti et al., 2013; Kang et al., 2016; Vardar et al., 2018; Zevallos and Carpio, 2015). As it has been mentioned before the problem of volatility clustering for stock market data, is one of the accentuated properties of financial time series. From the cointegration point of view short-term deviations are really important too, therefore the research is in compliance with papers cited in the literature review. In this case short-term deviations from the long-term equlibrium are vital to identify the error correction mechanism. It is worth noting that the mentioned before publication by Fijałkowski (2011) contains quite different research results. This is understandable because of quite different methodology and periods analyzed. Both studies were focused on the examination of the dependence between KGHM share and copper prices. However, it should be noted that the time series modelling requires procedures coinciding with the so-called new econometrics (Charemza and Deadman, 1997). The aim of the article was to show the application of such procedures in relation to time series of KGHM shares and copper prices. In this context, it seems that this intention has been achieved.

5. CONSLUSIONS

Concluding the conducted study of the cointegration relationship for the $KGHM_t$ and HG_t variables, it should be emphasized that financial data including stock quotes are most often non-stationary. Their analysis is hampered by the presence of stochastic trends and high volatility.

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Under the conditions, techniques for modeling integrated variables, such as the Engle-Granger and Johansen procedure, seem to be helpful. Although they are characterized by slightly different numerical properties, the example of the article shows that they can lead to very similar conclusions. Let's point out that the analysis presented here, by design, concerned only two variables $KGHM_t$ and HG_t therefore Johansen's multivariate method did not present its full capabilities. At the same time, a rather large set of stock quotes covering almost a decade made it possible to capture the common trajectories of the studied time series. Of course, the Engle-Granger method used in the first instance has several limitations mentioned in the body of the article, so its use should be confirmed by a more advanced method (the Johansen procedure). This is what happened in this case, where in the end very similar values of long-run equilibrium parameter estimates were obtained for both methods. Thus, on this basis, it is necessary to confirm the main hypothesis formulated at the beginning of the study, that is, in the case of KGHM Polska Miedź operating within the mining industry, the quotation of its shares is long-term dependent on the quotation of HG copper prices. Limitations of the study conducted include the inclusion of only two variables in the analysis. However, given the fairly unambiguous results on common long-term trends, expansion of the presented models with new variables will become a direction for future research. The purpose of the paper is to try to explain the long-term price volatility of KGHM shares. In the context, it seems that the intented purpose has been achieved. The main findings are related to the obtained cointegration vector that describes the long-term dependence between the variables examined. According to the research results there are quite sensitive values of KGHM share prices to changes of copper prices. This can be said based on the value of the long-term parameter around 1.3. The significance and relevance of the results are related to many areas with economic consequences. The presented modeling results can be an inspiration for investors and experts in company valuation. In addition, the analysis of highly volatile stock market data can contribute to a better understanding of processes on the WSE. In addition, the conducted empirical study enables making careful comparisons with the previous publications in the literature of time series cointegrating of the stock and commodity markets.

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Corporate Philanthropy in Shaping the Financial Efficiency of Cooperative Banks in Poland – Empirical Research

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ABSTRACT

The aim of the study is to establish and analyze the relationship between expenditure on corporate philanthropy and financial performance of cooperative banks in Poland. The study covered 70 cooperative banks, where the level of expenditure on social activities was above the average in the sector. The research was carried out using the Generalized Method of Moments for the years 2013–2020. Studies have shown positive importance of corporate philanthropy in shaping financial efficiency. The research also indicated a significant shift in the time between spending on social activities and the financial results obtained. The obtained results indicate an area that has not been thoroughly analyzed: the impact of spending on social activities in the context of shaping the financial efficiency considering time shifts. This knowledge may result in more intentional creation of the efficiency policy by management, considering corporate philanthropy, Corporate Social Responsibility, or the introduction of appropriate regulatory changes for the assessment of financial institutions by responsible entities.

JEL classification: A13, G41, M14, G21

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1. INTRODUCTION

In the first and second decades of the 21st century, discussions on the essence of the functioning of enterprises in the economy intensified in world markets. Changing consumer expectations and legislation is placing increasing emphasis on the protection of human rights and the environment, which forces companies to respond to the question of whether they are to be focused solely on profit maximization or their actions should take into account the subjective element of a human being and their needs, including non-business ones. Simultaneously, it is important that combining market orientation and focusing on achieving social goals requires enterprises to efficiently manage their resources, taking into account the implementation of social missions (Bruder, 2021). It is also worth noting that despite global crises, corporate philanthropy has significantly maintained its momentum as a growing phenomenon of global importance (Gautier and Pache, 2015).

In literature, we can observe the focus on voluntary actions related to the transfer of funds to individuals or non-governmental organizations, and combining business and social activities in such a way that they create value and increase the company's efficiency (Bruton, Ahlstrom and Obloj, 2008). In the context, in literature, one can find examples of research indicating a connection between interest in the implementation of social activities and the desire to improve the financial efficiency of enterprises.

Research on the relationship between corporate philanthropy carried out by companies and the rate of return on their stock shares (Ferrell, Liang and Renneboog, 2016), financial efficiency (Lins, Servaes and Tamayo, 2017) and financial risk (Kim, Li and Li, 2014) does not lead to unambiguous conclusions (Wu and Shen, 2013).

In addition, as indicated by researchers, an important element of research on the impact of social activity on the effectiveness of banks is the consideration in the analysis of sub-components used to measure social activity, such as environmental, social, and governance, because, in their opinion, they may have a different significance for the development of banks' financial performance of banks (Wu and Shen, 2013). From the perspective, corporate philanthropy, defined as the practice of organizations donating resources and funds to support social and environmental causes (Smith, 1994), is also undergoing a significant transformation. In recent years, there has been a growing awareness of interconnections between corporate philanthropy and Environmental, Social, and Governance (ESG) principles.

ESG represents a set of criteria that assess an organization's performance in three key dimensions: environmental sustainability (E), social responsibility (S), and ethical governance (G). The principles have gained importance as frameworks for evaluating long-term sustainable development and ethical corporate practices (Aluchna, Roszkowska-Menkes and Kamiński, 2022). Investors, consumers, and stakeholders now perceive ESG factors as fundamental indicators of an organization's commitment to responsible business practices. Therefore, ESG factors should be considered when making decisions by financial institutions.

Nonetheless, research conducted by the European Banking Authority (EBA) indicated that among the key factors motivating banks to adopt sustainable development strategies there are ethical considerations, perceived business opportunities, and the demands of customers and investors. Significant motivations are also related to new risk factors and expected regulatory changes (Coleton et al., 2020). According to Marcinkowska (2022), ESG factors must be incorporated into the management of financial institutions, regardless of the institutions' perception of sustainable development issues and their willingness to engage in them.

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At the same time, one of the least frequently analyzed is one of the most important examples of combining business and social activities in their functioning, locally operating financial institutions (McKillop et al., 2020). It is a special type of enterprise due to the Rochdale principles, included in the operating strategy of the type of enterprise, which impose on cooperatives the obligation to conduct social and cultural activities for their members, their families, and the local social environment (Tremblay, Hupper and Waring, 2019). They also unequivocally indicate that the activity should not only be incidental to the conducted economic activity, aimed at satisfying the economic needs and interests of its members, but should be parallel (Rand and Nowak, 2013). However, we do not know how locally operating financial institutions implement the postulates in practice because so far, there has been little research on their social activity (Taylor and Goodman, 2020). The research area of financial institutions focuses mainly on the financial efficiency of the institutions without paying special attention to their social dimension (Ferreira, 2021) and, as Belasri, Gomes and Pijourlet (2020) indicate, the analysis of such a complex institution what a bank is, requires a multidimensional approach that also takes into account social activity. Moreover, some research is fragmentary and usually involves a review of the literature in the area in question. There is little empirical research linking financial and social spheres, especially in terms of the differentiation of the level and scope of social banking activities depending on their financial effectiveness (Aramburu and Pescador, 2019).

Issues related to social activities conducted by cooperative banks are particularly important because of their importance in the European banking sector (Clark, Mare and Radić, 2018). The average share of the institutions in domestic markets in Europe increases year by year, exceeding 23% (*World Cooperative Monitor – Report 2021*, 2022), and in Austria, Germany, and Finland, it exceeds 30% (*The Co-operative difference: Sustainability, Proximity, Governance*, 2021). They also have a significant share in the market of loans and deposits for the SME sector, which is related to the fact that the banks are much smaller entities than commercial banks (Clark, Mare and Radić, 2018) and because of their specificity, they are also an important link in mitigating the effects of monetary tightening (Ferreira, 2021) and in the development of local markets. Moreover, the financial situation of cooperative banks is crucial for households, small and medium-sized enterprises, and farmers, as they are often important providers of financial services in rural areas and small towns (Kozłowski, 2016). They are also examples of financial institutions directly involved in the functioning of local markets and communities through active social activities. The aim of the study is to establish and analyze the relationship between expenditure on corporate philanthropy and financial performance of cooperative banks in Poland. Therefore, recognizing the importance of cooperative banks' social activity is an important issue, both in theory and practice, and the key value in this regard is the answer to the question of whether this activity strengthens or weakens the financial efficiency of the entities.

2. LITERATURE REVIEW AND HYPOTHESES DEVELOPMENT

Previous analyses of the banking sector mainly concerned quantitative data analyses or changes in the legal regulations of this part of the market (Mare and Gramlich, 2021). Nevertheless, a cooperative bank is both a bank and a cooperative; therefore, it is a good example of an institution that should not be perceived solely through the prism of financial efficiency in the strict sense because it is at the same time an element of modern financial markets and an integral part of local communities.

In literature, one can find a variety of studies indicating motives for undertaking corporate philanthropic activities by enterprises, among which three basic trends can be distinguished. The first concerns altruistic motives, they are inspired by social awareness and altruism, and their main goal is to improve social welfare, even if it has little or no impact on a company's profits (Sánchez, 2000). The second trend concerns the social pressure exerted on organizations

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resulting from certain expectations on the part of stakeholders legitimizing the functioning of a given organization (Gao, 2011). The third trend indicates that such activities are undertaken because of the benefits the enterprise can obtain (Ji, Tao and Deng, 2021). Among them, we can distinguish the enhancement of brand reputation and recognition (Pan et al., 2018); increasing customer loyalty and trust (Pivato, Misani and Tencati, 2007), strengthening the market position and strategic potential (Porter and Kramer, 2003), establishing political connections (Su and He, 2010) and improving financial efficiency (Belasri, Gomes and Pijourlet, 2020). An example of entities affected by the three trends may be cooperative banks, which, as entities strongly associated with communities and local markets, often operate under environmental pressure. The environment, perceiving them as cooperatives, puts a lot of pressure on undertaking social activities for their benefit, but cooperative banks, as financial entities, should also consider the financial goal of increasing the efficiency in their activities. Keeping this in mind, the following hypothesis is formulated:

H1: The social activity of a financial institutions contributes to the increase of its financial efficiency.

A manifestation of social activity carried out by cooperative banks supports local social initiatives by spending funds for social purposes. The expenses are referred to as corporate philanthropy, which has been gaining importance in recent years and has been the subject of numerous studies (Zhang, Chen and Mo, 2016; Christou, Hadjielias and Farmaki, 2019; Li et al., 2021). Despite the many available studies, the issue is still not fully explored and requires further research (Zhang, Chen and Mo, 2016). According to literature, corporate philanthropy consists of activities resulting from voluntary commitment to the good of others or to the common good by providing funds, goods, or services (Li et al., 2021). Therefore, the activities are perceived as a manifestation of altruism resulting from goodness, and are not aimed at obtaining additional benefits. However, studies available in literature suggest that activities related to social activity undertaken by enterprises are not only altruistic, but are also motivated by the desire to obtain tangible benefits (Ji, Tao and Deng, 2021). It is worth noting that the activities have an impact not only on the society and environment in which the company operates, but also on the organization itself that undertakes the activities (Chen and Lin, 2015).

In literature, one can find examples of research on the relationship between corporate philanthropy and effectiveness of companies; however, as has already been noted, results are not unequivocal. Some studies indicate that corporate philanthropy has a positive impact on companies' financial efficiency of companies (Su and Sauerwald, 2018), while other studies indicate that the relationship is negative or insignificant (Masulis and Reza, 2015). The discrepancies may result from a different approach to conducting research. Because cooperative banks are obliged to undertake social activities, it is worth considering the extent to which the activities affect their financial effectiveness and whether the interaction between business and social activities changes over time, as in the case of factors such as liquidity or bank asset growth, whose impact on bank efficiency is lagged (Alabbad and Schertler, 2022).

Profit maximization can be a primary motivation for corporate executives to engage in corporate philanthropy, in accordance with the perspective of enlightened self-interest. In the view, corporate philanthropy becomes part of the profit maximization equation by creating goodwill and enhancing the company's image and reputation (Baumol, 1970; Galaskiewicz, 1985; Stendardi, 1992; Shaw and Post, 1993; Abzug and Webb, 1996). It is expected that such actions will enhance the company's performance. Many CSR strategies suggest that company profitability can be a fundamental motivation for philanthropic activities (Gautier and Pache, 2015).

Accordingly, the following hypothesis is formulated:

H2: Financial efficiency in a given year is significantly dependent on the expenditure on social activity in preceding years.

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3. RESEARCH METHODS

3.1. Research sample

Banks differ from each other not only in terms of organizational and legal forms, management methods, ownership, and capital structure but also in terms of business goals (Dilger, Konter and Voigt, 2017). However, increasingly all financial institutions attempt to combine strategic objectives related to maximizing shareholder value with corporate social responsibility. They focus on maximizing value for a larger and more diverse group of stakeholders representing different interests, such as environmental concerns, as well as markets and local communities. They are less inclined to pursue short-term profit maximization and instead prioritize maintaining a stable and sustainable level of profitability, which is a necessary condition for ensuring the bank's continuity, security, resources for further development, and simultaneous achievement of non-financial goals. To maintain comparability of banks, we selected cooperative banks whose level of spending on social activities during the years 2013–2020 was above the arithmetic average of the amount of expenditure on social activities in the studied group. In the way, the analyses included banks actively involved in social activities, and their level of expenditure was not classified as incidental or marginal. According to a state dated December 31, 2020, 70 banks met the above condition, they constituted nearly 13.21% of all cooperative banks operating in Poland and the share in total assets of whole sector reached 16,87% as of the date of the analysis. The analysis was limited to 2013–2020 due to the occurrence of emergencies in 2021–2022, such as the COVID-19 pandemic or Russia's aggression against Ukraine, which could significantly distort the results of the analysis.

Analyses were conducted using dynamic panel models (Arellano and Bond, 1991) to explain selected measures of financial efficiency of cooperative banks. It made it possible to determine the factors influencing efficiency of banks, taking into account the time shifts between expenditure on social activities and the resulting financial benefits, as well as the elimination of the causal effect, the occurrence of which could lead to erroneous conclusions (Leszczensky and Wolbring, 2022). Dynamic models were estimated using the Generalized Method of Moments (GMM), which enables the model parameters to be estimated directly from the moment conditions.

3.2. Dependent variables

Profitability of total assets was used as the main measure of the efficiency of cooperative banks, which made it possible to compare the results of cooperative banks considering the entire scope of their activities and to eliminate the effect related to the diversification of the structure and size of equity (Li et al., 2021). Returns on total assets were calculated as the ratio of net income to total assets in general (Ramzan, Amin and Abbas, 2021). Additionally, to verify durability of the results (proxy variable), the rate of return on equity (ROE) [%] was used, which was calculated as the quotient of the net financial result (net income) and bank equity (total equity) (Finger, Gavious and Manos, 2018). The inclusion of ROE as an additional measure of the effectiveness of a cooperative bank allows for the omission of the risk related to the bank's off-balance sheet activities, which may be included in the ROA (Nizam et al., 2019), as well as the recognition of the bank's financial security level, which is a derivative of its equity (Folorunsho Monsuru and Adetunji Abdulazeez, 2014).

3.3. Independent and moderating variables

Expenditure on corporate philanthropy (CP) was the key explanatory variable that was used in the model and it was the sum of expenditure of a cooperative bank over a period of one year on all kinds of bank subsidies for investments and social projects, the recipients of which are members of local communities (Idasz-Balina et al., 2020). Such an approach to spending on social activities

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was included in research on the long-term effects of banks' social activities and to eliminate one-off donations for social purposes resulting from random events (Li et al., 2021). Thus, the natural logarithm of the sum of expenses of a cooperative bank on social activities during the calendar year (L_CP) was used as an independent variable. The use of such an approach made it possible to include in the research the factors related to the impact of the size of the cooperative bank on the value of funds allocated to social activities. Larger entities tend to be more active in taking up social activities than smaller entities.

The natural logarithm of the number of supported community initiatives within corporate philanthropy (L_NSI) is used as a proxy variable for social activity. The variable indicates the scope of social activity and its potential range in the local environment, which is particularly important for cooperative banks operating in local markets (Idasz-Balina et al., 2020).

Due to the nature of the research and the probability of delays in the impact of variables determining the social activity of cooperative banks on their financial efficiency, one-year ($t-1$) and two-year ($t-2$) delays were taken into account.

3.4. Control variables

In the study, control variables were used, which can impact the functioning of the bank and its financial efficiency.

Loan growth (L_G) leads to an increase in income and interest commissions in a bank, which translates into an increase in its efficiency with a simultaneous decrease in the total capital ratio that determines the level of the bank's financial security (Abedifar, Molyneux and Tarazi, 2018).

Deposit growth (D_G) also affects bank profitability and security. This is due to the fact that deposits are the main source of obtaining funds by banks, which then charge the risk on their own account, and the cost of obtaining them affects the level of interest costs that are the basis for determining the bank's financial efficiency (Nizam et al., 2019).

Total capital ratio (TCR) determines the level of financial security of a bank because the ratio shows whether the bank is able to cover the minimum capital requirement for credit, operational, and other risks with its own funds; the higher the value of the ratio, the greater the security of the bank's shareholders, which, as indicated by Paroush and Schreiber (2019), translates into a bank's financial efficiency, with a simultaneous negative correlation between the level of the capital ratio and ROA and ROE.

The non-performing loans (NPL) ratio measures a bank's exposure to the risk of loans and credits that are out of date as well as the quality of its borrowers. This ratio is particularly important for the level of a bank's financial efficiency because its low quality translates into a direct decrease in the level of financial efficiency and a decrease in the bank's financial security (Kil, Ciukaj and Chrzanowska, 2021).

The number of members of a cooperative bank (L_NBM) determines its ownership structure, which is the main mechanism of corporate governance (de Haan and Vlahu, 2016). However, research shows that a concentrated ownership structure may play a significant role in shaping financial results of economic entities. The ownership structure is particularly important in the case of cooperative banks, which, owing to their specific nature resulting from the cooperative nature of their activities, may significantly affect local markets and communities. Moreover, research indicates that ownership structure may have a significant impact on bank efficiency of banks (Huang, 2020).

Location of a bank (LOC) is also important for its ability to generate financial results (Degl'Innocenti, Matousek and Tzeremes, 2018). In the case of cooperative banks, which mainly operate in local markets, it was important to take into account the level of urbanization of the area in which they operate. Therefore, banks were grouped according to the area in which they operated, i.e. rural areas [value 1 assigned], rural-urban [value 2 assigned] and urban [value 3 assigned], in accordance with the Regulation of the Council of Ministers of November 14, 2007 on the introduction

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of the Nomenclature of Territorial Units for Statistical Purposes (NTS) at the NTS-5level of detail. The location of a bank (LOC) is also important for its ability to generate financial results. In the case of cooperative banks, which mainly operate in local markets, it is important to consider the level of urbanization of the area in which they operate. Therefore, banks were grouped according to the area in which they operated, that is, rural areas, rural–urban areas, and urban areas.

In addition to the above-mentioned variables, a variable related to the macroeconomic situation is also used. It reflects the conditions under which cooperative banks operated during the research period. The variable determined the growth of gross domestic product (GDP), which was included in the analyzes due to the fact that it is one of the most important external factors influencing the functioning of banks, and the so far conducted studies have shown that the activity of banks is sensitive to gross domestic product change (Dietrich and Wanzenried, 2014). Shi et al. (2021) also showed that GDP growth has a positive impact on banks' financial efficiency due to the increased demand of enterprises for external financing and the low risk of customer insolvency in good economic times.

Table 1 presents descriptive statistics for the study population, and Table 2 presents the Pearson linear correlation matrix between the analyzed variables. The results indicate validity of the selection of variables for assessing the impact of social activity on the efficiency of cooperative banks.

Table 1
Descriptive statistics

Variable	Mean	Median	Standard deviation	Minimum	Maximum
ROE (%)	7.220	7.140	2.520	0.980	15.500
ROA (%)	0.883	0.868	0.340	0.151	1.950
CP (PLN thousand)	79.900	56.000	67.200	13.900	494.000
NSI (number)	78.300	67.000	58.000	3.000	365.000
L_G (%)	4.310	2.100	7.290	–17.800	38.300
D_G (%)	6.600	5.770	8.470	–20.900	45.100
TCR (%)	16.500	15.200	4.330	10.100	37.500
NPL (%)	4.680	3.720	3.810	0.181	28.500
NBM (number)	2200	1750	1750	149	9760
GDP (%)	3.030	3.800	2.440	–2.500	5.400

Source: Author's calculation.

Table 2
Correlation matrix

	ROE	ROA	L_CP	L_NSI	L_G	D_G	TCR	NPL	L_NBM	LOC	GDP
ROE	**1.0000**										
ROA	0.7508***	1.0000									
L_CP	0.0674*	0.05439*	1.0000								
L_NSI	–0.0088	–0.1348***	0.4590***	1.0000							
L_G	0.0766**	0.0184	–0.0300	0.0038	1.0000						
D_G	–0.0159	–0.1065***	0.0003	0.0609**	0.5903***	1.0000					
TCR	–0.0684**	0.3174***	–0.2018***	–0.1210***	–0.0588*	–0.0322	1.0000				
NPL	–0.2529**	–0.2271***	0.1750***	0.0874**	–0.1154***	–0.0279	–0.1894***	1.0000			
L_NBM	–0.1131***	–0.2447***	0.2208***	0.3630***	–0.0286	–0.0091	–0.2270***	0.0582*	1.0000		
LOC	–0.1358***	–0.0461	0.1256***	0.1581***	–0.0294	–0.0182	–0.0077	0.0612*	0.2995**	1.0000	
GDP	0.0846**	0.0388	0.0065	0.0360	0.2493***	0.3742***	–0.0190	–0.0511	–0.0042	–0.0022	1.0000

Significance level at: * 10%, ** 5%, *** 1%.

Source: Author's calculation.

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The analysis of the Pearson linear correlation coefficients (Schober, Boer and Schwarte, 2018) between financial efficiency and the explanatory variables adopted for the study showed that ROA and ROE were significantly correlated with L_CP, TCR, NPL, and L_NBM. Additionally, in the case of ROA, apart from the variables mentioned above, L_NSI and D_G were significant, and L_G, LOC, and GDP were important for ROE. It indicates that effectiveness of cooperative banks depends on many factors that strictly determine their activities in organizational, financial, and social areas. It is worth paying attention to the correlation between expenditures on social activities (L_CP) and other variables. The results indicated that the level of the expenses was strongly correlated with the bank's financial security level (TCR), asset quality (NPL), the number of bank members (L_NBM), and its location. The number of supported initiatives (L_NBM) was strongly correlated with the same factors as expenditure on social activities (L_CP); however, the change in deposits (D_G) was also significant. Therefore, it can be concluded that hypothesis 2 was confirmed.

3.5. Model specification

Preliminary data analysis made it possible to propose an empirical model that defines the impact of social activity on a bank's financial efficiency using the dynamic nature of the interactions between the elements. The model is expressed as follows:

$$ROA_{it} = \gamma ROA_{i.t-1} + \beta_1 L\_CP_{i.t0} + \beta_2 L\_CP_{i.t-1} + \beta_3 L\_CP_{i.t-2} + \beta_4 L\_G_{i.t0} + \beta_5 D\_G_{i.t0} + \beta_6 TCR_{i.t0} + \\ + \beta_7 NPL_{i.t0} + \beta_8 L\_NBM_{i.t0} + \beta_9 LOC_{i.t0} + \beta_{10} GDP_{i.t0} + \varepsilon_{i.t}$$

$$ROE_{it} = \gamma ROE_{i.t-1} + \beta_1 L\_CP_{i.t0} + \beta_2 L\_CP_{i.t-1} + \beta_3 L\_CP_{i.t-2} + \beta_4 L\_G_{i.t0} + \beta_5 D\_G_{i.t0} + \beta_6 TCR_{i.t0} + \\ + \beta_7 NPL_{i.t0} + \beta_8 L\_NBM_{i.t0} + \beta_9 LOC_{i.t0} + \beta_{10} GDP_{i.t0} + \varepsilon_{i.t}.$$

To assess durability of the estimation results, two subsequent models were estimated, considering proxy variables.

$$ROA_{it} = \gamma ROA_{i.t-1} + \beta_1 L\_NSI_{i.t0} + \beta_2 L\_NSI_{i.t-1} + \beta_3 L\_NSI_{i.t-2} + \beta_4 L\_G_{i.t0} + \beta_5 D\_G_{i.t0} + \beta_6 TCR_{i.t0} + \\ + \beta_7 NPL_{i.t0} + \beta_8 L\_NBM_{i.t0} + \beta_9 LOC_{i.t0} + \beta_{10} GDP_{i.t0} + \varepsilon_{i.t}$$

$$ROE_{it} = \gamma ROE_{i.t-1} + \beta_1 L\_NSI_{i.t0} + \beta_2 L\_NSI_{i.t-1} + \beta_3 L\_NSI_{i.t-2} + \beta_4 L\_G_{i.t0} + \beta_5 D\_G_{i.t0} + \beta_6 TCR_{i.t0} + \\ + \beta_7 NPL_{i.t0} + \beta_8 L\_NBM_{i.t0} + \beta_9 LOC_{i.t0} + \beta_{10} GDP_{i.t0} + \varepsilon_{i.t},$$

where i denotes the next bank among all analyzed banks (N), t specifies the year of the analysis in the entire research period (T), $t0$ is the current period, $t-1$ data delay by one year and $t-2$ data delay by two years. Parameters γ and β represent the regression coefficient in the estimated function between the explanatory variables and the bank's efficiency level, and $\varepsilon\_$ *(i.t)* denotes the remainder of the model.

4. EMPIRICAL MODELS ANALYSIS AND DISCUSSION

Tables 3 and 4 present the empirical model of the financial efficiency of cooperative banks measured with ROA, taking into account two variants of the variable defining the social activity of cooperative banks. The first model uses the level of expenditure on the bank's social activity (L_CP), taking into account the time shifts (Table 3), while in the second model, the number of social initiatives supported by the bank (L_NSI) was used as a measure of social activity (L_NSI), also considering the time shift between the bank's operation and its effects (Table 4).

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Table 3
Impact of corporate philanthropy (CP) on banks' ROA

Variable	Coefficient	Standard error	p-value
ROA(t – 1)	0.19352	0.07821	0.01340**
L_CP(t0)	0.01326	0.03810	0.72790
L_CP(t – 1)	0.01063	0.03229	0.74190
L_CP(t – 2)	0.05888	0.02659	0.02680**
L_G	0.00356	0.00231	0.12340
D_G	0.00004	0.00004	0.33500
TCR	0.02437	0.00544	<0.0001***
NPL	–0.01212	0.00482	0.01190**
L_NBM	–0.00173	0.01900	0.92730
LOC	–0.01106	0.03296	0.73730
GDP	0.00196	0.00218	0.36930

Significance level at: * 10%, ** 5%, *** 1%.
Statistical test for estimated model: Test for AR(1) errors: z = –2,46302 [0,0138], Test for AR(2) errors: z = –1,47205 [0,1410], Sargan over-identification test (18) = 177,924 [0,0000], Wald (joint) test: Chi-square (11) = 1378,14 [0,0000].

Source: Author's calculation.

Table 4
Impact of corporate philanthropy (NSI) on banks' ROA

Variable	Coefficient	Standard error	p-value
ROA(t – 1)	0.22682	0.07153	0.00150***
L_NSI(t0)	0.00989	0.02602	0.70370
L_NSI(t – 1)	–0.01390	0.02147	0.51730
L_NSI(t – 2)	0.05559	0.02939	0.05860*
L_G	0.00276	0.00238	0.24610
D_G	0.00005	0.00004	0.29670
TCR	0.02397	0.00561	<0.0001***
NPL	–0.01085	0.00473	0.02190**
L_NBM	0.00367	0.02225	0.86890
LOC	–0.00988	0.03280	0.76330
GDP	0.00335	0.00206	0.10380

Significance level at: * 10%, ** 5%, *** 1%.
Statistical test for estimated model: Test for AR(1) errors: z = –2,58948 [0,0096], Test for AR(2) errors: z = 1,36605 [0,1719], Sargan over-identification test (18) =160,217 [0,0000], Wald (joint) test: Chi-square (11) =1465,27 [0,0000].

Source: Author's calculation.

Four variables were found to be significant in the first model. The first variable determines profitability of total assets of a cooperative bank from the previous year (ROA(t – 1)), the increase of which in the previous year positively contributed to increased efficiency in the current year, because banking activity is closely related to the bank's effectiveness and, in many cases, constitutes the basis for creating the equity of a cooperative bank necessary to develop lending and ensure its financial security. The second variable determines expenditure on social activities

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of a cooperative bank from two years before the year of analysis (L_CP($t-2$)). It means that the expenses incurred for social purposes by the examined cooperative banks two years earlier contributed to increased effectiveness in the current year. Therefore, it can be assumed that increasing spending on social activities by cooperative banks will contribute to a further increase in their financial efficiency, as measured by ROA, and the delay in the effect of the expenses was as long as two years. Thus, cooperative banks striving to increase profitability of total assets should pay attention to the level, scope, and continuity of their social activities. However, attention should be paid to the law of diminishing the efficiency of subsequent expenditures, which is associated with the need to exercise moderation to support subsequent social activities.

Moreover, the model includes a variable specifying the TCR level, the increase of which, in accordance with the estimated model, resulted in an increase in the bank's financial efficiency measured by ROA. It was due to the fact that increases TCR, on the one hand, indicated an increase in the bank's financial security, and at the same time was associated with an increase in its efficiency, which resulted from the fact that the bank had a stable capital base, thanks to which it could increase lending, which was the main source of its revenues. Therefore, in the case of cooperative banks, it is important to find a point at which the bank achieves a satisfactory level of effectiveness with an acceptable level of financial security. Another important variable included in the model determining NPL also had a direct and negative impact on ROA. It means that cooperative banks should limit their operating costs and credit risk, that is, factors directly affecting their net financial results and profitability of total assets. The record of the variable as a destimulant may indicate that cooperative banks should focus their activities more on granting high-quality loans and that they need to have a high-quality loan portfolio. It is also important that when using a proxy variable L_NSI to determine the impact of social activity on ROA, the obtained results indicate the significance of the number of social initiatives supported by the bank for two years earlier (L_CP($t-2$)) and the same variables as in the case of the model for ROA estimated using L_CP, that is, TCR and NPL. Thus, it confirms the importance of social activity in shaping profitability of cooperative banks, and the impact shifted in time, as social activities undertaken in a given year resulted in an increase in ROA after two years.

Table 5 presents the model of efficiency of a cooperative bank expressed in return on equity (ROE), using the adopted set of control variables and expenditure on social activity (L_CP). The results showed the statistical significance of the five variables. The first variable was the ROE level from the year before the analysis period. The next variable concerned social activity expenditure two years before the analysis period, similar to the ROA model. The model also included a variable determining increased deposits, the impact of which on the level of ROE was positive, meaning that when a bank increased its deposit base, its asset profitability improved. It was due to the fact that cooperative banks, based on customer deposits obtained at low cost resulting from the low level of interest rates on the market, could use them to grant loans or invest in securities that generated income for the bank at a level exceeding their capital acquisition costs. It should be emphasized that increasing the deposit base without the bank's lending and investment activities enabling the use of collected deposits would lead to a decrease in the bank's efficiency by generating excessive costs associated with their acquisition. NPL was another significant variable in shaping the ROE. It indicates that lending activities by cooperative banks should be associated with limiting credit risk, which negatively affects ROE and ROA. The last variable in the model concerned the number of members of the cooperative bank (L_NBM), the increase in which contributed to the increase in ROE. It was due to the fact that a greater number of members of the cooperative bank contributed to the growth of the bank's capital, which could constitute the basis for the development of the bank's lending activity and improvement of its financial security by strengthening the bank's capital base. Thus, it allows banks to increase lending while maintaining their current level of capital requirements.

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Table 5
Impact of corporate philanthropy (CP) on banks' ROE

Variable	Coefficient	Standard error	p-value
ROE(t – 1)	0.18908	0.07841	0.01590**
L_CP(t0)	0.47376	0.34148	0.16530
L_CP(t – 1)	–0.34633	0.30085	0.24970
L_CP(t – 2)	0.67572	0.31691	0.03300**
L_G	0.01697	0.02031	0.40340
D_G	0.00088	0.00039	0.02450**
TCR	0.04310	0.03227	0.18160
NPL	–0.15313	0.04159	0.00020***
L_NBM	0.38278	0.16192	0.01810**
LOC	–0.43336	0.28830	0.13280
GDP	0.02867	0.02250	0.20250

Significance level at: * 10%, ** 5%, *** 1%.
Statistical test for estimated model: Test for AR(1) errors: z = 2,62228 [0,0087], Test for AR(2) errors: z =- 0,729455 [0,4657], Sargan over-identification test (18) = 169,295 [0,0000], Wald (joint) test: Chi-square (11) = 1475,22 [0,0000].

Source: Author's calculation.

In the course of the study, the ROE model was also estimated with using the number of supported social initiatives (L_NSI) as a measure of the social activity of cooperative banks. The results of the model estimations are presented in Table 6. In the model, as in the case of the ROE model estimated with the use of L_CP, the following were significant: ROE level last year, increase in deposits (D_G), asset quality (NPL), and the number of bank members (L_NBM). It confirms the significance of the impact of the variables on the profitability of cooperative banks' equity. Additionally, the variable determining GDP, which had a positive impact on ROE, turned out to

Table 6
Impact of corporate philanthropy (NSI) on banks' ROE

Variable	Coefficient	Standard error	p-value
ROE(t – 1)	0.23052	0.07140	0.00120***
L_NSI(t0)	0.17515	0.27575	0.52530
L_NSI(t – 1)	–0.41896	0.20356	0.03960**
L_NSI(t – 2)	0.91362	0.32953	0.00560***
L_G	0.01005	0.02027	0.62010
D_G	0.00085	0.00040	0.03120**
TCR	0.04618	0.03345	0.16740
NPL	–0.13760	0.03958	0.00050***
L_NBM	0.39007	0.18729	0.03730**
LOC	–0.40223	0.29300	0.16980
GDP	0.04627	0.02141	0.03060**

Significance level at: * 10%, ** 5%, *** 1%.
Statistical test for estimated model: Test for AR(1) errors: z = –2,85184 [0,0043], Test for AR(2) errors: z = –0,374757 [0,7078], Sargan over-identification test (18) = 169,295 [0,0000, Wald (joint) test: Chi-square (11) = 1540,58 [0,0000].

Source: Author's calculation.

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be significant in the model. It means that the improvement in the economic situation contributed to an increase in the profitability of banks' equity. In periods of economic boom, the demand for external financing sources increases, especially in the sector of small and medium-sized enterprises, which constitute an important group of customers for cooperative banks. It translates into an increase in the volume of loans granted by banks, which, while maintaining an appropriate level of quality, contributes to an increase in its financial results. What is really important is that the model also included two variables concerning the social activity of cooperative banks, and as in the case of the previous models, their impact shifted over time. In the model, the number of supported social initiatives from two years ago (L_NSI($t-2$)) and one year before the analysis period (L_NSI($t-1$)) was significant, while the influence of L_NSI($t-2$) was positive for ROE, and L_NSI($t-1$) was negative. However, regression coefficients for L_NSI($t-2$) were more than twice as high as those for L_NSI($t-1$), indicating that the time shift between the social activities undertaken by cooperative banks was delayed.

The results indicate that financial institution aiming to improve their financial efficiency, in addition to conducting basic banking activities related to granting loans and accepting deposits, should support the social activities that are part of the cooperative profile of the banking sector. It is worth emphasizing that excessive increase in spending on social activities may have a weaker or even negative impact on the financial efficiency of cooperative banks due to the law of decreasing marginal efficiency, and the fact that in order to successfully carry out the activity, cooperative banks must primarily conduct basic banking activities, which significantly shapes the level of financial efficiency of cooperative banks. Therefore, an important element of the banks' activities should be to find a balance between the implementation of the social mission and the pursuit of increasing financial efficiency at an acceptable level of risk. Nevertheless, the study showed that the activity of cooperative banks in financial support for social activity in their local environment generally favors an increase in their financial efficiency, which indicates the desirability of further tightening relations and cooperation with local communities. Therefore, it was found that hypotheses 1 and 2 were confirmed. The results of the study are consistent with those of other studies on the factors determining financial efficiency of banks. In the developed proprietary models, the level of TCR appears to be very important, as its growth positively influences ROA in general. Similar results have been obtained in previous studies by Bouzgarraiu et al. (2018), Rumel and Waschniczek (2016) and Bertay et al. (2013). It confirms that the appropriate level and structure of a cooperative bank's capital are necessary elements for shaping the growth of its ROA. The study also indicated significance of NPL, the increase of which contributes to the decline in'nksthe ROA and ROE, which is also indicated by the research conducted by Petria et al. (2015). In the developed models, D_G was also significant as it had a positive effect on ROE growth. The direction of the relationship between the factors was also indicated by Rachdi (2013). To increase their profitability, banks must effectively use the accumulated funds. The number of bank members (L_NBM) also appears among the important explanatory variables in the models developed for the ROE. The increase in the variable contributed to the improvement of the ROE by increasing banks' equity, which formed the basis for conducting lending activities (Huang, 2020). Moreover, in the model for ROE, GDP turned out to be significant, as its increase resulted in an increase in ROE in the cooperative banks surveyed. As emphasized by Kuc and Teplý (2022), the relationship is caused by the relationship between enterprises' need for financial resources in the form of loans and the economic cycle. Thus, an improvement in the economic situation leads to an increase in the profitability of the capital of cooperative banks.

In the ROE and ROA models for the analyzed banks, the variables determining the social activity of cooperative banks (L_CP and L_NSI) also turned out to be of key importance. The results obtained are partially consistent with those of other studies on the impact of social activity on banks' financial efficiency (Orlitzky and Shen, 2013). Similar results were obtained by Djalilov et al. (2015), who pointed to the existence of a positive relationship between banks'

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social activity and ROE and ROA. Nizam et al. (2019), Belasri et al. (2020), and Tran et al. (2020) also confirmed that the social activity of banks has a positive effect on the return on equity (ROE), return on total assets (ROA), bank's interest income (NII), and interest margin (NIM). Moreover, in the literature, one can find studies indicating the existence of a negative relationship between social activity and the effectiveness of banks (Oyewumi, Ogunmeru and Oboh, 2018) and studies indicating no such relationship (Masulis and Reza, 2015).

Our study introduced a new element to the discussion in the field of assessing the impact of social activity on the effectiveness of banks, which considers time shifts between expenditure on social activity and the resulting effects. This may partially explain the variation in the results obtained thus far by other researchers who have not conducted such analyses. Our study shows that it may be an important aspect in the context of managing financial institutions and assessing their financial effectiveness.

5. ROBUSTNESS ANALYSIS

To increase credibility of our conclusions, a robustness test was performed using stepwise regression for the generalized method of moments (GMM) (Dahir et al., 2019). It was to confirm the significance of the results obtained and eliminate the potential problem of endogeneity resulting from the degree of corporate philanthropy being conditioned by the financial results of a bank. Table 7 shows the estimation results. They confirmed the significance of all the variables that were important in the developed models determining the relationship between CP and CF.

Table 7
Robustness tests GMM step-wise regression

Variable	Model 1		Model 2		Model 3		Model 4	
	Coefficient	Significance level	Coefficient	Significance level	Coefficient	Significance level	Coefficient	Significance level
ROA(t – 1)	0,42503	<0,0001***	0,45354	<0,0001***				
ROE(t – 1)					0,2419	0,0001***	0,2677	<0,0001***
L_CP(t0)								
L_CP(t – 1)								
L_CP(t – 2)	0,05437	0,0006***			0,6337	0,0029***		
L_NSI(t0)								
L_NSI(t – 1)								
L_NSI(t – 2)			0,04757	0,0006***			0,4690	0,0154**
L_G								
D_G	–0,00625	0,0002***	–0,00643	0,0002***			–0,0263	0,0929*
TCR	0,01714	<0,0001***	0,01683	<0,0001***				
NPL	–0,01254	0,0016***	–0,01109	0,0027***	–0,1681	<0,0001***	–0,1530	<0,0001***
L_NBM					0,4256	0,0005***	0,4914	0,0001***
LOC								
GDP	0,01856	<0,0001***	0,01990	<0,0001***	0,0977	<0,0001***	0,1434	<0,0001***

Significance level at: * 10%, ** 5%, *** 1%.
Statistical test for estimated models: Model 1: Test for AR(1) errors: $z = 5{,}15632$ [0,0000], Test for AR(2) errors: $z = 0{,}743043$ [0,4575], Sargan over-identification test (19) = 165,366 [0,0000], Wald (joint) test: Chi-square (6) = 2395,65 [0,0000]; Model 2: Test for AR(1) errors: $z = –5{,}13587$ [0,0000], Test for AR(2) errors: $z = –0{,}788327$ [0,4305], Sargan over-identification test (19) = 163,782 [0,0000], Wald (joint) test: Chi-square (6) = 2732,23 [0,0000]; Model 3: Test for AR(1) errors: $z = –4{,}10823$ [0,0000], Test for AR(2) errors: $z = –0{,}69858$ [0,4848], Sargan over-identification test (19) = 166,749 [0,0000], Wald (joint) test: Chi-square (6) = 1495,31 [0,0000]; Model 4: Test for AR(1) errors: $z = –3{,}92027$ [0,0001], Test for AR(2) errors: $z = –0{,}468182$ [0,6397], Sargan over-identification test (19) = 166,39 [0,0000], Wald (joint) test: Chi-square (6) = 1538,24 [0,0000].

Source: Author's calculation.

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in particular, the delayed impact of corporate philanthropy on the financial efficiency of cooperative banks yet also emphasized the importance of D_G, which was also significant for ROE in the performed robustness tests. Moreover, the GMM models indicate the significance of the dynamics of GDP, which occur in all the estimated models determining the financial efficiency of cooperative banks. The results indicate the accuracy of the applied methodological approach and durability of the obtained results.

6. CONCLUSIONS

Financial institutions, including banks try to fulfill their social mission by acting for the benefit of all their shareholders, undertaking tasks that other economic entities are not always interested in. The strength of the sector is often its individual approach to a client and its ability to build relationships. Therefore, in the era of sustainable development policies in global, national, or regional economies, our research can be a significant aspect that has not been analyzed before and can influence the shaping of, among others, a conscious sustainable development policy regarding the social activities of financial institutions. The study demonstrated the importance of social factors in shaping financial efficiency, which can be relevant for those managing financial institutions. Moreover, it indicates that in the case of expenses related to social activities, a common occurrence is a time shift associated with their occurrence of up to two years. Awareness of the time shift can lead to a more intentional construction of the bank's efficiency policy by the management, taking into account social activity, or the introduction of regulatory changes concerning cooperative banks in this regard. However, it should be emphasized that the probable excessive increase in spending on social activities can have a limiting impact on the financial efficiency of banks due to the law of decreasing effectiveness of subsequent financial expenditures. It is also important to note that corporate philanthropy is not the core banking activity and should not be treated as the basis for the functioning of financial institutions. Therefore, finding a good balance between the scope of social activities and the pursuit of increased financial efficiency with an acceptable level of risk should be an essential element of the activities of all financial institutions nowadays.

7. LIMITATIONS AND FUTURE RESEARCH DIRECTION

Although our study has shown its theoretical and practical significance, it has some limitations. The first is the limited scope of the study related to the geographic area covered.

As a result, the outcome of our study cannot be directly transferred to other regions; therefore, future research could focus on the impact of corporate philanthropy on the financial efficiency of financial institution, considering the conditions in different countries or regions. Another limitation concerns the narrowing of the analyses to the relationship between social activity and the effectiveness of cooperative banks, as shown in the literature. The activity also translates into other areas of the bank's activity and determines the impact of corporate philanthropy on such elements as the quality of management, the quality of customer relations, or liquidity or solvency of a bank, which could fill the gap in this respect. In addition, among the variables determining corporate philanthropy and financial performance, two measures were used to determine the level of spending on social activity (CP) and the number of supported initiatives (NSI), and two measures for cooperative bank financial performance (ROA, ROE) which, despite their frequent use by other researchers, have certain limitations. Therefore, future research could include an attempt to clarify the measures that allow for the measurement of corporate philanthropy in the financial sector or apply other methods of assessing financial performance. Our study also does not take into consideration the impact of financial crises, the COVID-19 pandemic or the

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war in Ukraine on social activities undertaken by financial institutions, which could also be an interesting and valuable aspect for further consideration in the analyzed area.

8. AUTHOR CONTRIBUTIONS

Conceptualization: Marta Idasz-Balina; methodology: Marta Idasz-Balina, Rafał Balina; formal analysis: Marta Idasz-Balina, Rafał Balina; investigation: Marta Idasz-Balina, Rafał Balina; writing – original draft preparation: Marta Idasz-Balina, Rafał Balina; writing – review and editing: Marta Idasz-Balina, Rafał Balina, Adam Zając, Krzyszof Smoleń; supervision: Marta Idasz-Balina, project administration: Rafał Balina

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Comparison of the financial market conditions in Poland and selected countries during the pandemic and the Russian-Ukrainian war

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ABSTRACT

The initial period of the COVID-19 pandemic and the outbreak of the Russian-Ukrainian war had a significant impact on financial markets. The aim of the work is to analyse the behaviour of selected segments of the Polish financial market in the initial period of the pandemic and war. The research is based on two measures of risk, relative and absolute. The research covers the period from 24 February 2022 to 22 August 2022. The research focuses on segments of the Warsaw Stock Exchange through the analysis of the following indices: WIG, WIG20, mWIG40, sWIG80, NCIndex, industry indices, and the TBSP treasury bond market index. The Warsaw Stock Exchange indices are also compared using variation coefficients and standard deviations with indices of selected world stock exchanges. After the start of the war, declines in the Warsaw Stock Exchange occurred with a delay of over a month. They can be linked to specific EU actions in the area of sanctions against Russia. Both risk measures indicate a similar response to the main indices, which can be interpreted as the high level of development of the Warsaw Stock Exchange, as compared to other stock exchanges.

JEL classification: C43, C49, D53, E44, G50

Keywords: capital market, financial market, COVID-19 pandemic, Russian-Ukrainian war, measures of risk, market indices.

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1. INTRODUCTION

The initial period of the COVID-19 pandemic and the outbreak of the Russian-Ukrainian war were unprecedented events that had a significant impact on financial markets. Due to earlier research on the beginning of the pandemic presented in the works (Żebrowska-Suchodolska et al., 2021, 2022) covering about half a year (from 19 November 2019 to 15 May 2020), it was also decided to cover the half-year period of the beginning of the Russian-Ukrainian war (from 24 February 2022 to 22 August 2022). It seems that the assumed period largely reflects the "shock" of financial markets to events such as the pandemic and war. In the latter case, they were dealing with a series of actions by the European Union in the form of sanctions imposed on all projects in which Russia was involved. Undoubtedly, they have an impact on the behaviour of entities operating on capital markets. Therefore, the aim of the work is to a large extent a qualitative, but supported by elements of quantitative analysis, comparison of reactions of selected segments of financial markets to the initial period of the pandemic and the Russian-Ukrainian war. The work focuses on the Polish stock market as the basic element of the capital market. Therefore, in the first place, the behaviour of the segments of the Warsaw Stock Exchange was compared by analysing the following indices: WIG, WIG20, mWIG40, sWIG80, NCIndex, and the TBSP treasury bond market index. It should be noted that the indices describe qualitatively different segments of the stock market, primarily determined by the size of companies (except for the TBSP index). Thus, the knowledge of a large or small correlation between the percentage changes of indices shows investors' interest in more or less risky markets. In the analysed periods, markets described by stock indices were relatively strongly positively correlated, but this is not always the case. This information is important in itself, as it shows that investors attributed a similar risk to individual market segments, understood qualitatively as the volatility of prices of fundamental companies (WIG20), relatively smaller ones (mWIG20, sWIG80) or even speculative ones (NCIndex). Due to editorial limitations, a detailed analysis of industry indices was abandoned, leaving only a comparison of variation and standard deviations (risks) in both analysed periods. The Polish capital market operates within the framework of the European Union, so it was decided that it is worth looking at how it behaves in comparison to other stock markets. For the purpose, quotations of the WIG20 index were compared with the German DAX index and the French CAC index. It was decided that the situation of the United Kingdom after its recent departure from the EU is worth analysing as it is still part of the European market. Moreover, due to the special involvement of the USA in helping Ukraine and leading the anti-Russian campaign, the DJIA index was included in the analysis. In the final part of the work, quotations of EUR, USD, and GBD to PLN were dealt with, as an indicator of the situation in the currency market from the point of view of the Polish zloty.

2. LITERATURE REVIEW

The global crisis caused by the COVID-19 pandemic, which started in China, has spread to all countries of the world. It affected the economies, and earlier, the world stock exchanges. The reaction of world stock exchanges took place earlier when information about the first cases of the disease in China appeared. It was then that the stocks began to fall. However, the biggest drop occurred on the days when individual countries began to announce lockdowns. Because the largest reaction of the stock markets took place in the initial phase of the pandemic, many studies focus on the initial period. Then, the pandemic hurt global stock exchanges (He et al., 2020), (Scherf et al., 2022). The impact in the initial period of the pandemic was the strongest in Asian emerging markets (Topcu, Gulal, 2020). The reaction of stock markets also changed over time depending on the stage of the pandemic and the number of cases (Ashraf, 2020). Research on the impact of COVID-19 was also conducted for individual industries. Natural gas, food, healthcare, and software stocks posted strong positive

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returns, while oil, real estate, entertainment, and hospitality stocks fell sharply (Mazur et al., 2021). Financial markets were also studied in terms of their volatility (Mirza et al., 2020), (Yousef, 2020) or in terms of the relationship between COVID-19 and basic economic indicators (Uddin et al., 2021).

The return of stock markets to their pre-pandemic state had barely occurred, and here was another reason for the crisis of the world's economies and stock markets. It was the year-long war in Ukraine. The reaction was instantaneous (Izzeldin et al., 2023). On the day of the attack, the S&P 500 index fell by more than 10% from October 2020. The reactions of most global stock indices were also negative with the largest reaction on the day of the invasion (Boungou, Yatié, 2022). The invasion of Ukraine in 2022, generated negative cumulative abnormal returns for global stock indices (Boubaker et al., 2022). Stock exchanges of Hungary and Russia were the first to react, then Poland and Slovakia. The reaction took place the day before the invasion. Stock markets of Australia, France, Germany, India, Italy, Japan, Romania, South Africa, Spain and Turkey reacted the day after the invasion (Yousaf, 2022). For commodity markets, the results were in line with those for stock markets except for intensity (Izzeldin et al., 2023). However, depending on the country and its involvement in the war, the response was different (Ahmed et al., 2022). Also, depending on the industries, the response varied. In the EU countries, the sector strongly affected by the war was the manufacturing sector. The finance and services sectors also showed a negative impact, even more so than the industry (Sun et al., 2022).

The events of the outbreak of the pandemic and the war in Ukraine are called black swans in literature (Antipova, 2020), (Mielus, 2022). The European financial market has been negatively affected by Russia's aggression (Nimani, Spahija, 2023). Comparing the effects of the COVID-19 pandemic and the war in Ukraine, there was a lot of exchange rate volatility during both crises. However, greater exchange rate instability occurred during the war than during the pandemic. Also, in the case of financial markets, the war had a more significant impact on them than the pandemic (Mielus, 2022). A negative relationship was observed between war and the rate of return of the EU stock exchanges (Burdekin, Siklos, 2022), (Frey and Kucher, 2001), (Goel et al., 2017), (Hudson, Urquhart, 2015, 2022).

In light of the cited review of literature, the authors analyse the behaviour of selected segments of the Polish financial market in the initial period of the pandemic and war, filling the gap in the latter period research. The basis is two measures of risk, relative (coefficient of variation) and absolute (standard deviation). Based on superficial observations, it can be hypothesized that the comparison of both periods depends significantly on which risk measure we use. In the case of the Polish market, the relative measure indicates a slightly different dominance of the war period, as compared to the pandemic period than the absolute measure. Another hypothesis results from the comparison of Polish market indices with selected foreign markets and states that the Polish market does not stand out from other markets.

3. METHODOLOGY AND DATA

The basis for the analysis of the initial periods of the pandemic and war is the risk measured by the variation of the time series percentage return rates of indices and currencies. For variation comparison purposes, quotients of statistical coefficients of variation were calculated:

$$V = \frac{\left(\frac{\sigma}{\bar{r}}\right)_{war}}{\left(\frac{\sigma}{\bar{r}}\right)_{pandemic}}. \tag{1}$$

Where σ is the standard deviation of the sample and $\bar{r}$ the mean rate of change. The traditional interpretation of the variation coefficient was taken as the standard deviation per unit value of

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the average daily rate of return, so it is a relative measure of risk. The sign of the V coefficient depends on the sign of the average rates of return in both periods, so a positive value is interpreted as compatibility of the trends prevailing in the markets during the pandemic and war, regardless of whether it is dealing with an upward or downward trend. On the other hand, a negative value proves opposite tendencies of changes. In addition, a positive value and less than 1 indicates less variation during the war, as compared to the pandemic period, and greater than one, the opposite situation. A negative value of V and less than –1 proves greater variation during the war (absolute values of variation are then greater than 1), while greater than –1 corresponds to the opposite situation. In both cases, it is dealing with opposite trends prevailing in both periods.

In addition, standard deviations of quotations during the pandemic and war period were compared by calculating their ratio:

$$R = \frac{\sigma_{war}}{\sigma_{pandemic}}. \tag{2}$$

The R factor is related to the risk measured by standard deviation and treated as an absolute measure of risk (not related to average rates of return). A value greater than 1 proves a greater dispersion of quotations during the war, as compared to the pandemic period, i.e. a riskier period from the point of view of market participants. On the other hand, less than 1 corresponds to the opposite situation. Of course, it should be remembered that greater risk does not have a negative interpretation, because greater variation creates greater investment opportunities, but also greater potential losses.

The data was taken from the stooq.pl portal. The considered time series concerned daily quotations. In the case of quotations of foreign indices, the lengths of the series slightly differed due to different days off from quotations, not the same in different countries. However, the fact was omitted as the differences concerned two days in the case of the DAX index and three days in the case of the WIG20. In addition, it should be noted that the period under investigation begins three months before the first case of the disease in Poland. In contrast, the period of war is exactly on the day of the aggression, not earlier. It is due to the obvious fact that cases in other countries preceded the emergence of a pandemic in Poland, and therefore everyone expected infections in our country as well. In contrast, the war broke out unexpectedly, so there is no waiting period for it to start in the data used.

4. RESULTS

4.1. Trend on the Warsaw Stock Exchange

First of all, the situation on the Warsaw Stock Exchange was examined by comparing the WIG, WIG20, mWIG40, sWIG80 NCIndex, and the TBSP treasury bond market index. A comparison of quotations is included in Table 1.

It is obvious at first glance that during the pandemic, they were dealing with a clear sharp decline in share indices, followed by a fairly steady increase. Only in the case of the New Connect market, the index quite abruptly exceeded the peak of quotations from the period of the beginning of the pandemic. The situation is different in the case of the bond market. The TBSP index recorded a slight decrease at the beginning of the pandemic and quickly began to grow dynamically, reaching a local minimum in mid-March, thus reacting to the emergence of the pandemic in Poland. The decline coincides with the period of sharp declines in stock markets measured by other indices.

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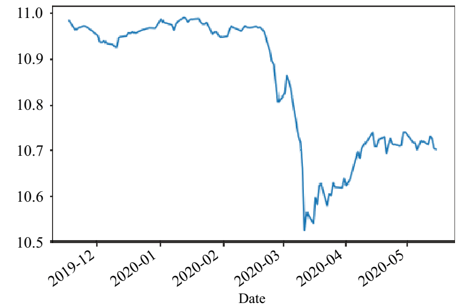
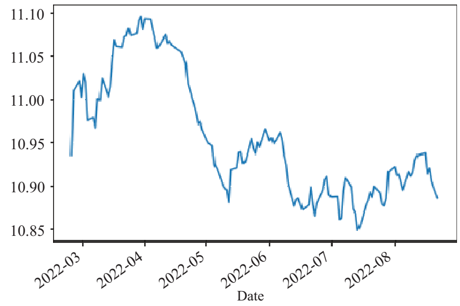
Table 1
Main indices of Warsaw Stock Exchange (logarithmic scale)

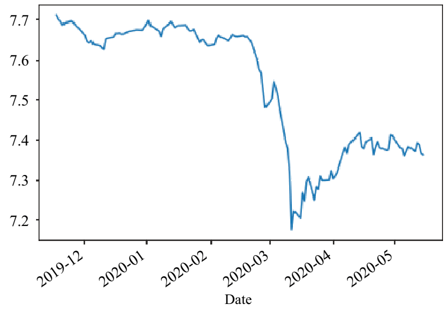
Pandemic | **War**

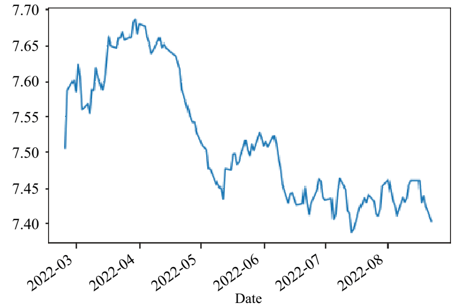
WIG



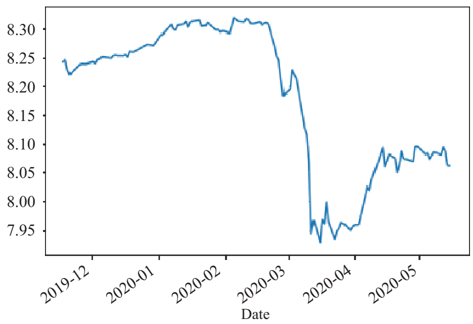


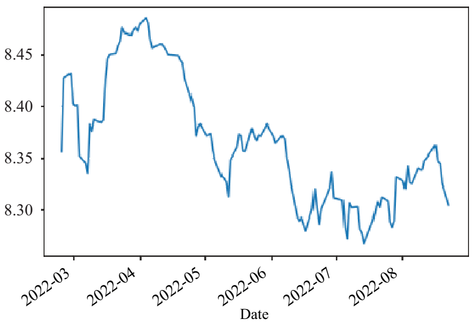
WIG20



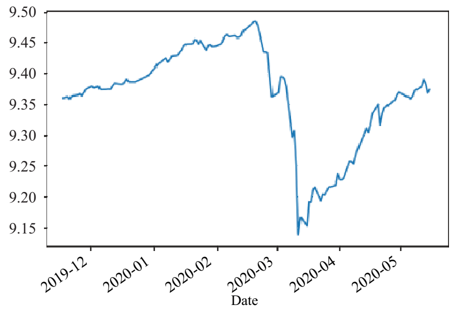


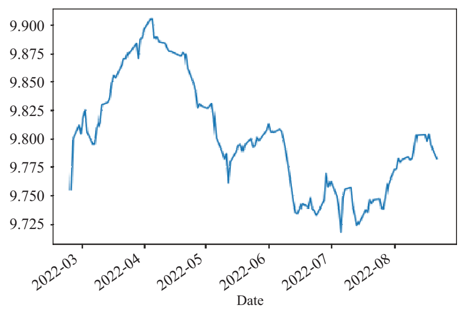
mWIG40



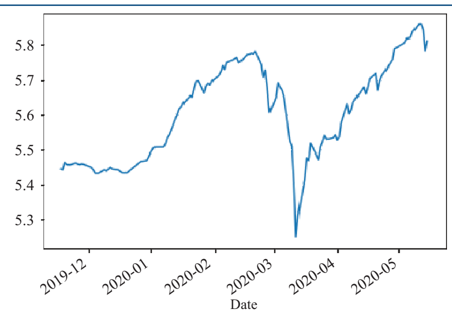


sWIG80





NCIndex



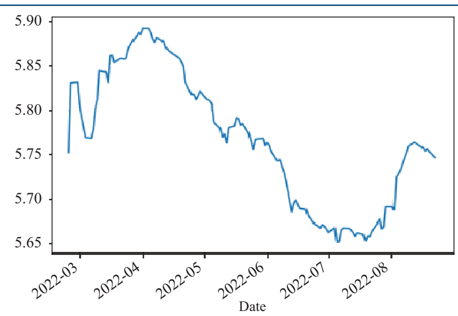


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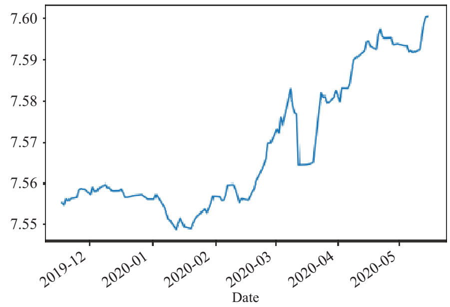
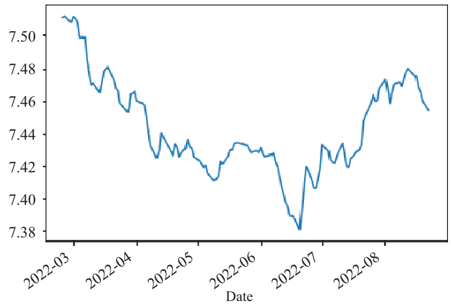
Table 1 – continued

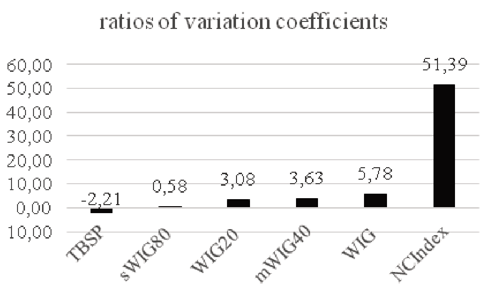
Pandemic	War
TBSP	
	

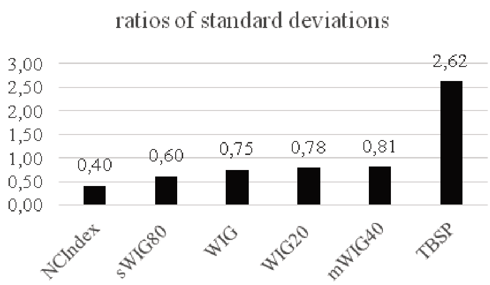
Source: own study.

The behaviour of stock indicators in the first days after the start of the war is completely different. Equity indices are clearly on the rise, reaching their maximum at the turn of March and April, followed by a rather rapid decline. In the next period, they are dealing with increases, but smaller than during the pandemic. In the case of the TBSP index, there is a gradual decline, which reaches its minimum only at the end of June. The results of the analysis of coefficients of variation and standard deviations are presented in Chart 1. In both cases, the values are arranged in ascending order.

Chart 1

Ratios of variation coefficients and standard deviations for the main WSE indices





Source: own calculation.

It should be noted that, except for the bond market, coefficients of variation ratios are positive, so the trends at the onset of the pandemic and the war are consistent. In the case of the TBSP index, the coefficient of variation during the pandemic period was more than two times higher in absolute value than during the war period, and the trends were opposite, which means that the variation during the war period was higher than during the pandemic period. Taking into account the other indices, only on the small companies market described by the sWIG80 index the pandemic period was characterized by greater variation than the period of the beginning of the war, the V coefficient is less than one. In other cases, the variation during the war was greater than during the pandemic. Against this background, the New Connect market stands out and the variation ratio is higher by an order of magnitude, as compared to other markets. The risk ratio graph shows a completely different situation. The New Connect market is characterized by the smallest value of the R-factor, so the dispersion of the rates of return around the average value during the pandemic was greater than during the war. A similar situation applies to other segments of the stock market, except the treasury bond market. In this case, the dispersion of quotations was much greater during the war.

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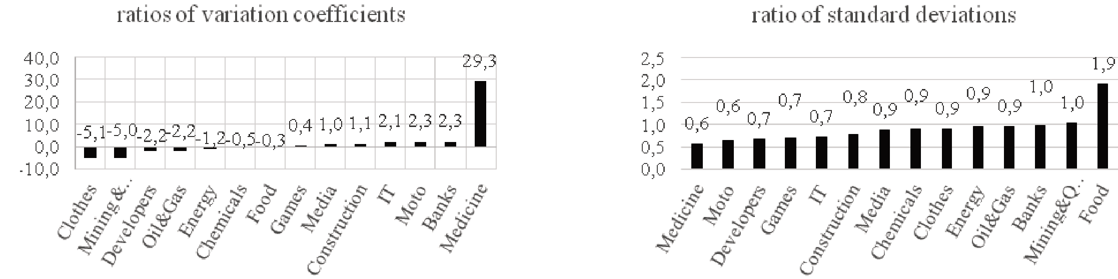
In the case of Polish industry indices, only variation coefficients and standard deviations in both periods are presented, as the presentation of the trends would take up too much space. As a consequence, results are shown in Chart 2 below.

Chart 2

Ratios of variation coefficients and standard deviations for WSE industry indices



Source: own calculation.

Ratios of variation coefficients for industry indices show predominance of variation during the war, as compared to the period of the pandemic. Clothing, mining, real estate, fuels and energy are characterized by negative variation ratios, which means that the trends characterized by the average daily rate of return were opposite in both periods. However, the values of the V coefficient are less than –1, so variability of the war period dominates variation during the pandemic. A similar situation applies to industries such as construction, IT, automotive industry, banks and medicines, but in this case, the trends in both periods are consistent. Against this background, the pharmaceutical industry stands out, for which the value of the V coefficient is equal to 29.3. This shows an almost 30-fold advantage of variation during the war, as compared to the pandemic period with the same trend. By contrast, in the case of the chemical, food, and gaming industries, a higher variation prevails during the pandemic, but in the first two cases the trends in both periods are opposite, and in the third they are consistent. At the adopted level of accuracy, the index of media companies was characterized by the same variation in both periods. For ratios of standard deviations R, the values vary in a small range from 0.6 to 1.9. The latter value applies to the food industry. In addition, banks and mining are characterized by the value of $R = 1$, which means the same risk in both periods. The remaining values of the ratios of standard deviations are less than 1, which proves that there is a slight risk advantage during the pandemic, as compared to the period of the war. As a consequence, it can be concluded that the risks in both borderlands were similar because no industry would stand out from the others.

4.2. Conditions of the Warsaw Stock Exchange and selected Foreign Stock Exchanges

Representative stock market indices are commonly recognized as indicators of the economic situation prevailing on the stock exchanges. Therefore, the research took into account the Polish WIG20 indices and the following indices of the leading European stock exchanges: DAX, CAC40 and FTSC100. Due to the involvement of the USA in helping Ukraine in the war with Russia, the DJIA index of the US stock exchange was additionally taken into account. The table 2 presents the logarithmic scale of market indicators in the initial period of the pandemic and the Russian-Ukrainian war.

In the authors' work (Żebrowska-Suchodolska et al., 2021) concerning the initial period of the pandemic, statistics of stock exchanges were analysed, among others, by comparing the dynamics of declines. However, in the case of the initial period of the war, such an analysis cannot be applied. there are no distinguished sessions when the indices reach their highest and lowest values. In the case of the presented indices, there is a similarity between the quotations during the pandemic and separately during the war. As a consequence, the following comparison is obtained, as shown in Chart 3.

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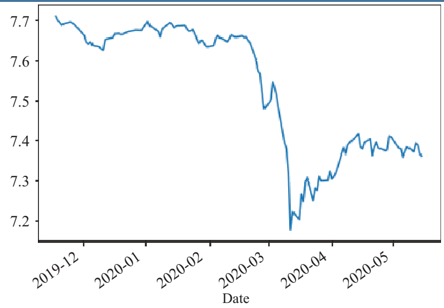
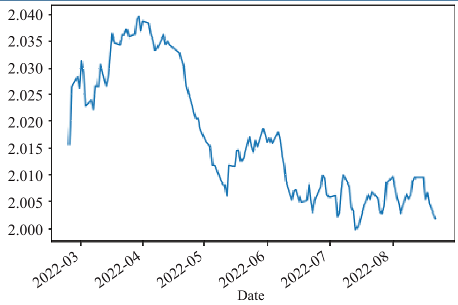
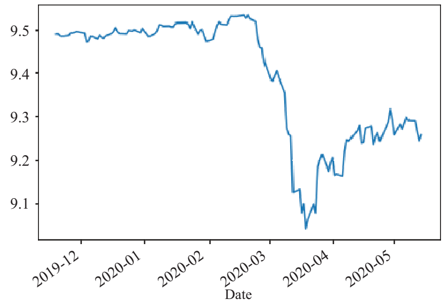
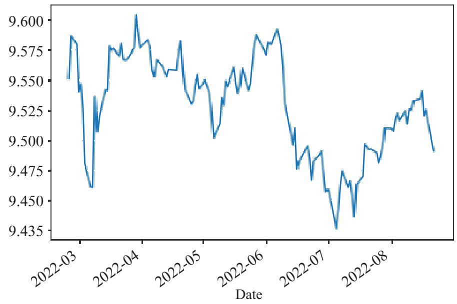
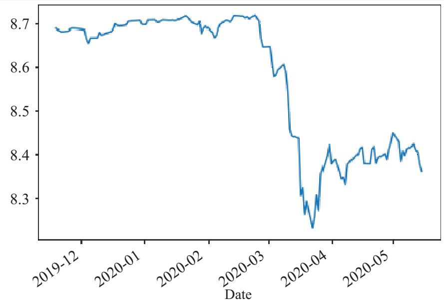
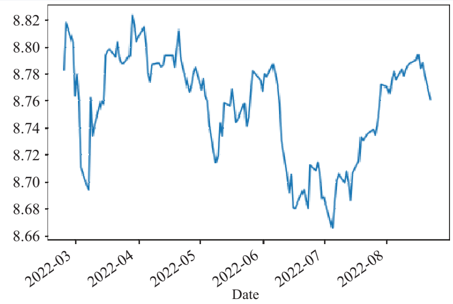
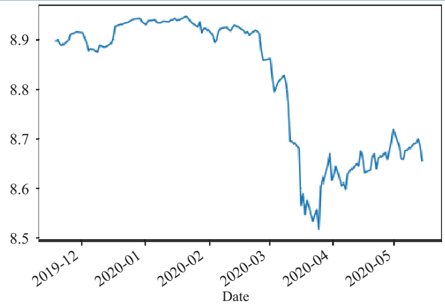
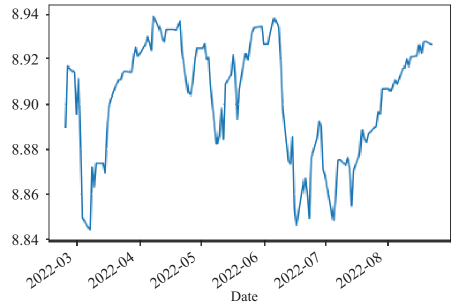
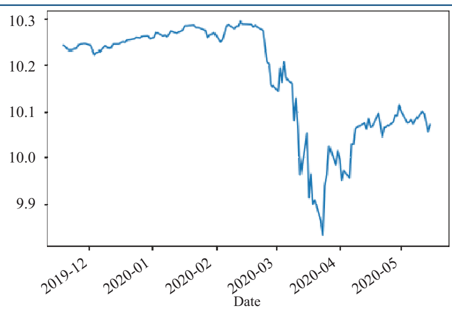
Table 2
Indices charts of selected European and American stock exchanges (logarithmic scale)

Pandemic	War
WIG20	
	
DAX	
	
CAC40	
	
FTSE100	
	
DJIA	
	

Source: own study based on stooq.pl.

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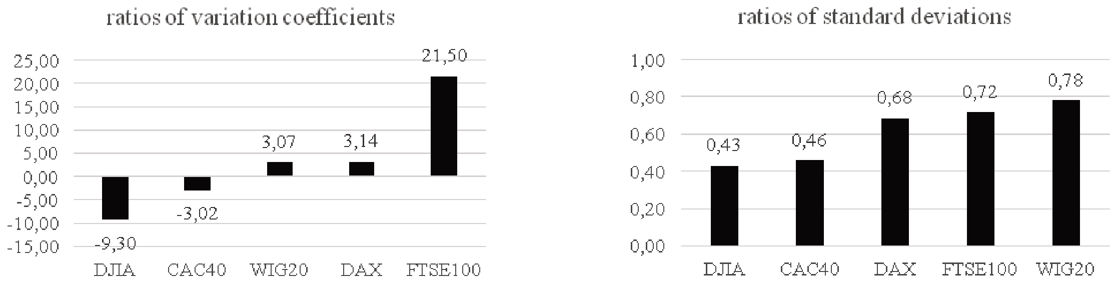
Chart 3
Ratios of variation coefficients and standard deviations of indices of selected stock exchanges



Source: own calculation.

It can be seen that in the case of the American and French stock exchanges, the trends during the pandemic and the war were opposite. Both values are less than –1, but the variation of the DJIA index is three times lower than the variation of the CAC40 index. In the remaining markets, the trends were consistent, but in all cases, the variation during the war was greater than during the pandemic. Against the background, the FTSE100 index stands out with its value approximately seven times higher than that of WIG20 and DAX. An interesting result is shown by the plot of the ratios of standard deviations. In all cases, the pandemic period was characterized by a higher absolute risk (measured only by the standard deviation, not related to the average rate of return), as compared to the war period. However, diversification is small, the smallest value concerns the American stock exchange (0.43), and the largest Polish one is (0.78).

4.3. Comparison of exchange rates of selected currencies to PLN

It was decided to compare exchange rates of EUR, USD and GBP, as a consequence, the chart presents the exchange rates of the currencies with the Polish zloty.

Table 3
EUR, USD and GBP to PLN exchange rates (logarithmic scale)

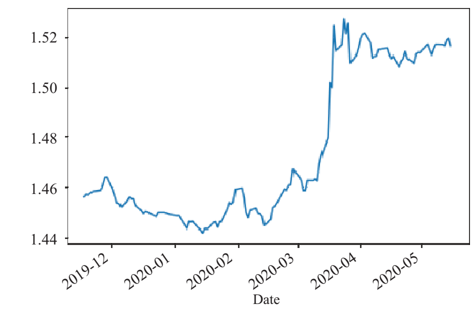
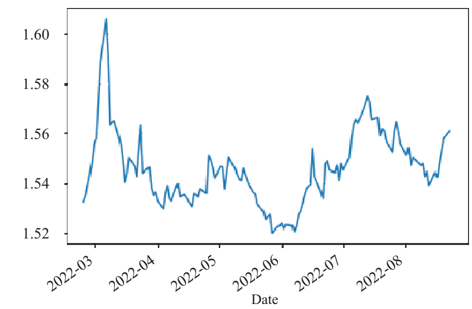
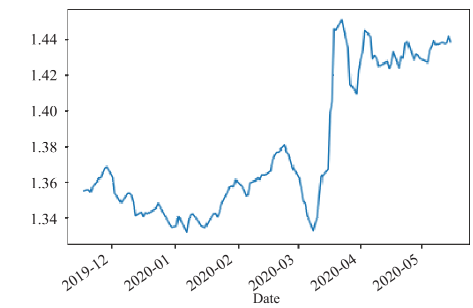
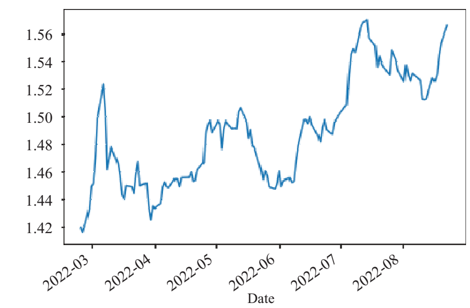
Pandemic	War
EUR/PLN	
	
USD/PLN	
	

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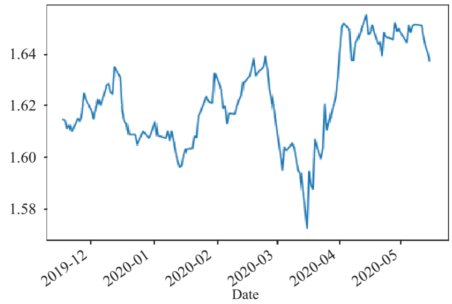
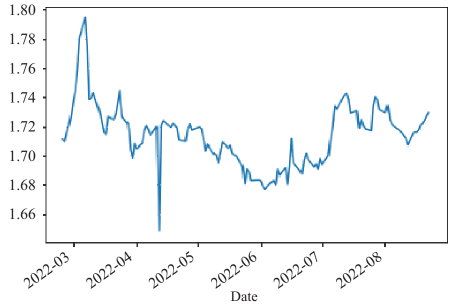
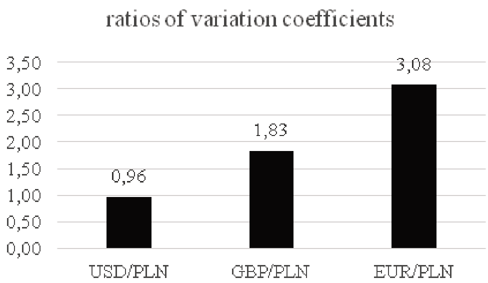
Table 3 – continued

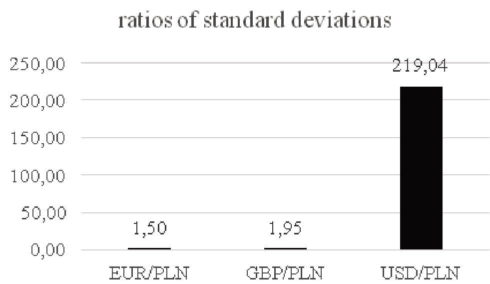
Pandemic	**War**
GBP/PLN	
	

Source: own study based on stooq.pl.

Coefficients of the variation of the above quotes are presented in Chart 4.

Chart 4
The ratios of variation coefficients and standard deviations of EUR, USD and GBP against PLN





Source: own calculation.

The charts above refer to the condition of the Polish zloty in the context of the three most important currencies. *V* coefficients are positive in all three cases, so the trends are the same. In the case of the USD/PLN exchange rate, a minimal advantage of variation was observed during the pandemic, as compared to the war period. For the British pound and the euro, variation prevails during the war, but is moderate. However, comparing the risks shows a completely different picture of the situation. Risks associated with the EUR/PLN and GBP/PLN exchange rates are small, not to say negligible, as compared to the USD/PLN exchange rate risk.

5. CONCLUSIONS

In summary, it can be stated that major drops in most of the main indices of the Polish stock exchange, apart from TBSP, occurred almost immediately after the outbreak of the pandemic in Poland, Table 1. On the other hand, the decline after the war started with a delay of over a month. It seems that it may be related to the emergence of sanctions against Russia, which did not immediately follow the aggression (similar conclusions are in (Ahmed et al., 2022)). Because they have only just begun to affect the capital market, mainly economic entities, the indices' decline can be associated with specific EU actions, and not with the period of consultations and arrangements. At the same time, there was uncertainty about the further actions of EU decision-makers and the course of the war. The situation began to clear up at the turn of July and August, then the index quotations began to slowly recover from the earlier drops. In the event of

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a pandemic, its global nature and negative impact on the financial market caused an immediate drop in index quotations.

On the other hand, during the pandemic the treasury bond market showed almost uniform growth, which seems understandable, there was a flight of capital to instruments that are much less risky than shares, because of bonds. In the case of the TBSP index, they are government bonds. During the war, however, the market for the instruments fell until the end of July, and then the trend reversed. Just like in the stock market. It seems that this time the factor determining quotations of the TBSP index was Poland's involvement in helping Ukraine and its proximity to the country, which creates additional risk. The analysis of coefficients of variation and standard deviations, Chart 1, seems to confirm the conclusion. During the pandemic, the "advantage" of war was observed, which explains the V value of –2.21. The minus sign indicates opposite trends in both periods. The ratio of standard deviations again indicates the advantage of risk during the war, and it is higher than the values for the other indices. The latter indicates an advantage of risk during the pandemic, all others have values lower than 1. It is worth noting that the ratio of coefficients of variation is negative only for the TBSP index, for the others it is positive. Thus, the stock markets showed the same economic situation in both analysed periods and, exceptionally, the small companies market described by the sWIG80 index was characterized by greater variation during the pandemic, as compared to the war period. Against this background, the New Connect market stands out, the variation coefficient exceeds the V coefficients for the other markets by an order of magnitude. On the other hand, the R coefficient for the market takes the lowest value. Therefore, the New Connect market showed a very high relative risk (coefficient of variation) during the war and a relatively low absolute risk (standard deviation) during the period.

In the case of the analysis of industry indices, variation of the economic situation can be noticed, the same number is characterized by positive and negative V coefficients. In this case, only three industries (chemical, food, and games) indicate the "advantage" of the pandemic. For other industries, values indicating a higher relative risk during the war or the same (media) can be observed. Against the background, drugs with a value of the V factor that is an order of magnitude higher than the other companies stand out. This can be interpreted as a much greater impact of the war on the industry, as compared to the pandemic period. Perhaps it is because during a pandemic there is a demand for a small variety of drugs, as compared to an armed conflict. The absolute risk analysis distinguishes the pandemic period, but not very strongly. For all industries, values are positive and range from 0.6 (drugs) to 1.0 (banks and mining), except for the food industry, for which $\boldsymbol{R} = \mathbf{1.9}$.

A comparison of both types of risk, relative and absolute, for the Polish stock exchange and selected foreign exchanges shows a similarity in the case of absolute risk, and the period of the pandemic "dominates". All R factors are less than 1, ranging from 0.43 (DJIA) to 0.78 (WIG20). On the other hand, dispersion of relative risks is relatively large, from $\boldsymbol{V} = \mathbf{9.3}$ (DJIA) to $\boldsymbol{V} = \mathbf{21.5}$ (FTSE100). However, it should be noted that all values indicate the advantage of the war period, as compared to the pandemic. It can also be said that the WSE does not stand out against the background of the analysed stock exchanges.

A comparison of EUR/PLN, USD/PLN, and GBP/PLN quotations shows that in the case of relative risk, the period of war predominates, only in the case of USD/PLN quotations the V factor is slightly lower than 1 and takes the value $\boldsymbol{V} = \mathbf{0.96}$. All values are positive, which indicates the same economic situation in both periods. Whereas absolute risk says the opposite, the R for USD/PLN is two orders of magnitude higher, as compared to EUR/PLN and GBP/PLN. This can also be explained by large changes in the currency markets, in particular large fluctuations in the EUR/PLN exchange rate. In addition, the role of the US during the war is completely different than during the pandemic, when the country faced difficulties analogous to those that affected other countries.

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In the context of the previously formulated research hypotheses, it can be concluded that the use of relative risk measures to compare initial periods of pandemic and war leads to different conclusions than those resulting from absolute margins. Consequently, one of the two risk measures should be used in more in-depth analyses. It seems that relative measures are more adequate because they take into account the situation in the markets, which is the source of significant interpretation conclusions. In conclusion, it should be mentioned that the Polish stock market does not stand out from the American, United Kingdom, German and French markets. Both risk measures used in the analysis of the pandemic and war periods indicate a similar reaction to the basic indices, which can be timidly interpreted as the high level of development of the Warsaw Stock Exchange.

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Cross Country Heterogeneity of Procyclicality of Bank Loans: Evidence from OECD Countries using the SURE Model

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ABSTRACT

Procyclicality of credit supply, which refers to the simultaneous movement of credit issued to the non-financial sector alongside economic activity indicators, can create a destabilizing feedback loop between the banking system and the real economy. The impact of credit supply on the financial and real sectors may vary across different economies, and the interconnectedness between countries can magnify the effect.

We conducted research examining procyclicality of loans provided by banks, analyzing data at the country level for 13 OECD countries for over 16 years (2005–2020). Our research findings indicate that the parameters measuring the procyclical effect are statistically insignificant when using the FE panel model. To showcase diversity of relationships under scrutiny across countries, we employed an OLS regression approach to estimate procyclicality for each country's loans. This approach assumes a lack of interconnectedness between economies.

We then introduced the Seemingly Unrelated Regression Equations (SURE) framework to examine how interconnectedness among countries affects the strength of loan procyclicality. Our analysis reveals the existence of procyclicality in many countries, and utilizing the SURE model further reinforces the phenomenon. Moreover, we found that bank-specific variables are more significant as loan supply determinants than macroeconomic variables.

JEL classification: E32, G21, G28

Keywords: procyclicality, credit supply, bank loans, capital management, risk management, seemingly unrelated regression.

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1. INTRODUCTION

Procyclicality refers to the feedback loops between the real economy and the financial system that amplify the business cycle. Alternative indicators can measure economic fluctuations (Boehl et al., 2016). According to Vanhoose (2010), the banking sector is inherently procyclical. This viewpoint was highlighted by Franco Modigliani, who believed that the tendency of credit markets toward instability, reflected in upswings and downturns, is natural (Modigliani et al., 1998).

During expansion periods, people can save more, leading to increased bank deposits. Additionally, consumers and entrepreneurs tend to increase spending, resulting in a higher overall demand for loans and credit from banks. Banks can meet the demand by supplying additional credit, which boosts their profits. Conversely, during a recession, individuals typically withdraw their funds to support themselves, leading to decreased deposits held by banks and a reduction in the overall supply of loans. The demand for loans also diminishes during a recession due to decreased consumption and reduced investment activities. Thus, it is expected that the aggregate level of loans supplied by banks will generally increase during expansion periods and decrease during recession periods.

Banks' procyclicality of credit supply is observed when it falls during economic downturns and rises during upturns (Borio et al., 2001). Literature on credit procyclicality can be broadly categorized into three groups, as Kouretas et al. (2020) outlined. The first group analyses various determinants of loans (Hempell and Sorensen, 2010; Jiménez et al., 2011; Cull and Martinez Pería, 2013; Cull et al., 2017; Kouretas et al., 2020). The second group focuses on the role of market structures in the real economy and their impact on credit (Petersen and Rajan, 1995; Kashyap and Stein, 2000; Altunbas et al., 2002; Claessens and Laeven, 2004; De Guevara and Maudos, 2011; Bikker and Leuvensteijn, 2014). The third group examines macro-financial linkages using panel vector autoregression models (Love and Zicchino, 2006; Marucci and Quagliariello, 2008; Bouvatier et al., 2012; Antonakakis et al., 2015; Apostolakis and Papadopoulos, 2019; Leroy and Lucotte, 2019).

Bank credit is determined by endogenous factors (bank-specific variables) and exogenous factors (macroeconomic variables). A robust financial system and a well-developed economy mutually support each other's growth. The ability of banks to expand long-term business loans depends on various factors, including capitalization, size, and the availability of long-term liabilities (Imran and Nishat, 2013). Vanhoose (2010) demonstrates that the aggregate level of loans in the economy typically increases during expansions and decreases during recessions. Goodhart and Segoviano (2004) explain that regulators are more stringent during recessions when reviewing banks due to higher default risk. This may lead to a contraction in loan supply to the economy. On the other hand, regulators are less strict during expansions, resulting in increased loan supply to the economy by banks. Bouheni and Hasnaoui (2017) show positive co-movements between bank lending and the business cycle for Eurozone banks, with differentiated impacts for larger and smaller banks.

The role of credit markets in the severe global recession of 2007–09 highlighted the need for a better understanding of the relationship between the financial sector and the real economy, which needed to be adequately incorporated into macroeconomic models (Gambetti and Musso, 2012). Comprehending the relative influence of supply and demand forces on credit and output is crucial, as this may require different responses from monetary and fiscal policy (Fourie et al., 2011). The global financial crisis, which triggered a severe worldwide recession, increased concerns about the procyclicality of bank risk-based capital requirements (Jokivuolle et al., 2015). Banks face more significant capital constraints than constant capital requirements as risk-based capital requirements rise during recessions. Consequently, banks may be compelled to significantly reduce lending, potentially exacerbating the recession (Kashyap and Stein, 2004).

While extensive research has focused on cyclicality of business credit, empirical evidence of credit supply cyclicality is available only for a limited number of countries (e.g., Gambacorta and

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Mistrulli, 2004; Marcucci and Quagliariello, 2008; Ivashina and Scharfstein, 2010; Becker and Ivashina, 2011; Jiménez et al., 2011). There is no evidence of cross-country linkages in analyzing credit growth procyclicality. Increasing interconnectedness of financial institutions and markets, along with more highly correlated financial risks, has intensified cross-border spillovers through various channels (Claessens et al., 2011; Olszak and Pipień, 2016; Fernandez-Gamez et al., 2020). Arčabić and Škrinjarić (2021) analyze spillovers and synchronization of business cycles in the European Union and find pronounced spillovers, highlighting the importance of studying cross-country linkages for the EU countries. Kouretas et al. (2020) investigate the impact of market structure on the EU bank loans and find heterogeneities between advanced and transitioning EU banking sectors.

Against this backdrop, our paper aims to investigate the link between bank loans and their determinants using a balanced panel dataset comprising 13 OECD countries (Belgium, Canada, Denmark, France, Germany, Italy, the Netherlands, Norway, Poland, Spain, Sweden, Switzerland, and the United States) from 2005 to 2020. We employ an empirical estimation approach in two steps. First, we apply the Fixed Effect (FE) panel regression approach to identify the common procyclicality effect across all analyzed countries. The approach considers data from multiple countries and assumes constant parameters across countries. We employ the Seemingly Unrelated Regression Equations (SURE) system in the second step. The novel approach allows for variable parameters of interest across countries, enabling the testing of cross-country heterogeneity in the procyclicality effect. Moreover, it is an econometric framework suitable for analyzing the empirical significance of the standard panel regression outcomes, assuming a similar procyclicality effect for each explored country. Our primary research hypothesis posits that each country's procyclicality effect is specific, exhibiting substantial variability across different economies.

Previewing our main findings, we uncover a procyclical nature of loans in 6 of the 13 countries examined. Moreover, bank-specific variables hold greater significance as loan supply determinants than macroeconomic variables. Notably, applying the Seemingly Unrelated Regression Equations (SURE) model reinforces the statistical significance of business cycle and banking sector-specific variables, thus bolstering the procyclical effect of loans. Our research contributes to the existing literature on the procyclicality of bank loans in two significant ways. Firstly, we employ the SURE approach, allowing for distinct procyclical effects across each country. Secondly, our analysis sheds light on the role of interconnectedness among countries in estimating the strength of the procyclicality effect.

The remainder of the paper is organized as follows. Section 2 discusses hypotheses about the determinants of loans, considering both bank-specific and macroeconomic variables. Section 3 outlines the dataset and provides an overview of the empirical methodology. Subsequently, in Section 4, we present empirical results. Finally, Section 5 concludes the paper, highlighting implications for further research.

2. DETERMINANTS OF LOAN SUPPLY AND HYPOTHESES DEVELOPMENT

Extensive empirical research has consistently demonstrated that loan growth tends to be positive during economic upswings and damaging during recessions. Consequently, periods of rapid loan expansion are often accompanied by a decline in credit quality (Caporale et al., 2014). It can be attributed to banks' ability to increase lending by reducing interest rates or relaxing credit screening criteria for prospective borrowers. When screening criteria are relaxed, individuals previously deemed lacking sufficient creditworthiness may now be eligible for loans. However, such borrowers typically carry higher risk and are more likely to default in adverse scenarios, such as an economic downturn. Understanding the relationship between business cycles and the banking system remains a significant challenge for researchers and economists, particularly in light of the global financial crisis dated from 2007 to 2009.

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Our study considers a set of variables traditionally employed to explain credit supply, considering the income smoothing hypothesis (Greenawalt and Sinkey, 1988; Beatty et al., 2002; Liu and Ryan, 2006). Additionally, we modify the variables by incorporating measures of the business cycle, as observed in previous studies (Laeven and Majnoni, 2003; Bikker and Metzemakers, 2005; Olszak and Pipień, 2016). The chosen variables are presented in Table 1.

Table 1
Definitions of variables

Variable	Measure	Notation	Expected effect on Loan supply
Dependent variable:			
Loan Supply	Gross loans & advances to customers divided by total assets	*LOANS/TA*	
Determinants:			
Banking sector-specific:			
Profit	Profit before taxes divided by total assets	*PROFIT/TA*	+
Credit risk	Loans loss reserves divided by total assets	*LLP/TA*	–
Deposits	Capital to assets ratio.	*DEP/TA*	+
Macroeconomic:			
Business cycle measure	Real GDP growth	*GDPG*	+
	Inflation (Consumer Prices Index)	*INF*	+/–
	Unemployment (% of total labor force)	*UNEMP*	–

Source: Author's own elaboration.

2.1. Macroeconomics determinants

Our analysis focuses on three macroeconomic determinants of loan supply, denoted as GDPG, INF, and UNEMP. GDPG represents real GDP growth and is a crucial indicator of loan procyclicality. It is widely preferred in investigating procyclicality at quarterly or annual frequencies (Banerjee, 2011). Empirical research consistently demonstrates a positive relationship between GDPG and credit supply (Gambacorta and Mistrulli, 2004; Bikker and Metzemakers, 2005; Jiménez et al., 2011; Banerjee, 2011; Kelly et al., 2013; Imran and Nishat, 2013). However, a negative correlation between GDP and loan supply may suggest countercyclical behavior by banks (Ibrahim, 2016; Albaity et al., 2020).

INF represents the inflation rate, measured by the consumer price index, which reflects the annual percentage change in the cost of a basket of goods and services for the average consumer. We include INF as an exogenous control variable. Previous studies have utilized inflation as a determinant of credit supply (Gambacorta and Mistrulli, 2004; Djiogap and Ngomsi, 2012; Klein, 2013). If INF is considered as an economic cycle variable, increasing during economic booms and decreasing during economic downturns, we would expect a positive relationship with loans, aligning with the procyclicality hypothesis that emphasizes loan supply increases during economic upswings and reduces during economic downturns (Klein, 2013).

UNEMP represents the unemployment rate, indicating the share of labor force without work but actively seeking employment. Unemployment and GDP series exhibit a negative correlation, meaning that unemployment tends to be higher during recessions and vice versa. Okun's law

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suggests that for every 3-percentage-point decrease in GDP from its long-run level (also known as potential GDP), the economy experiences a one-percentage-point increase in unemployment. Conversely, a 3-percentage-point increase in GDP from its long-run level is associated with a one-percentage-point decrease in unemployment. As an economic variable, unemployment is expected to affect loan supply negatively. Previous studies have considered unemployment a determinant of loans (Gambacorta and Mistrulli, 2004; Klein, 2013; Donaldson et al., 2015).

2.2. Bank determinants

As determinants of loan supply specific to the banking sector, we consider three bank-related variables: Profit, Loan Loss Provisions, and Deposits.

PROFIT represents the operating profit before provisions and taxes divided by the bank's total assets (PROFIT/TA). We examine the variable to assess whether profits lead to bank credit expansion. Richter and Zimmermann (2019) find that profits increase banks' net worth and lending capacity, thereby increasing the supply of loans. Profitability is positively correlated with credit supply (Barona and Xiong, 2017). Bank profitability could motivate banks to expand their loans, suggesting a positive correlation between profits and credit supply (Awdeh, 2017; Alihodžić and Ekşi, 2018).

Loan Loss Provisions divided by total assets (LLP/TA) are introduced as an independent variable to proxy for credit risk. Changes in total loans outstanding are related to changes in default risk (as well as credit risk). If banks use Loan Loss Provisions (i.e., their allocation to cover expected losses) to manage credit risk, the relationship between LLP and LOANS is expected to be positive. Conversely, if banks exhibit imprudent loan loss provisioning behavior, the supply of loans may have a negative impact on LLP. Empirical findings regarding the relationship vary. Some studies find a positive influence of real loan growth on LLP (Bikker and Metzemakers, 2005; Fonseca and González, 2008), implying that banks set aside provisions to cover risks accumulated during economic booms. Other studies document a negative coefficient on loans (Laeven and Majnoni, 2003), which rejects the prudent loan loss provisioning behavior hypothesis. Shala and Toçi (2021) explored banks in SEE (South-Eastern Economies) and their use of LLPs. They investigated procyclicality, capital management, and income smoothing. The authors recommend a dynamic provisioning system to enhance efficiency during business cycles. Transparency on provisioning could enhance proper provisioning and counter-procyclicality, which would help market discipline.

Total Deposits normalized by total assets (DEP/TA) are included to test the liquidity hypothesis. Traditionally, the amount of credit provided by banks was directly linked to the level of deposits they held. However, financial innovation in the past decade has severed the link between credit and deposits. The decoupling has been identified as a primary contributing factor to the 2007–09 financial crisis (Kelly et al., 2013). Given the structure of the banking system, lending typically generates deposits. When a bank grants a loan to a household or firm, the loan proceeds are initially credited to the borrower's bank account. It means that lending is initially offset by corresponding deposits, increasing the money stock (Bang-Andersen, 2014). Therefore, deposits exhibit a direct and positive relationship with credit supply in the economy.

Shala et al. (2020) provided empirical evidence that banks in nine South-Eastern European countries use loan loss provisions to smooth their incomes and that components of LLPs do matter in growth in bank lending. However, the study does not support the hypothesis that LLPs are used for capital management by banks in the region. Shala et al. (2022) analyzed NPL determinants using macroeconomic, structural, and bank-specific data from 17 CEE countries from 2006–2017. It includes legal environment indicators and assesses the GFC's effect on NPLs. The findings suggest strengthening microprudential supervision, considering credit growth and regulatory quality, and ensuring accurate indicator measurements for policy implications. Ozili (2017) examined whether Western European banks' discretionary provisioning is driven by credit risk

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or income smoothing. After the 2007–2009 financial crisis, bank regulators in Europe introduced strict rules on bank provisioning and risk-taking behavior. However, it is unclear whether Western European banks' provisioning behavior is driven by credit risk or income smoothing incentives. The study by Ozili (2017) finds that Western European banks' discretionary provisioning is driven by both income smoothing and credit risk considerations.

3. DATA AND RESEARCH METHODOLOGY

Our analysis utilizes aggregated yearly bank balance sheet and income statement data covering the period of 16 years (2005–2020). The dataset comprises information from 4870 banks across 13 OECD countries: Belgium, Canada, Denmark, France, Germany, Italy, the Netherlands, Norway, Poland, Spain, Sweden, Switzerland, and the United States. The comprehensive dataset is sourced from Moody's Analytics BankFocus at the bank level and then aggregated at the country level for our analysis.

We also incorporate macroeconomic variables from the World Bank Development Indicators database to supplement our analysis. The variables include the GDP growth rate, inflation rate, and unemployment rate, which provide essential contextual information for our study.

The basic model, based on Olszak and Pipień, 2016, is formulated within a panel regression framework as follows:

$$\frac{LOANS_{t,j}}{TA_{t,j}} = \alpha_0 + \alpha_1 \Delta GDPG_{t,j} + \alpha_2 INF_{t,j} + \alpha_3 UNEMP_{t,j} + \\ + \alpha_4 \frac{PROFIT_{t,j}}{TA_{t,j}} + \alpha_5 \frac{LLP_{t,j}}{TA_{t,j}} + \alpha_6 \frac{DEP_{t,j}}{TA_{t,j}} + \varepsilon_{t,j}, \tag{1}$$

where all variables are observed for the j-th country ($j = 1, \ldots, n$) at year $t = 1, \ldots, T$. The dependent variable is the loan supply (LOANS) of a bank divided by the bank's total assets (TA). Independent variables can be subdivided into two groups. In the first group we collect macroeconomic variables, like annual growth of the real Gross Domestic Product ($\Delta GDPG_{t,j}$), inflation rate ($INF_{t,j}$), and unemployment rate ($UNEMP_{t,j}$). The second group of variables consist of various bank-specific variables, like profits before taxes $\frac{PROFIT_{t,j}}{TA_{t,j}}$, loans loss provisions $\frac{LLP_{t,j}}{TA_{t,j}}$, and customer deposits of the bank $\frac{DEP_{t,j}}{TA_{t,j}}$. All banking sector specific variables are normalized by the bank total average assets (TA) to mitigate potential estimation problems with heteroscedasticity. Following Olszak and Pipień (2016), we move forward by relaxing assumption in (1) about cross-country homogeneity of parameters α_i. In order to perform this task, we refer to the system of Seemingly Unrelated Regression Equations (SURE) elaborated by Zellner (1962). To start let us rewrite equation (1) making all regression parameters variable across j:

$$\frac{LOANS_{t,j}}{TA_{t,j}} = \alpha_{0,j} + \alpha_{1,j} \Delta GDPG_{t,j} + \alpha_{2,j} INF_{t,j} + \alpha_{3,j} UNEMP_{t,j} + \alpha_{4,j} \frac{PROFIT_{t,j}}{TA_{t,j}} + \\ + \alpha_{5,j} \frac{LLP_{t,j}}{TA_{t,j}} + \alpha_{6,j} \frac{DEP_{t,j}}{TA_{t,j}} + \varepsilon_{t,j}, \tag{2}$$

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The standard assumption that, for each t, Gaussian error terms $\varepsilon_{t,j}$ and $\varepsilon_{t,i}$ in (2) are uncorrelated if $i \neq j$, makes the system of equations (2) independent. We denote the case by M_0. An application of such a system corresponds to the econometric strategy based on estimation of regression parameters separately for each country analysed. However, in general, error terms $\varepsilon_{t,j}$ and $\varepsilon_{t,i}$ can be correlated and system (2) can be treated as falling under the Seemingly Unrelated Regression Equations (SURE) model. We define the case as M_1, while $\varepsilon_t = (\varepsilon_{t,1}, \ldots, \varepsilon_{t,n})$ stands for the vector of error terms at time t with the covariance matrix $\sum$. In the case of model M_1 the matrix $\sum$ is symmetric and positively definite with $n(n+1)/2$ free elements (σ_{ij}^2), $i = 1, \ldots, n$ and $j = 1, \ldots, n$, such that $\sigma_{ij}^2 = \sigma_{ji}^2$. In the standard notation the variance of the error terms in the i-th country is denoted by $\sigma_{ii}^2 > 0$ and the covariance between error terms in j-th and i-th country is denoted by $\sigma_{ij}^2 \in R$. We apply the following notation to the dependent variable and the vector of explanatory variables:

$$y_{t,j} = \frac{LOANS_{t,j}}{TA_{t,j}},$$

$$x_{t,j} = \left(1, GDPG_{t,j}, INF_{t,j}, UNEMP_{t,j}, \frac{PROFIT_{t,j}}{TA_{t,j}}, \frac{LLP_{t,j}}{TA_{t,j}}, \frac{DEP_{t,j}}{TA_{t,j}}\right).$$

The system of equations (2) can be formulated in the following closed form:

$$y^{(j)} = x^{(j)} \alpha^{(j)} + \varepsilon^{(j)}, \qquad j = 1, \ldots, n,$$

where $y^{(j)} = (y_{1,j}, \ldots, y_{T,j})'$, $x^{(j)} = (x'_{1,j}, \ldots, x'_{T,j})'$, $\varepsilon^{(j)} = (\varepsilon_{1,j}, \ldots, \varepsilon_{T,j})'$ and $\alpha^{(j)} = (\alpha_{0,j}, \alpha_{1,j}, \ldots, \alpha_{6,j})'$. In the next step we stack the observations presenting the system of equations as a regression of the following form:

$$Y = X\alpha + \varepsilon, \tag{3}$$

where: $Y_{[nTx1]} = (y^{(1)'}, \ldots, y^{(n)'})'$, $\varepsilon_{[nTx1]} = (\varepsilon^{(1)'}, \ldots, \varepsilon^{(n)'})'$, $\alpha_{[n7x1]} = (\alpha^{(1)'}, \ldots, \alpha^{(n)'})'$, and:

$$X_{[nTxn7]} = \begin{pmatrix} x^{(1)} & 0_{[Tx7]} & \cdots & 0_{[Tx7]} \\ 0_{[Tx7]} & x^{(2)} & \cdots & \vdots \\ \vdots & \ddots & \ddots & 0_{[Tx7]} \\ 0_{[Tx7]} & \cdots & 0_{[Tx7]} & x^{(n)} \end{pmatrix}.$$

Simple calculations yield the following form of the covariance matrix for the error term ε in (3):

$$V(\varepsilon) = \sum \otimes I_n,$$

where $\otimes$ denotes the Kronecker product. The form of the covariance matrix of ε makes the system (2) a generalized linear regression. Given $\sum$, the Aitken Generalized Least Squares estimator of all parameters in the system can be expressed in the following form:

$$\hat{\alpha} = (X'(\sum \otimes I_n)^{-1} X)^{-1} X'(\sum \otimes I_n)^{-1} y.$$

In the M_0 case, where $\sum = \mathrm{diag}(\sigma_{11}^2, \ldots, \sigma_{nn}^2)$ we have:

$$\hat{\alpha} = \hat{\alpha}_{OLS} = (X'X)^{-1} X'y, \tag{4}$$

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which is equivalent to the application of the OLS estimator to each equation separately. In the general case, M_1, we have to estimate the covariance matrix $\sum$. In the empirical part of the paper, we apply the Zellner (1962) method, and estimate elements of matrix $\sum$ on the basis of OLS residuals, denoted by $\hat{\varepsilon}_{[nTx1]} = (\hat{\varepsilon}^{(1)\prime}, \ldots, \hat{\varepsilon}^{(n)\prime})$. The Estimated GLS, elaborated by Zellner (1962) takes the following form:

$$\hat{\alpha}_{EGLS} = (X'(S \otimes I_n)^{-1} X)^{-1} X'(S \otimes I_n)^{-1} y. \quad (5)$$

where:

$$S = \frac{1}{T}(\hat{\varepsilon}^{(1)}, \ldots, \hat{\varepsilon}^{(n)})'(\hat{\varepsilon}^{(1)}, \ldots, \hat{\varepsilon}^{(n)}).$$

The nondiagonal matrix S explains correlations between error terms from different equations. Hence, it can be treated as a measure of the strength of cross-country linkages. In the empirical part of the paper, we discuss the importance of the SURE specification in explaining the heterogeneity of the relationship between loans and the banking sector – i.e. their specific determinants.

4. RESULTS AND DISCUSSION

We constructed a balanced panel dataset using annual data from 2005 to 2020 for 13 OECD countries, specifically Belgium, Canada, Denmark, France, Germany, Italy, the Netherlands, Norway, Poland, Spain, Sweden, Switzerland, and the United States. The countries were chosen to ensure comparability with the study conducted by Olszak and Pipień (2016). Descriptive statistics for the variables employed in our analysis are presented in Tables 2 and 3. On average, the loan-to-assets ratio in the banking sectors of our sample was at 52% over the entire period, although the ratio varied considerably across countries, ranging from 5.7% to 93.2%. Notably, Sweden, Poland, and Denmark exhibited high loan-to-assets ratios, with values of 71.9%, 65.5%, and 63.7%, respectively, while Switzerland and France had lower proportions of 32.4% and 38.8%, respectively. The levels of profits, deposits, and loan loss provisions relative to total assets also displayed significant variation across countries.

Table 2
Descriptive statistics of the analyzed series

	LOANS/TA	GDPG	INFLATION	UNEMPLOYMENT	PROFIT/TA	LLP/TA	DEPOSIT/TA
Mean	0.515511	1.31521	1.515383	7.474712	0.006587	0.008544	0.383678
Median	0.510628	1.842526	1.531628	6.975000	0.006352	0.005285	0.388337
Maximum	0.931885	7.061544	4.489444	26.09000	0.019751	0.040059	0.73337
Minimum	0.056583	–10.8229	–1.14391	2.490000	–0.012108	0.000416	0.02884
Std. Dev.	0.137083	2.616323	1.13641	3.981535	0.004762	0.008522	0.154807
Skewness	0.070785	–1.50307	0.12625	2.200512	–0.306749	1.734758	0.142825
Kurtosis	3.009366	6.590182	2.649199	9.234274	4.509617	5.408717	2.414647
Jarque-Bera	0.17446	190.0274	1.619088	504.705	23.0128	154.6087	3.676701
Probability	0.916466	0.000000	0.445061	0.000000	0.00001	0.000000	0.15908
Sum	107.2264	273.5636	315.1997	1554.74	1.37001	1.77725	79.80498
Sum Sq. Dev.	3.889871	1416.945	267.3254	3281.493	0.004693	0.015033	4.960808
No. of observations	208	208	208	208	208	208	208

Source: Author's own calculation.

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Table 3
Country-wise statistics of the analyzed series

Country	Avg. of LOANS/TA	Avg. of GDPG	Avg. of INFLATION	Avg. of UNEMPLOYMENT	Avg. of PROFIT/TA	Avg. of LLP/TA	Avg. of DEPOSIT/TA
Belgium	0.49	1.03	1.86	7.45	0.01	0.01	0.41
Canada	0.49	1.94	1.70	6.98	0.01	0.00	0.57
Denmark	0.64	1.14	1.43	5.88	0.01	0.01	0.16
France	0.39	0.64	1.22	8.93	0.00	0.01	0.30
Germany	0.44	1.12	1.39	6.03	0.00	0.00	0.40
Italy	0.59	0.53	1.35	9.49	0.01	0.02	0.32
Netherlands	0.50	1.16	1.60	5.14	0.00	0.00	0.43
Norway	0.45	1.30	2.06	3.61	0.01	0.00	0.19
Poland	0.65	3.59	2.08	8.38	0.01	0.03	0.67
Spain	0.57	0.51	1.59	17.21	0.01	0.01	0.49
Sweden	0.72	1.80	1.17	7.39	0.01	0.00	0.24
Switzerland	0.32	1.83	0.28	4.45	0.01	0.00	0.41
United States	0.45	1.57	1.99	6.23	0.01	0.01	0.40

Source: Author's own calculation.

Among the three macroeconomic determinants examined, the mean levels of two variables, GDP growth and unemployment rates, were relatively comparable at 1.3% to 1.5%. However, the GDP growth rate exhibited greater diversity across countries, with minimum and maximum values of –10.82% and 7.06%, respectively. The average GDP growth rate across all countries during the analysis period was 1.32%. Italy was the only country with a negative average GDP growth rate (–0.53%) during the period, while Poland experienced the highest average GDP growth rate (3.59%). The average unemployment rate across the sample was 7.47%, showing substantial heterogeneity across countries, with the possibility of outliers indicated by a sample kurtosis greater than 9. Spain recorded the highest average unemployment rate (17.2%), whereas Norway had the lowest average unemployment rate (3.6%).

Table 4 presents the correlation matrix of the regression variables. The overall sample suggests relatively weak associations between the loan-to-assets ratio and the explanatory variables. Notably, the variables representing the banking sector's condition, namely profits, deposits, and loan loss provisions, exhibit strong correlations. Surprisingly, the variables do not display as strong a correlation with the loan-to-assets ratio as expected. However, as we will demonstrate later in the article, reporting weak relationships between the analyzed variables would be erroneous. The nature of the relationships between the variables is primarily driven by cross-country heterogeneity, as revealed within the SURE model.

Table 4
Matrix of sample correlation of the analyzed series

	LOANS/TA	GDPG	INFLATION	UNEMPLOYMENT	PROFIT/TA	LLP/TA	DEPOSIT/TA
LOANS/TA	1	0.0139	0.0437	0.0373	0.0051	0.0258	0.0070
GDPG	0.0139	1	0.0721	0.0398	0.0128	0.0384	–0.0067
INFLATION	0.0437	0.0721	1	0.0006	–0.0002	0.0044	0.0003
UNEMPLOYMENT	0.0373	0.0398	0.0006	1	0.0020	–0.0002	0.0004
PROFIT/TA	0.0051	0.0128	–0.0002	0.0020	1	–0.3626***	0.5667***
LLP/TA	0.0258	0.0384	0.0044	–0.0002	–0.3626***	1	–0.4158***
DEPOSIT/TA	0.0070	–0.0067	0.0003	0.0004	0.5667***	–0.4158***	1

Note: *** denotes significance at 1% level.

Source: Author's own calculation.

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Tables 5, 6, and 7 present estimation results of the parameters in equations (1) and (2). Initially, we estimated the parameters in equation (1) using fixed-effect (FE) panel regression techniques, as shown in Table 5. In the case, the FE approach does not account for cross-country diversity in the impact of explanatory variables on the loan-to-assets ratio. Among all the factors that potentially influence loan variability, profits and loan loss provisions are empirically important, as indicated by statistically significant parameter estimates. Surprisingly, deposits do not appear as significant as the LLP measure and profits. In contrast, macroeconomic variables do not contribute significantly to the loan-to-assets ratio, with relatively small and insignificant point estimates for the corresponding parameters.

Table 5

Determinants of $\frac{LOANS_{t,j}}{TA_{t,j}}$ – the FE panel estimates of parameters in equation (1)

	Intercept	GDPG	INFLATION	UNEMPLOYMENT	PROFIT/TA	LLP/TA	DEPOSIT/TA
FE panel regression estimates							
Est.	0.4416***	–0.0040	0.00357	0.002161	9.2319***	5.3059***	–0.1267***
Std. error	0.0315	0.00377	0.00794	0.00259	2.0830	1.3123	0.0661
p-value	0.0000	0.2970	0.6561	0.4111	0.0001	0.0003	0.0643

Note: *** denotes significance at 1% level.

Source: Author's own calculation

The empirical insight into cross-country heterogeneity is presented in Tables 6 and 7. In Table 6, we report estimation outcomes for parameters in models M_0 and M_1. In model M_0, which assumes no correlations between $\varepsilon_{t,j}$ and $\varepsilon_{s,i}$ in the system of equations (2) (model M_0) we run separate OLS regressions for each country, treating the regression for each country independently and disregarding interactions among equations. Table 7 presents the results of estimation in the case of M_1, which accounts for a non-diagonal covariance matrix in the system (2) using the Zellner (1962) estimator. In both cases, M_0 and M_1, the data strongly supports cross-country diversity in the relationships between the loan-to-assets ratio and other variables of interest. In most cases, the relationships are statistically significant, as measured by the corresponding parameter estimates. The lack of significance between the GDP growth rate and the loan-to-assets ratio in the case of the United States is noteworthy. There are also countries where fluctuations in economic growth hurt the loan-to-assets ratio. Among the countries, which include Canada, Denmark, Spain, Sweden, and Switzerland, only Sweden exhibits vital significance, with some evidence for Switzerland.

Table 6

Cross country heterogeneity of determinants of $\frac{LOANS_{t,j}}{TA_{t,j}}$ – estimates of parameters in equation (2) provided the model M_0 (independent regressions)

Country		Intercept	GDPG	INFLATION	UNEMPLOYMENT	PROFIT/TA	LLP/TA	DEPOSIT/TA
The system of independent regressions; M_0								
Belgium	Est.	–0.0711	0.0047	–0.0021	0.0350	–6.1422	–2.2310	0.8253
	Std. error	0.0866	0.0018	0.0036	0.0076	1.4003	3.2544	0.1350
	p-value	0.4178	0.0136	0.5611	0.0001	0.0001	0.4979	0.0000
Canada	Est.	0.4600	–0.0019	–0.0124	–0.0086	5.9418	–32.8700	0.3281
	Std. error	0.1550	0.0039	0.0092	0.0066	4.6054	6.6498	0.2875
	p-value	0.0056	0.6270	0.1898	0.2052	0.2062	0.0000	0.2623

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Table 6 – continued

Country		Intercept	GDPG	INFLATION	UNEMPLOYMENT	PROFIT/TA	LLP/TA	DEPOSIT/TA
Denmark	Est.	0.9561	–0.00086	–0.0121	–0.0087	0.0913	–0.1550	–1.5486
	Std. error	0.0305	0.0016	0.0042	0.0084	1.7588	4.5363	0.1063
	p-value	0.0000	0.5871	0.0068	0.3094	0.9589	0.9730	0.0000
France	Est.	0.2216	0.0042	–0.0096	–0.0254	0.2904	19.75	0.6941
	Std. error	0.0358	0.0014	0.0047	0.0042	1.6235	1.2707	0.0628
	p-value	0.0000	0.0055	0.0510	0.0000	0.8591	0.0000	0.0000
Germany	Est.	–0.5604	0.0091	–0.0308	0.0274	–28.93	9.3119	2.2475
	Std. error	0.2000	0.0035	0.1350	0.0059	7.9103	8.6623	0.3744
	p-value	0.0086	0.0149	0.0299	0.0001	0.0009	0.2904	0.0000
Italy	Est.	0.2642	–0.0031	0.0539	0.0090	17.35	1.2588	0.1582
	Std. error	0.2320	0.0111	0.0269	0.0342	7.7527	6.5491	0.3813
	p-value	0.2632	0.7845	0.0537	0.7949	0.0323	0.8488	0.6811
Netherlands	Est.	–0.0664	0.0041	0.0178	–0.0026	1.5027	29.45	1.0088
	Std. error	0.0921	0.0030	0.0095	0.0082	1.5549	9.24	0.1682
	p-value	0.4762	0.1916	0.0713	0.7532	0.3411	0.0032	0.0000
Norway	Est.	0.0257	0.0307	0.0008	–0.0072	–4.32	90.42	1.2583
	Std. error	0.1437	0.0157	0.0180	0.0305	6.57	26.1949	0.3418
	p-value	0.8594	0.0600	0.9636	0.8155	0.5149	0.0016	0.0008
Poland	Est.	1.0974	–0.0027	–0.0164	–0.0230	6.29	15.83	–1.0482
	Std. error	0.1476	0.0043	0.0054	0.0045	0.95	2.696	0.2543
	p-value	0.0000	0.5432	0.0048	0.0000	0.0411	0.0000	0.0002
Spain	Est.	0.0168	–0.0012	–0.0022	–0.0157	7.7049	26.24	0.8151
	Std. error	0.0330	0.0013	0.0039	0.0016	1.0477	1.5228	0.0654
	p-value	0.6140	0.3455	0.5852	0.0000	0.0000	0.0000	0.0000
Sweden	Est.	0.3220	–0.0146	0.00072	0.0062	54.94	40.04	–0.5960
	Std. error	0.1162	0.0025	0.0069	0.0111	6.08	8.69	0.1309
	p-value	0.0092	0.0000	0.9172	0.5829	0.0000	0.0001	0.0001
Switzerland	Est.	0.0085	–0.0056	0.0103	0.0110	–0.2791	9.24	0.6449
	Std. error	0.0478	0.0025	0.0059	0.0124	0.7271	13.64	0.0588
	p-value	0.8604	0.0306	0.0902	0.3781	0.7040	0.4948	0.0000
United States	Est.	0.4276	–0.00012	0.0062	–0.0273	4.8280	22.27	–0.0771
	Std. error	0.1156	0.0081	0.0055	0.0112	3.4537	8.1279	0.2491
	p-value	0.0008	0.9884	0.2682	0.0205	0.1718	0.0100	0.7590

Source: Author's own calculation.

Table 7

Cross country heterogeneity of determinants of $\frac{LOANS_{t,j}}{TA_{t,j}}$ – estimates of parameters in equation (2) provided the model M_1 (SURE specification)

Country		Intercept	GDPG	INFLATION	UNEMPLOYMENT	PROFIT/TA	LLP/TA	DEPOSIT/TA
The SURE model; M_1								
Belgium	Est.	–0.0604	0.0038	–0.0011	0.0342	–5.5754	–1.2658	0.7889
	Std. error	0.0668	0.0015	0.0025	0.0059	1.1326	2.5806	0.1065
	p-value	0.3726	0.0145	0.6572	0.0000	0.0000	0.6271	0.0000
Canada	Est.	0.4732	–0.0019	–0.0122	–0.0125	5.1763	–33.38	0.3694
	Std. error	0.0797	0.0013	0.0033	0.0028	1.9310	4.2648	0.1400
	p-value	0.0000	0.1338	0.0009	0.0001	0.0115	0.0000	0.0128

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Table 7 – continued

Country		Intercept	GDPG	INFLATION	UNEMPLOYMENT	PROFIT/TA	LLP/TA	DEPOSIT/TA
Denmark	Est.	0.9597	–0.0010	–0.0134	–0.0081	–0.2718	–0.3267	–1.5613
	Std. error	0.0280	0.0015	0.0038	0.0075	1.4402	3.9712	0.0955
	p-value	0.0000	0.5003	0.0012	0.2906	0.8515	0.9349	0.0000
France	Est.	0.1982	0.0039	–0.0059	–0.0219	0.7255	18.6232	0.6842
	Std. error	0.0259	0.0009	0.0032	0.0026	1.0701	0.9052	0.0474
	p-value	0.0000	0.0002	0.0760	0.0000	0.5027	0.0000	0.0000
Germany	Est.	–0.5215	0.0074	–0.0304	0.0229	–21.8209	13.2469	2.1279
	Std. error	0.1187	0.0022	0.0080	0.0040	5.1639	5.1010	0.2231
	p-value	0.0001	0.0021	0.0006	0.0000	0.0002	0.0141	0.0000
Italy	Est.	0.2865	0.0068	0.0416	0.0134	14.3284	–2.0080	0.2857
	Std. error	0.1112	0.0071	0.0146	0.0155	3.3844	2.8156	0.2394
	p-value	0.0148	0.3424	0.0078	0.3936	0.0002	0.4809	0.2413
Netherlands	Est.	–0.0428	0.0032	0.0187	–0.0024	2.0691	20.6508	1.0215
	Std. error	0.0672	0.0026	0.0069	0.0061	1.1665	6.7065	0.1202
	p-value	0.5291	0.2156	0.0108	0.7002	0.0856	0.0042	0.0000
Norway	Est.	0.1296	0.0240	0.0053	–0.0298	–5.1175	79.3370	1.3000
	Std. error	0.0913	0.0121	0.0124	0.0202	4.0089	19.3885	0.2268
	p-value	0.1655	0.0568	0.6748	0.1497	0.2110	0.0003	0.0000
Poland	Est.	0.9508	0.0007	–0.0113	–0.0177	3.7840	13.4419	–0.7867
	Std. error	0.1098	0.0030	0.0042	0.0034	2.0948	2.0524	0.1912
	p-value	0.0000	0.8195	0.0112	0.0000	0.0803	0.0000	0.0003
Spain	Est.	0.0024	–0.0008	–0.0009	–0.0164	7.3853	25.5120	0.8878
	Std. error	0.0249	0.0010	0.0027	0.0011	0.6882	1.1137	0.0541
	p-value	0.9242	0.4644	0.7514	0.0000	0.0000	0.0000	0.0000
Sweden	Est.	0.4028	–0.0144	–0.0030	–0.0064	54.8008	39.9204	–0.5237
	Std. error	0.0790	0.0017	0.0048	0.0071	3.9758	4.5496	0.0926
	p-value	0.0000	0.0000	0.5281	0.3782	0.0000	0.0000	0.0000
Switzerland	Est.	0.0187	–0.0039	0.0041	0.0058	–0.2482	22.91	0.6261
	Std. error	0.0299	0.0019	0.0037	0.0075	0.3379	9.5677	0.0411
	p-value	0.5364	0.0510	0.2733	0.4476	0.4679	0.0226	0.0000
United States	Est.	0.3526	0.0060	0.0060	–0.0157	2.4820	14.7957	0.0993
	Std. error	0.0803	0.0058	0.0039	0.0075	2.3874	5.6534	0.1718
	p-value	0.0001	0.3031	0.1354	0.0441	0.3063	0.0134	0.5673

Source: Author's own calculation.

Tables 8 and 9 present the statistical significance and the direction of the relationship between the loan-to-assets ratio and explanatory variables. The tables provide qualitative insights into the procyclicality effects and the strength of the analyzed linkages, complementing the information from Tables 6 and 7. Initially, when examining the role of economic growth in explaining loan fluctuations, we reported a very weak, slightly negative, but statistically insignificant impact in the FE panel regression outcomes shown in Table 5. However, according to the system (2) in both stochastic settings (M_0 and M_1), the relationship between economic growth and loans exhibits substantial diversity across countries. Overall, the SURE model (M_1) provides more precise estimates, resulting in stronger inferences about the statistical significance of the parameters compared to the independent regressions in M_0. Among the countries with a positive impact of

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economic growth fluctuations on the loan-to-assets ratio there are Belgium, France, Germany, Norway, and the United States. However, based on model M_1, the impact can be considered decisively significant only for Belgium, France, and Germany. The lack of significance between the growth rate of GDP and loans (to total assets) in the case of the United States is worth noting. On the other hand, there are countries where economic growth fluctuations have a negative impact on the loan-to-assets ratio. Table 9 shows that among the countries, including Canada, Denmark, Spain, Sweden, and Switzerland, only in the case of Sweden can we report strong significance, with some evidence also found for Switzerland.

Table 8
The sign and significance of impact in (2) – the model M_0 (independent regressions)

Country	Intercept	GDPG	INFLATION	UNEMPLOYMENT	PROFIT/TA	LLP/TA	DEPOSIT/TA
Belgium	–	+*	–	+***	–***	–	+***
Canada	+	–	–	–*	+	–***	+
Denmark	+***	–	–***	–	+	–	–***
France	+***	+***	–*	–***	+	+***	+***
Germany	-***	+**	–**	+***	–***	+	+***
Italy	+	–	+*	+	+**	+	+
Netherlands	–	+	+*	–	+	+***	+***
Norway	+	+*	+	–	–	+***	+***
Poland	+***	–	–***	–***	+**	+***	–***
Spain	+***	–	–	–***	+***	+***	+*
Sweden	+***	–***	+	+	+***	+***	–***
Switzerland	+	–**	+*	+	–	+	+***
United States	+***	–	+	–**	+	+***	–

Note: A particular variable's positive/negative impact is denoted by +/– respectively. We also put notation reporting the significance at levels 0.01, 0.05 and 0.1 by ***, ** and *.
Source: Author's own calculation.

Table 9
The sign and significance of impact in (2) – the model M_1 (independent regressions)

Country	Intercept	GDPG	INFLATION	UNEMPLOYMENT	PROFIT/TA	LLP/TA	DEPOSIT/TA
Belgium	–	+**	–	+***	–***	–	+***
Canada	+***	–	–***	–***	+**	–***	+**
Denmark	+***	–	–***	–	–	–	–***
France	+***	+***	+*	–***	+	+***	+***
Germany	–***	+***	–***	+***	–***	+**	+***
Italy	+**	+	+***	+	+***	–	+
Netherlands	–	+	+**	–	+*	+***	+***
Norway	+	+*	+	–	–	+***	+***
Poland	+***	+	–**	–***	+*	+***	–***
Spain	+	–	–	–***	+***	+***	+***
Sweden	+***	–***	–	–	+***	+***	–***

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Table 9 – continued

Country	Intercept	GDPG	INFLATION	UNEMPLOYMENT	PROFIT/TA	LLP/TA	DEPOSIT/TA
Switzerland	+	–*	+	+	–	+**	+***
United States	+***	+	+	–**	+	+**	+

Note: A particular variable's positive/negative impact is denoted by +/– respectively. We also put notation reporting the significance at levels 0.01, 0.05 and 0.1 by ***, ** and *.

Source: Author's own calculation

Regarding the impact of inflation on the loan-to-assets ratio, the FE estimates presented in Table 5 indicate insignificance. However, when considering model M_1, we observe some diversity. Negative and statistically significant impacts are found for Canada, Denmark, and Germany, while positive and statistically significant influences are reported for Italy and the Netherlands. In the case of other countries, the relationship between inflation and loans (to total assets) is insignificant.

Unemployment rate is the third and final macroeconomic variable in our analysis. Like the GDP growth rate and inflation, the FE panel estimates indicate the empirical insignificance of the relationship between the unemployment rate and the loan-to-assets ratio. According to Table 9, in the SURE model (M_1), the predominant statistically significant impacts are negative and can be attributed to Canada, France, Poland, Spain, and, to some extent, the United States. On the other hand, Belgium and Germany exhibit a positive and significant impact of the unemployment rate on the loan-to-assets ratio.

In addition to the variables discussed earlier, we also considered some observed categories representing the activity of the banking sector as explanatory variables in the panel regression (1) and the system (2). Among the variables, profits and loan loss provisions (relative to total assets) exhibit significance in the FE panel regression at the 0.01 level. The results obtained for model M_1 shed light on the nature of the analyzed relationships. The impact of profits on loans (relative to total assets) is positive and statistically significant, at least at the 0.1 level, for Canada, Italy, the Netherlands, Poland, Spain, and Sweden. However, in the cases of Belgium and Germany, the impact is also statistically significant but negative. Similarly, regarding the relationship between loan loss provisions and loans (relative to total assets), model M_1 (Table 9) indicates an increasing relationship for nine countries. In most cases, the identified associations are characterized by statistically significant parameters in equation (2). The only substantial evidence supporting a negative impact of loan loss provisions on loans (both relative to total assets) is found for Canada. FE panel regression estimates in Table 5 show a negative relationship between deposits and loans (relative to total assets), which is statistically significant at a level no smaller than 0.1. The SURE specification (M_1) strengthens the level of statistical significance for negative relationships in Denmark, Poland, and Sweden, as compared to the FE panel outcomes.

Table 10 presents estimation results for the elements of the covariance matrix Σ, which are necessary for the SURE specification. The point estimates of the variances (shown in bold font), contemporaneous covariances (shown in italics, above the diagonal), and contemporaneous correlations (displayed below the diagonal) of error terms are reported. Analyzing the correlation estimates, it becomes evident that the system regression approach employed in model M_1 is empirically important. SA simple analysis based on country-independent regressions, which formally assume a diagonal covariance matrix, overlooks substantial cross-country financial linkages. The strongest correlations in the system are positive, with some exceptions. The strongest financial linkage, as measured by the correlation of the error terms, is observed between the United States and Norway (0.91). Additionally, pairs such as Sweden-Canada, Spain-Italy, Poland-the Netherlands, Sweden-the Netherlands, and Sweden-Poland exhibit positive correlations exceeding 0.5. A few exceptions with strong negative correlations include Germany-Canada, France-Switzerland, Italy-Germany, Germany-Spain, and Italy-Switzerland.

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Table 10
Point estimates of the variances (bold), contemporaneous covariances (in italics above the diagonal) and contemporaneous correlations (shaded cells below the diagonal) of error terms in case of model M_1 (SURE specification), calculated based on matrix S

Country	Belgium	Canada	Denmark	France	Germany	Italy	The Netherlands	Norway	Poland	Spain	Sweden	Switzerland	US
Belgium	**0.0001150**	*0.0000284*	*–0.0000005*	*0.0000195*	*0.0000295*	*–0.0004204*	*0.0000106*	*0.0001407*	*0.0000063*	*–0.0000730*	*0.0000212*	*0.0000446*	*0.0000938*
Canada	0.1658733	**0.0002544**	*–0.0000120*	*0.0000113*	*–0.0003180*	*0.0006955*	*0.0001168*	*0.0002504*	*0.0002279*	*0.0000641*	*0.0002168*	*–0.0000508*	*0.0000757*
Denmark	–0.0042713	–0.0672037	**0.0001244**	*–0.0000071*	*–0.0000248*	*–0.0001280*	*–0.0000525*	*0.0000141*	*–0.0000506*	*–0.0000253*	*–0.0000972*	*0.0000068*	*–0.0000109*
France	0.2229476	0.0870313	–0.0781787	**0.0000668**	*0.0000512*	*–0.0000185*	*0.0000089*	*0.0002016*	*0.0000648*	*0.0000112*	*0.0000036*	*–0.0000596*	*0.0000760*
Germany	0.1036583	–0.7518819	–0.0836982	0.2361591	**0.0007032**	*–0.0012450*	*–0.0002041*	*0.0003153*	*–0.0003779*	*–0.0002156*	*–0.0001372*	*0.0000940*	*0.0001156*
Italy	–0.4337105	0.4822548	–0.1268689	–0.0249895	–0.5192681	**0.0081748**	*–0.0004161*	*0.0013682*	*–0.0001995*	*0.0008029*	*–0.0001913*	*–0.0008492*	*0.0001446*
Netherlands	0.0475832	0.3533172	–0.2272467	0.0528037	–0.3714958	–0.2220645	**0.0004294**	*–0.0004881*	*0.0004020*	*–0.0000108*	*0.0002761*	*0.0000205*	*–0.0001140*
Norway	0.2395314	0.2864947	0.0230606	0.4500130	0.2169736	0.2761481	–0.4298034	**0.0030028**	*–0.0003864*	*–0.0000689*	*0.0001882*	*–0.0001493*	*0.0010126*
Poland	0.0204586	0.4947619	–0.1571236	0.2745120	–0.4934451	–0.0763929	0.6717509	–0.2441794	**0.0008339**	*0.0002192*	*0.0003546*	*–0.0000383*	*–0.0001089*
Spain	–0.4285458	0.2526436	–0.1426980	0.0862174	–0.5114171	0.5585861	–0.0326624	–0.0790681	0.4775088	**0.0002527**	*–0.0000071*	*–0.0001117*	*–0.0000353*
Sweden	0.0888881	0.6100625	–0.3911547	0.0197464	–0.2322535	–0.0949728	0.5980872	0.1541711	0.5510650	–0.0200013	**0.0004965**	*0.0000820*	*0.0000959*
Switzerland	0.3210598	–0.2458503	0.0472764	–0.5623761	0.2734880	–0.7246409	0.0763722	–0.2102010	–0.1023475	–0.5419391	0.2839402	**0.0001680**	*–0.0000226*
US	0.4291456	0.2328357	–0.0480180	0.4557898	0.2138238	0.0784228	–0.2697751	0.9064217	–0.1849465	–0.1090304	0.2110975	–0.0853804	**0.0004156**

Source: Author's own calculation.

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5. CONCLUSIONS

The paper utilizes a balanced panel database of aggregated financial statements from the banking sector in 13 OECD countries (Belgium, Canada, Denmark, France, Germany, Italy, the Netherlands, Norway, Poland, Spain, Sweden, Switzerland, and the United States) to examine procyclicality of credit supply by banks from 2005 to 2020. We investigate bank loan determinants by considering bank-specific and macroeconomic variables. Firstly, we address the critical question of identifying determinants of bank loans at the country level using panel regression analysis. Secondly, we employ the Seemingly Unrelated Regression Equations (SURE) methodology to explore the impact of interconnectedness among countries on the diversity of the strength of the procyclicality of bank loans.

Our research contributes to the existing literature on the procyclicality of bank loans by utilizing the SURE approach, which allows us to empirically measure the interconnectedness between countries as a determinant of bank loans. As compared to panel regression models, which serve as the reference econometric framework, and the regression analysis conducted independently for each country, applying the SURE model enhances the statistical significance of the business cycle and banking sector-specific variables in countries with a procyclical effect.

Our findings provide empirical evidence supporting the procyclicality of loans in 6 out of the 13 countries included in the analysis. Furthermore, we observe that bank-specific variables have greater significance as loan supply determinants than macroeconomic variables.

A potential path for further research is to analyze whether the procyclicality of bank loans differs based on the type of loans, such as consumer loans, commercial loans, and residential mortgage loans.

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Employment effects of minimum wage changes across regions, age groups, and sectors

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ABSTRACT

Increases in minimum wages in many developed and developing economies in recent years raise the question of whether and how they impact employment. We analyze the employment effects of minimum wage increases for different age groups of workers simultaneously. We construct a panel using three-dimensional cells formed by three age groups, two economic sectors, and 16 regions, separately for each year. We use individual data on employee and employer characteristics from the Structure of Earnings Survey and aggregated data from the Local Data Bank in Poland. The research period covers 2006–2020.

Our results confirm the differences in employment elasticity for different groups of workers. We discover latent heterogeneities with regions simultaneously experiencing both negative and positive employment effects of minimum wage changes for different groups of workers and sectors. Negative employment effects are observed mostly for youths, positive employment effects are predominant in the groups of workers aged 50 and over. The employment reaction to changes in the minimum wage is the result of a combination of regional labor market features. Negative employment effects are more likely in areas with larger proportion of workers in the private sector, in industries in which it is more difficult to increase the prices of goods or services produced, and where small firms are widespread.

The results show that previous analyses at the aggregated level might underestimate the employment effects of the minimum wage. The results also show that the overall minimum wage effects cannot be easily predicted by policymakers.

JEL classification: J21, R23, J31, J38

Keywords: employment elasticity, minimum wage, regional labor markets, multidimensional panel analyses, intra-regional differences, Poland.

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1. INTRODUCTION

Increases in minimum wages in many developed and developing economies in recent years raise the question of whether and how they impact employment. The importance of the topic has grown in times of high inflation and minimum wage raises through various indexation mechanisms. The amount of research on the minimum wage–employment relationship is enormous, but neither the direction of the relationship nor its strength has been unequivocally determined.

Increases in the minimum wage raise production costs. There are a few ways that firms can deal with higher labor costs, including reducing employment or the non-financial benefits for workers to decrease total costs. They may pass higher labor costs on to prices, increasing their income. Firms can also maintain employment levels, non-financial benefits, and prices, but then their markups and profits would be reduced (Lemos, 2008).

The overall effect of minimum wage increases depends on several factors. Since employment reduction is costly, firms first insist on passing the costs to consumers by increasing prices of their products (Harasztosi and Lindner, 2019). Therefore, negative employment effects would be expected in those labor market segments where prices cannot be increased or where worker turnover costs are low, e.g., firms in tradable sectors, firms facing high market competition, small firms, and firms that employ young, low-educated, and less experienced workers.

Literature shows that negative employment effects are observed among less-skilled and less-experienced (younger) workers (see, e.g., Kiss, 2018; Wolfson and Belman, 2019; Neumark and Shirley, 2021). However, the size of the effect depends on the market competition. Munguía Corella (2020) found significant negative employment effects due to minimum wage changes for youth under perfect competition, and insignificant effects under full monopsonistic labor markets. Harasztosi and Linder (2019) found that employment reaction varies across countries and industries, and that unemployment effects were greater in industries that had more difficulty passing wage costs onto consumers. Moreover, a growing number of authors underline heterogeneity of labor markets across regions as the main source of non-significant employment elasticity at the aggregate level (Thompson, 2010; Autor et al., 2019). All of the studies confirmed differences in regional employment reactions due to minimum wage increases, but they did not indicate the reasons for the differences.

Our study follows the approach suggested by Card (1992), which relies on the extent to which regional labor markets are affected by the minimum wage. The minimum wage is intended to affect less skilled and less experienced workers. Therefore, the uneven distribution of young or less-educated workers across economic sectors and regions may result in differences in how minimum wage changes impact employment in particular sectors of regional labor markets. Additionally, structures of both employers and employees may determine the extent to which minimum wage increases will affect employment.

Literature on minimum wage effects on employment is abundant but prior analyses are partial in nature. They usually considered only one or two factors (dimensions), investigating employment effects of minimum wage increases over time and across age groups, educational groups, sectors, or regions separately. Additionally, due to data limitations, effects at the firm level were typically analyzed in isolation from analyses that used workers' characteristics. This paper addresses the gaps by studying employment effects of minimum wage increases on most important factors simultaneously. It creates a novel four-dimensional approach. We estimate the employment elasticities of minimum wage increase for different segments of the labor market in Poland and analyze the factors behind them. Our aim is to investigate and explain employment effects of minimum wage increases by age groups, taking into account sectoral and regional structure of the workforce simultaneously. In particular, we want to explain why in some regions negative employment effects for young workers are observed, while in others they

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are not. The value added of the study is finding and explaining heterogeneities among minimum wage elasticities, it would have been impossible without implementing our four-dimensional approach.

We estimate that employment effects of minimum wages increase for different age groups of workers. In particular we aim to verify the following research hypotheses: (1) In all regions and sectors analyzed negative employment effects in the group of young workers are observed; in the case of other age groups employment effects are not significant. (2) Negative employment effects are more pronounced in the industry sector than in the market services due to higher international competition and lower possibilities to pass higher costs to consumers. (3) Negative employment effects are higher in regions with higher share of workers in industries facing higher international competition.

As a case study we use Polish data. It is worth exploring the topic using Poland for several reasons. First, the minimum wage policy conducted at a national level is simple and has a long history; moreover, there is one minimum wage rate for all regions, occupations, and sectors. Second, according to Eurostat data, Poland's share of minimum wage workers is one of the highest of all European economies. Third, there has been a sustained increase in the national minimum wage in Poland in recent years. Fourth, Poland is one of the largest EU economies, and the minimum wage coverage is extensive. Finally, Poland exhibits large and enduring regional differences.

We use individual data on employee and employer characteristics from the Structure of Earnings Survey, supplementing them with regional data from the Local Data Bank in Poland. The research period covers 2006–2020. We construct a panel using three-dimensional cells formed by three age groups, two economic sectors, and 16 regions, separately for each year. The cells are our units of observation. Using a cell-level approach allows for multiple factors to be taken into account simultaneously. This is a novel approach.

We begin by estimating the average employment elasticity for the whole sample of workers before applying the slope homogeneity test for panel data developed by Blomquivst and Westerlund (2013). After rejecting the homogeneity of the employment effect, we allow the parameter of the minimum wage variable to vary across cells (age group, economic sector, and region simultaneously). In the second stage, we try to explain differences in the minimum wage elasticity estimates. We apply cluster analysis to the three-dimensional cells of workers. Finally, we verify how different labor market structures affect employment reactions to minimum wage changes.

The multidimensional approach has an inevitable advantage over prior studies in that it allows for a more detailed picture of the analyzed phenomenon. The paper makes several contributions to the minimum wage literature. We study differences among regions together with heterogonous reactions to policy changes within regions, observing how regional differences in sectoral and age composition of the workforce affect employment reaction to minimum wage increases. To the best of our knowledge, it is the first study of its kind.

Our results confirm differences in employment elasticity for minimum wages across regions. We also discover latent heterogeneities in the regional employment effect, with regions simultaneously experiencing both negative and positive employment effects of minimum wage changes for different groups of workers and sectors. Negative employment effects are observed mostly for the youth, while positive employment effects are predominantly in the groups of workers aged 50 and over. Stronger negative effects are observed in the industry than market services sector.

We found that the employment effect of changes in minimum wage levels is the result of a combination of regional labor market features. Negative employment effects are more probable in regions with small, private sector firms in the tradable sector, where it is more difficult to increase prices of goods or services produced. Conversely, positive employment effects are

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more probable in regions with a high share of workers employed in the public sector or in large enterprises. Significantly, the two completely different labor market environments can coexist within a given region, which may explain why empirical analyses at a regional level often indicate insignificant values of employment elasticity for minimum wage changes. We have not found similar findings in the literature.

The remainder of the study proceeds as follows. Section 2 contains a literature review. Section 3 describes data and an empirical approach. Subsequently, Section 4 reports results and robustness analyses. Section 5 concludes.

2. LITERATURE REVIEW

Considerable research has been conducted on the relationship between minimum wage changes and employment; however, there is still an ongoing debate on the direction and strength of the relationship. Wolfson and Belman (2019) and Neumark and Shirley (2021) present the most recent summaries of evidence from the US. Campolieti (2020) provides a meta-analysis for Canada, and Dube (2019) summarizes the international evidence. Broecke, Forti and Vandeweyer (2017) and Neumark and Mungiua Corella (2021) studied employment effects of minimum wages in developing countries. Most research indicates a negative impact of minimum wage growth on employment among the most vulnerable groups of workers, i.e., the young and less educated (see, e.g., Kiss, 2018 or Marimpi and Koning, 2018).

In theory, negative employment effects are expected in a competitive price-taker setting, but the effect of minimum wages is ambiguous under monopsonistic labor markets. Manning's (2003) model indicates three possible scenarios: (1) firms are unconstrained because the minimum wage is not binding; (2) firms are supply-constrained and increases in minimum wages have positive effects on employment; and (3) firms are demand-constrained, and a high minimum wage has a negative effect on employment (Munguía Corella, 2020).

Many studies have used the monopsony model to explain non-negative results, including Katz and Krueger (1992) and Card and Krueger (1994), and more recently, Dube, Lester, and Reich (2010). Azar et al. (2019) provided empirical evidence to support the monopsony model as an explanation for the near-zero minimum wage employment effect. They suggest that the aggregate minimum wage employment effects estimated in literature may mask heterogeneity across different levels of labor market concentration. Munguía Corella (2020) constructed a Herfindahl-Hirschman Index (HHI) that measures the concentration of industrial employment in the US at the county level and estimated the effect for different levels of the bindingness of the minimum wage. He found negative and significant elasticity of youth employment due to minimum wage changes under perfect competition, and positive, but insignificant, effects under full monopsonistic labor markets (Munguía Corella, 2020). Moreover, the effect on employment was found to increase with the level of bindingness of the minimum wage.

In addition to the labor market structure, product market structure also matters in the employment effects of minimum wage. Harasztosi and Linder (2019) found that the reaction of employment varies across countries and industries, and that unemployment effects were greater in industries that had more difficulty passing wage costs onto consumers. Therefore, raising the minimum wage can be more costly in countries where low-wage jobs are concentrated in manufacturing (e.g., Germany) than in countries where low-wage workers are concentrated in the services sector (e.g., the US).

Using Hungary as a case study, Harasztosi and Linder (2019) confirmed that the first best option for firms as a response to minimum wage increases is to raise product prices. Similar results were obtained by Bodnár et al. (2018), they analyzed firms' reactions to minimum wage increases across Central and Eastern European countries. They found that the most popular adjustment

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channels were raising product prices, cutting non-labor costs, and improving productivity. Despite this, Poland had the highest share of firms that reported laying people off as the relevant adjustment channel. The results also indicated that firm size matters in adjustment; the layoff channel was more relevant in small firms (20–49 employees) than in firms with more workers. Similar results were found by Céspedes and Sánchez (2014), they showed employment effects monotonically decreasing in absolute terms by firm size: moderate in big firms and higher in small firms. However, Arrowsmith, Gilman, Edwards, and Ram (2003) underlined that the impact of the national minimum wage can be mediated by informality of employment relations in small firms. What is also important is that large enterprises pay higher wages than small firms (see Gibson and Stillman, 2009), so their share of workers affected by minimum wage changes is lower than in small enterprises.

Moreover, there is evidence that the size of the public sector in regional labor markets matters for wages and employment in the private sector. Nalban and Smădu (2021) showed that public job creation crowds out private sector employment, while increases in public wages lead to muted spillover effects. Alfonso and Gomes (2014) showed that growth in public sector wages and employment positively affects the growth in private-sector wages. In contrast, the International Labor Organization underlined that changes in minimum wage can have far-reaching effects on wages in the public sector, especially when different groups of workers are paid a multiple of the minimum wage, increasing the public sector wage bill.[2] Lemos (2004) explained that minimum wage increases can have different effects on employment in the private and public sectors. In the private sector, the effects are predicted by standard neoclassical theory and rely on a profit-maximizing firm, while a government employer can cover the increased wage bill by raising taxes or reducing expenditures. Lemos (2004) also noted that if the public sector has inelastic labor demands, the associated non-negative employment effect might offset some of the negative employment effects observed in the private sector, making the overall employment effect less adverse. She estimated the effects of the minimum wage on wages and employment in both private and public sectors. Adverse employment effects were found in the private sector, but no evidence of adverse employment effects was uncovered in the public sector. Navarro and Tejada (2022) recently confirmed the findings using data from Chile. They found that the institutional features of public sector employment reduce labor market frictions and mitigate the negative effect of the minimum wage on unemployment and welfare.

The differences in personal and firm characteristics translate to differences in the distribution of low-wage workers across regions, as well as the differences in the employment response to minimum wage changes at the regional level. Autor, Manning, and Smith (2016) confirmed that changes in minimum wages may have different impacts across regions and their effect on employment can induce heterogeneous responses. Williams (1993) found that elasticity of employment due to minimum wage changes in the US is highly heterogeneous among states, with the lowest (more negative) elasticity observed in the least developed regions. Thompson (2009) confirmed differences in employment elasticity for minimum wages across US counties. Ahlfeldt, Roth, and Seidel (2018) and vom Berge and Frings (2020) found that the minimum wage caused a contraction in employment growth in eastern Germany with a relatively high bite, while the west of the country experienced no change in employment. Significant differences in employment or unemployment elasticity across regions due to minimum wage increases were also found in the literature for Poland (Majchrowska and Żółkiewski, 2012; Broniatowska, Majchrowska, and Żółkiewski, 2015; Majchrowska, Broniatowska, and Żółkiewski, 2016; Albinowski and Lewandowski, 2020).

2 https://www.ilo.org/wcmsp5/groups/public/---ed_protect/---protrav/---travail/documents/genericdocument/wcms_474533.pdf

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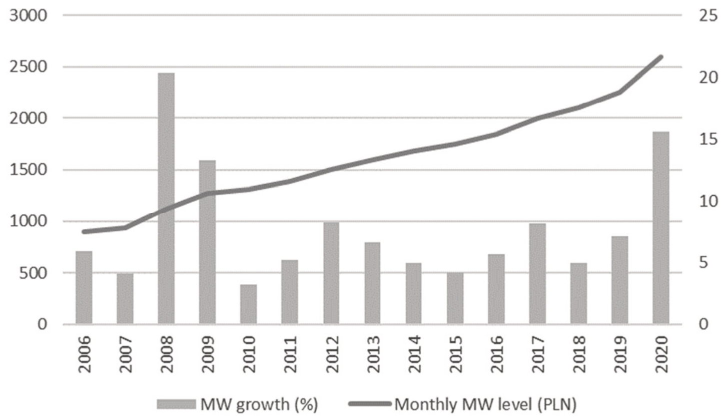
3. MINIMUM WAGE POLICY IN POLAND

The national minimum wage in Poland is regulated by law. The monthly gross minimum wage level is established every year through negotiations within the Social Dialogue Council, composed of representatives chosen from the government, employer organizations, and trade unions. If the Council is unable to reach a consensus, the minimum wage level for the following calendar year is decided solely by the Council of Ministers no later than September 15th. Since 2010, the Social Dialogue Council has not reached an agreement, and each year the decision on increasing the minimum wage has been made solely by the Council of Ministers.

The minimum wage in Poland is established at the national level; it is not differentiated by region, sector, or occupation. There is also no subminimum wage rate for younger workers. The minimum wage legislation does not cover several public sector services (teachers, health, and military services), where wages are determined by separate regulations.

The annual minimum wage increase is guaranteed to at least match the increase in price levels (CPI) projected for the following year. Additionally, in 2005, the Polish government introduced an additional rule for the minimum wage increase, reflecting two-thirds of the forecasted GDP growth rate. This rule is set until the minimum wage reaches half of the average wage in the national economy (Minimum Wage Act of October 10th, 2002, with changes). Minimum wage growth was around 7–8% on average between 2006 and 2020 (see Figure 1), and usually, the actual annual minimum wage growth exceeded the minimum value required by law.

Figure 1
Minimum wage level (PLN, left axis) and minimum wage growth (y/y, %, right axis) in Poland, 2006–2020



Source: Eurostat and the Statistics Poland.

After joining the EU in 2004, the minimum-to-average wage ratio in Poland remained around 35%. The permanent increase in minimum wage observed in the analyzed period led to an increase in the minimum-to-average wage ratio up to 50% in 2020 (Eurostat). In the same time, the ratio of the national minimum wage to average wages differs across age groups, sectors, and regions. In the 18–29 age group, this ratio exceeds 60% in industry and 70% in market services in some regions (see Figure 2).

The permanent growth of the minimum wage level also led to an increase in the share of minimum wage workers in Poland. In 2006, they accounted for 2.5% of all workers in Poland employed in firms with at least ten workers; the proportion reached 7.8% in 2020 (Table 1). Importantly, almost all minimum wage workers in Poland are employed in the private sector, while the share of minimum wage workers in the public sector is negligible. In 2020, more than 11% of private sector workers[3] received no more than minimum wage. Thus, analysis of the effect of minimum wage changes in Poland on employment in private sector is of particular importance.

[3] Employed in firms with at least 10 employees.

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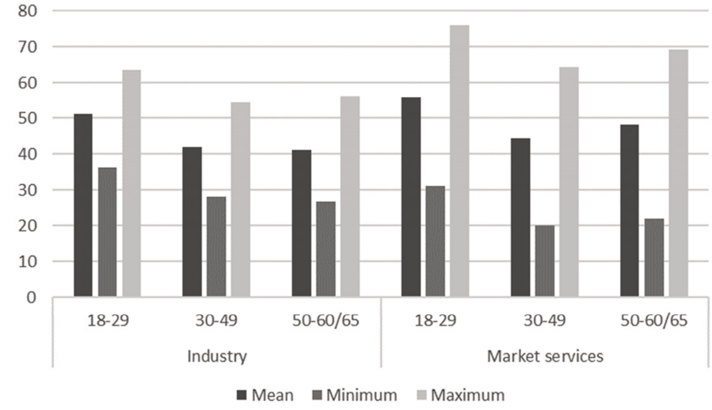
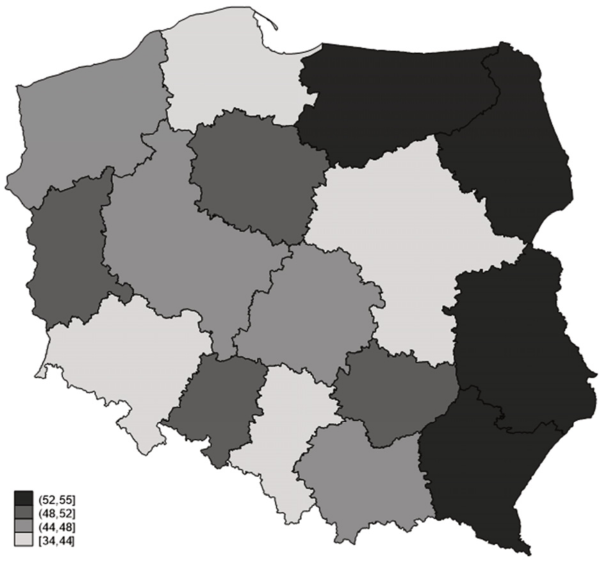
Figure 2
Mean of minimum-to-average wage ratio across age groups and sectors in Poland in 2006–2020 (%)



Source: Eurostat and Statistics Poland.

Figure 3
Mean of minimum-to-average wage ratio across NUTS-2 regions in Poland in 2006–2020 (%)



Source: Eurostat and Statistics Poland.

Table 1
Share of minimum wage workers and workers receiving more than minimum wage but less than 50% of the average wage in Poland, 2006–2020* (%)

	Share (%) of workers receiving:					
	No more than the minimum wage			More than the minimum wage but less than 50% of the average wage		
	Total	Public	Private	Total	Public	Private
2006	2.5	0.1	4.2	17.4	6.6	24.7
2008	4.2	0.1	6.7	14.3	7.2	18.7
2010	5.0	0.1	8.4	12.8	5.5	17.8
2012	7.6	0.4	11.6	11.3	6.4	14.0
2014	8.6	0.5	12.7	10.4	5.1	13.1
2016	9.0	0.6	12.8	8.5	4.2	10.5
2018	7.6	0.4	10.8	8.6	4.3	10.5
2020	7.8	0.3	11.2	5.6	1.5	7.3

* Data related only to firms with at least ten workers. Data on the share of workers earning minimum wage or more are collected biennially.
Source: Structure of Earnings Survey, different editions from 2006–2020.

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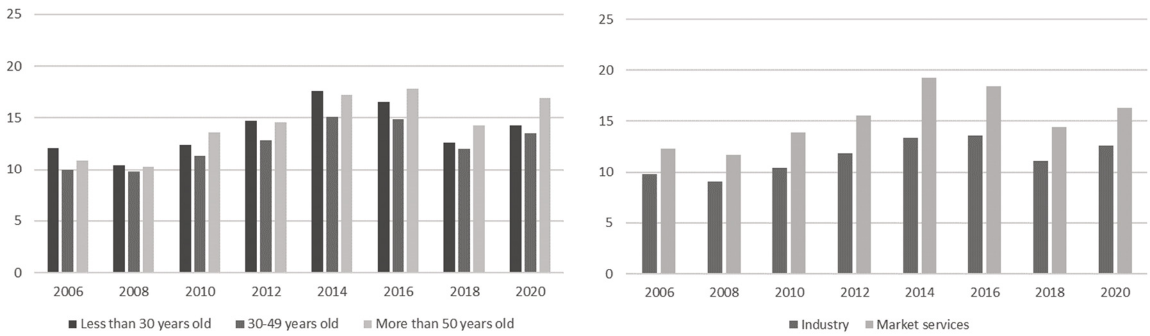
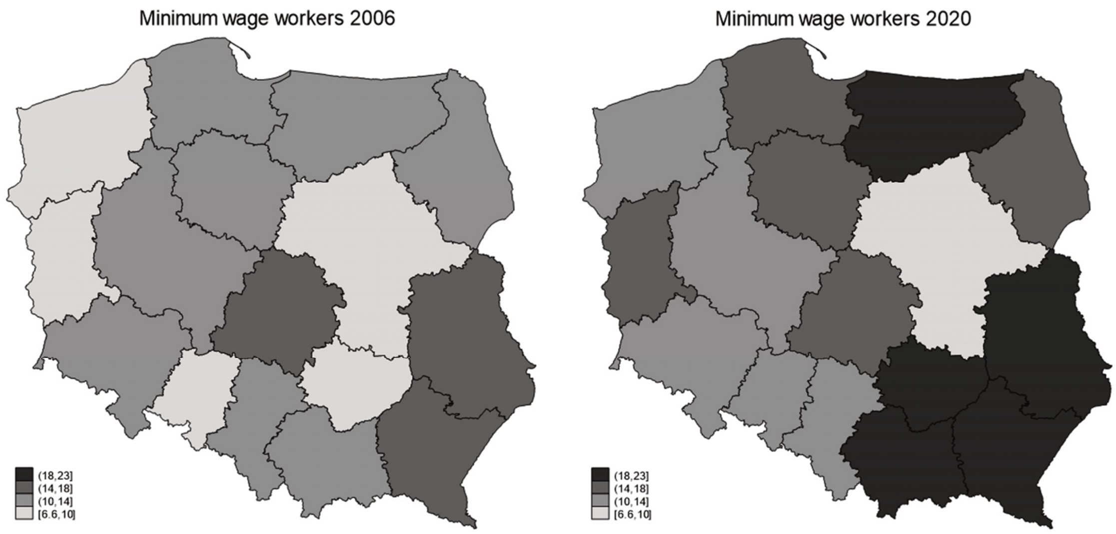
Looking at the distribution of minimum wage workers across regions, age groups, and economic sectors we can notice that both their between-regions and within-region variance is high (see Figure 4 and 5). In contrast to fully developed economies, minimum wage workers in Poland are not concentrated only among young workers; they are in all other age groups. Moreover, in all age groups, the share of minimum-wage workers increased. The share of minimum wage workers is higher in market services than in the industry. Regional differences are significant. In less developed eastern regions of Poland, the share of minimum wage workers reaches or even exceeds 20%. In the Mazowieckie (capital) region, it is below 10% (see Figure 5).

Figure 4
Share of minimum wage workers* in Poland across age groups and economic sectors in 2006–2020 (%)



* According to Eurostat, minimum wage workers are those earning not more than 105% of the minimum wage in a given year.
Source: Structure of Earnings Survey, various editions.

Figure 5
Share of minimum wage workers* in Poland across NUTS-2 regions in 2006–2020 (%)



* According to Eurostat, minimum wage workers are those earning not more than 105% of the minimum wage in a given year.
Source: Structure of Earnings Survey, various editions.

4. DATA AND EMPIRICAL APPROACH

4.1. Data

To identify how the minimum wage affects employment across subgroups of workers, we need comprehensive and reliable wage data on the eligible population and their employment level; thus, we use individual data on wages and employment characteristics from the Structure of Earnings Survey (SES) in Poland. This is part of the large European-wide survey coordinated

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by Eurostat. The SES is a large enterprise sample survey that provides detailed and comparable information on the relationships between remuneration and individual worker characteristics (gender, age, occupation, work experience, and the highest educational level attained, among others) and those of their employers (economic activity, ownership sector, NACE section, size, and enterprise location). The SES covers around 12–15% of all enterprises that employ more than nine workers. We select data from 2006 to 2020; as the SES is conducted biennially, we have eight periods.

The advantages of the database include its high reliability (wages are reported by the accounting departments of the enterprises) and scope. Each sample is very large: over 660,000 observations in 2006 and over 760,000 in 2020. Although the database represents only entities employing more than nine workers, the employment structure in Poland has a very high share of self-employed individuals operating without job contracts (own-account workers). We estimate that the SES database covered 84% of all contract workers in Poland in 2020.[4]

We made adjustments to the initial database. We focused on workers for whom the minimum wage is binding; we excluded workers younger than 18 and workers over retirement age (60 for women and 65 for men) from the initial sample. We included only private sector workers because many public sector workers are not covered by the minimum wage legislation (see section 3 and Table 2). We concentrated on workers in the industry and market services sector, including both full-time and part-time workers; we recalculated the wages of part-time workers as full-time equivalents.

The SES database is our data source for the number of employed workers and their average wages. The other data included in the model (regional and sectoral gross value added, population by age group, and unemployment rate) is based on the 16 regions according to the NUTS2 level of regional classification. They are taken from the Local Data Bank of Statistics Poland, Poland's largest publicly available database on the economy, society, and environment.

4.2. Modelling approach

The SES database provided information about monthly salaries and individual worker characteristics. Since the survey sample is randomly drawn in every reporting period, it is impossible to create a panel of individuals, although it is possible to create different sub-groups of workers, e.g., by age group, economic sector, and region. We constructed three-dimensional cells separately for each year comprising three age groups (up to 30 years, 30–50 years, and 50 years and older), two economic sectors (industry and market services), and 16 regions at the NUTS2 level. We cannot construct finer groups due to the low number of observations in some cells. The cells are our unit of analysis.

We followed the standard approach proposed in the literature and estimated the parameters of the log-linear relationship between employment, our minimum wage measure, and other variables. Following Dickens, Machin, and Manning's (1999) theoretical model, we included both demand and supply-side variables in the model. We used gross value added in economic sectors and regions as a measure of regional and sectoral demand shocks. We also included country time effects[5] to control for aggregate demand shocks. Population size approximates supply shocks. Unemployment rate controls for the size of the labor force available in the regional labor market. To consider differences in the market concentration in regional labor markets, the Herfindahl-Hirschman index (HHI) calculated for 2-digit occupational groups at each cell was included (Munguía Corella, 2020).

4 According to the data from Statistics Poland, only 34% of workers in micro firms in 2016 were employed on a job contract. Source: https://stat.gov.pl/obszary-tematyczne/podmioty-gospodarcze-wyniki-finansowe/przedsiebiorstwa-niefinansowe/dzialalnosc-gospodarcza-przedsiebiorstw-o-liczbie-pracujacych-do-9-osob-w-2016-roku,1,11.html (in Polish).

5 For the robustness check we estimated also model with regional trends included. The results are similar and available upon request.

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The 4-dimensional panel data model used in our analyses is expressed as follows:

$$empl_{R,N,A,T} = \alpha_0 + \alpha_1 wrel_{R,N,A,T} + \alpha_2 gva_{R,N,T-1} + \alpha_3 pop_{R,A,T} + \alpha_4 urate_{R,T} + \\ + \alpha_5 HHI_{R,N,A,T} + \sum \delta_{R,N,A} + \sum T_T + \varepsilon_{R,N,A,T}, \quad (1)$$

where:

$empl_{R,N,A,T}$[6] – indicates the logarithm of the number of workers employed in region R ($R = 1, 2, \ldots, 16$), economic sector N ($N = 1$ – industry, 2 – market services), age group A ($A = 1$: less than 30 years, 2: 30–49, 3: 50 and above) in year T ($T = 2006$, 2008, 2010, 2012, 2014, 2016, 2018, 2020);

$wrel_{R,N,A,T}$ – represents the logarithm of the relative minimum wage (minimum-to-average wage ratio) in region R, economic sector N, age group A, at time T;

$gva_{R,N,T-1}$ – indicates the logarithm of gross value added in region R, economic sector N, at time $T-1$ (millions of PLN, constant 2010 prices);

$pop_{R,A,T}$ – denotes the logarithm of the population in region R, age group A, at time T (thousands of people);

$urate_{R,T}$ – is the logarithm of the unemployment rate of male workers of working age in region R, at time T (%);

$HHI_{R,N,A,T}$ – is the standardized Herfindahl-Hirschman index calculated at the 2-digit occupational groups in region R, economic sector N, age group A, at time T;

$\delta_{R,N,A}$ – is the cell specific effect;

T_T – is country time effects;

$\varepsilon_{R,N,A,T}$ – represents the error term.

As a measure of employment, we took the number of workers in a given cell – those employed in enterprises with at least ten workers in the private sector in Poland. Following Caliendo et al. (2018), we used the log employment level, not the employment-to-population ratio, because the latter reflects changes in both employment level and population. We included the population at the cell level as a control variable.

Our minimum wage bite measure is the simplified Kaitz index – the relative minimum wage calculated as the ratio of minimum wage in a given year to the average wage in the previous year for a given cell. We used the difference between the log of the nominal minimum wage level applicable in a given year and the log of nominal average wages in the previous year[7] in a given cell. Since the minimum wage is unique to all workers, the variation in the minimum wage bite measure comes from minimum wage differences over time and the differences in average wages across cells over time.

Our model used the values of current minimum wage bite variables divided by the average wage lagged by one year. In Poland, information on the minimum wage increase for the next year is available usually in September of the previous year (see section 3). By lagging the average wage, we consider that entrepreneurs need time to adjust their firms' policies to upcoming changes in labor costs.

We used gross value added in a given economic sector and region as a measure of demand shock, it can affect employment. It is measured at 2010 constant prices and lagged one period, i.e., two years, to avoid simultaneity problems – a recently increased minimum wage may influence both employment and production levels. Production can be modeled as a persistent stochastic process, and changes in the current minimum wage level do not affect production levels in the previous periods. We used the regional male unemployment rate to approximate the existing

[6] We use small letters for the variables in logarithms, and capital letters for the variables in real values.

[7] Average wages are calculated as the mean of the monthly wage of individuals in a given cell without bonuses.

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surplus of the available labor force. The unemployment rate in the group of men of working age is perceived as more vulnerable to changes in the business cycle (see An et al., 2022).

Following literature, we added measures of supply shocks that affect employment. In particular, information regarding population size in a given age group and region is used to capture the idiosyncratic differences among regions. Population is measured in thousands of inhabitants.[8] The HHI measures the market concentration at the 2-digit level of classification of occupations at every cell defined by region, age group and sector. Descriptive statistics of all the variables used in the model across the cells are presented in Table A1 in the Appendix.[9]

Our main parameter of interest in model (1) is α_1, it shows the direction and strength of the relationship between the minimum wage bite and employment. Our identification strategy is based on Card's (1992) observation that "a rise in local (state) minimum wage will typically affect a larger fraction of workers in some regions (states) than in others". The induced variation creates a simple natural experiment for measuring the effect of a minimum wage change. Intensity of how wages need to change under a new minimum wage should be related to the fraction of workers initially earning less than the new minimum wage (Caliendo et al., 2018). Specifically, intensity with which wages need to change following minimum wage changes is heterogeneous among regions, age groups, and economic sectors. In the cells where the minimum wage *bites* the hardest, adaptations in wages will be stronger, as will those in labor demand.

To test it empirically, we first estimated parameters of equation (1) for the full sample to obtain an average value of the parameter of interest. We assumed homogeneity of the employment elasticity concerning the minimum wage variable across cells, used as the units of observation. However, both theoretical considerations and previous empirical results emphasize that minimum wage increases affect different groups of workers to different extents.

Thus, our second step was to test the slope homogeneity of the coefficient of the minimum wage bite measures across cells using Bersvendsen and Ditzen' (2020) Stata procedure. This method makes it possible to verify slope homogeneity in a panel data context with no correlation (Pesaran and Yamagata, 2008) or use the heteroscedasticity and serial correlation version (Blomquist and Westerlund, 2013), as employed due to the differences in our cell sizes. The influence of control variables such as gross domestic product, population, and the unemployment rate is held constant. We started with 4-dimensional cells to reduce dimensionality if homogeneity were rejected. As the test requires a panel setting, we were unable to eliminate the time dimension.

In the third step, we relaxed the assumption that employment elasticity of the minimum wage variable is homogeneous and allowed parameter α_1 in model (1) to vary first, separately across age groups, sectors, and regions and second, simultaneously across all dimensions. To choose the model that best fits the empirical data, we tested several specifications.[10] We started from the ordinary least squares, tested the presence of fixed and random effect, and finally a generalized least squares (GLS) technique that enables a heterogeneous error structure and panel-specific AR1 autocorrelation was used to correct for heteroscedasticity arising from aggregation and potential autocorrelation. We did not weight the units of observations in the model, and treated each cell as a separate observation since we were interested in estimating employment elasticity separately for each cell and comparing them with each other.

[8] We used yearly average for population and biennial data for the working population so that the data is not influenced by temporary migrations or seasonal work.

[9] Studies on minimum wage impact on employment often include a measure of other institutional variables, such as unemployment benefits, which may impact individuals' employment decisions. Majchrowska and Strawiński (2021) analyzed the impact of unemployment benefits on employment in local labor markets in Poland. They showed that social security benefits do not affect employment decisions there. The replacement ratio of unemployment benefits to minimum wage in Poland is low (41% in 2020), much lower than in Germany (78%) or France (65%; OECD data).

[10] We do not present all estimation results in the text due to limited space; all results are available upon request.

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In the fourth step, we performed a cluster analysis to find out which factors may explain differences in employment elasticity across cells. Then we expanded our model to account for those factors which explain differences in employment elasticity to highest extent.

5. RESULTS

5.1. Employment elasticity across various groups of workers

We first estimated the parameters of model (1) for the sample of private-sector workers grouped in cells. The sample included workers from all 16 NUTS-2 regions, three age groups, and two economic sectors.

Table 2
Results of model (1) with average elasticity of employment in the sample (a) and allowing employment elasticity to vary across age groups (b) and economic sectors (c)

	(a)	(b)	(c)
wrel	0.267*** (0.077)		
*wrel*age1829*		–0.581*** (0.142)	
*wrel*age3049*		–0.549*** (0.135)	
*wrel*age50plus*		0.348*** (0.073)	
*wrel*industry*			0.072 (0.093)
*wrel*market services*			0.375*** (0.082)
Lagged gross value added	0.391*** (0.081)	0.285*** (0.075)	0.511*** (0.084)
Population	0.903*** (0.047)	0.874*** (0.068)	0.891*** (0.045)
Unemployment rate	–0.063** (0.025)	–0.057** (0.023)	–0.053** (0.024)
Herfindahl-Hirschman index	0.032*** (0.007)	0.034*** (0.007)	0.030*** (0.007)
Constant	–4.914*** (1.111)	–0.581*** (0.142)	0.072 (0.093)
N	672	672	672
Cell specific effects	Yes	Yes	Yes
Country time effects	Yes	Yes	Yes

Note: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Note: *wrel*age1829* – minimum-to-average wage ratio in the 18–29 age group old; *wrel*age3049* – minimum-to-average wage ratio in the 30–40 age group; *wrel*age50plus* – minimum-to-average wage ratio in the 50–59/64 age group; *wrel*industry* – minimum-to-average wage ratio in the industry sector; *wrel*market services* – minimum-to-average wage ratio in the market services sector.
Source: Own calculations.

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Table 2 presents the estimation results. Column (a) in Table 2 presents the average values of parameters for the analyzed sample. The parameter by the gross value added variable is significant at the 1% significance level and positive. The value of 0.4 indicates that an increase in GVA by 1% was on average accompanied by an increase in total employment by 0.4%, on average. Employment is also positively correlated with population and workers' concentration measure (HHI for occupational groups). The latter shows that higher market concentration comes with higher employment. We also found a negative correlation between the regional *unemployment rate* and the level of employment. All the results are in line with economic theory and other research findings.

Our main parameter of interest (minimum wage employment elasticity) equals 0.27 and is significant at the 1% significance level. The positive sign indicates that, on average, in the analyzed period, a higher minimum-to-average wage ratio was accompanied by higher employment. The positive sign may be because the sample is based on information from all workers: those for whom the minimum wage is binding and those for whom it is not. Literature shows that negative and significant values of minimum wage employment elasticity apply only to the most vulnerable groups of workers (i.e., the young and less educated).

Our model estimated the average employment elasticity affected by minimum wage changes for the full sample, indicating that we assumed homogeneity of employment effects across age groups, economic sectors, and regions (cells) in time – the assumption is not necessarily valid. Therefore, we perform the Blomquivst and Westerlund's (2013) homogeneity test using Bersvendsen and Ditzen's (2020) Stata procedure. Results are summarized in Table A2. They indicate that when observations are divided into 4-dimensional cells, substantial differences in the impact of minimum-to-average wage on employment are observed. A different picture arises for the 3-dimensional cells. The most significant factor that causes diversity of employment effects is regional variation in industry composition and, to a lesser extent, age structure. When regional variation is completely removed from the model, the impact of relative minimum wage on employment remains identical in each cell defined by age group, economic sector, and time. In the model with cells defined by regions and time, homogeneity of the employment effect is not rejected. It implies that the interaction of industry composition and local characteristics is likely to be responsible for the heterogeneous reaction of employment to changes in the minimum wage.

Therefore, in the third step, we relaxed the assumption of homogeneity of employment elasticity for the minimum wage variable and allowed the parameter by the minimum wage variable to vary separately across regions, age groups, and economic sectors. Column (b) in Table 2 presents results for age groups. The parameter by the minimum wage variable is significant for all age groups. The sign of the parameter is negative for young and middle-aged workers and positive for workers aged 50+. The results suggest that if firms dismiss workers, they reduce employment among those who are least costly, i.e., young and less experienced workers. The results indicate that employers do not dismiss experienced older workers since their layoff costs are higher. Subsequently, we allowed the parameter by the minimum wage variable to vary across the two economic sectors (see Column (c) in Table 2). The parameter estimate by the minimum wage variable is insignificant for the industry sector but significant and positive for market services. Lastly, we allowed the parameter by the minimum wage variable to vary across 16 NUTS2 regions. In most regions, the parameter is significant, but interestingly, the sign of the parameter estimate differs; in two regions, it is negative, and in eight, it is positive (see Table 3). The results show that the reaction of employment to minimum wage changes is diversified across age groups, economic sectors, and regions.

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Table 3
Results of model (1) allowing employment elasticity to vary across 16 NUTS2 regions

	Estimated parameters	Standard errors
*wrel*dolnoslaskie*	0.357*	(0.193)
*wrel*kujawsko-pomorskie*	–0.382**	(0.169)
*wrel*lubelskie*	0.340**	(0.160)
*wrel*lubuskie*	–0.385*	(0.224)
*wrel*lodzkie*	1.197***	(0.270)
*wrel*malopolskie*	0.506***	(0.128)
*wrel*mazowieckie*	0.725***	(0.123)
*wrel*opolskie*	0.717***	(0.187)
*wrel*podkarpackie*	0.781***	(0.203)
*wrel*podlaskie*	0.239	(0.179)
*wrel*pomorskie*	0.212*	(0.110)
*wrel*slaskie*	0.462***	(0.102)
*wrel*swietokrzyskie*	–0.291*	(0.156)
*wrel*warminsko-mazurskie*	–0.811***	(0.186)
*wrel*wielkopolskie*	0.840***	(0.138)
*wrel*zachodniopomorskie*	0.121	(0.157)
Lagged gross value added	0.141	(0.097)
Population	0.902***	(0.045)
Unemployment rate	–0.080***	(0.028)
Herfindahl-Hirschman index	0.027***	(0.007)
Constant	–2.088*	(1.266)
N	672	672
Cell specific effects	No	No
Country time effects	No	Yes

Note: * $p<0.1$, ** $p<0.05$, *** $p<0.01$, *wrel*name of the region* – minimum-to-average wage ratio in the given NUTS2 region.
Source: own calculations.

Therefore, in the next step, we extended the analysis and allowed the parameter by the minimum wage variable to vary across age groups, economic sectors, and regions simultaneously:

$$empl_{R,N,A,T} = \beta_0 + \beta_{R,N,A} wrel_{R,N,A,T} + \beta_2 gva_{R,N,T-1} + \beta_3 pop_{R,A,T} + \beta_4 ur_{R,T} + \\ + \beta_5 HHI_{R,N,A,T} + \sum \gamma_{R,N,A} + \sum T_T + \epsilon_{R,N,A,T}. \quad (2)$$

Due to the relatively small number of observations in time,[11] we estimated the average employment elasticity for each cell. Figure 6 presents results for the group of young workers. We can observe negative elasticity of employment for young workers in the industry sector in 11 out of 16 Polish regions. In the other five regions the impact of minimum wage changes

[11] SES is conducted every two years. The research period covers 2006–2020, and we have lagged gross value added in the model; therefore, the number of periods is reduced from eight to seven.

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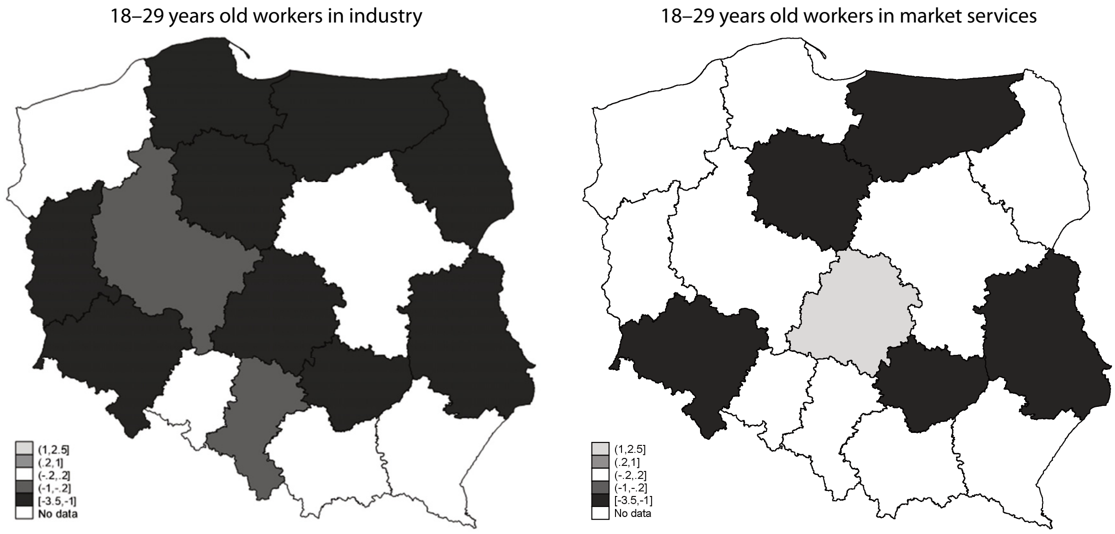
on youth employment was insignificant. The opposite situation was noted in market services. A negative employment reaction among young workers was found only in five regions. In the others the relationship was insignificant.

We observe similar picture in the case of workers aged 30–49. In most of the regions increased minimum-to-average ratio was accompanied by decreased employment of 30–49 years old workers in the industry sector. Similarly as in the case of young, the employment reaction was less pronounced in the market services (see Figure 7).

Figure 6

Elasticity of employment with respect to minimum wage changes for the group of workers aged 18–29 across sectors and regions in Poland (on average, 2006–2020)



Note: Dark colors indicate regions with negative employment elasticity across given age groups and sectors (black: employment elasticity lower than –1; dark grey: employment elasticity between –1 and 0.2). Light colors indicate regional labor markets with positive employment responses (light grey: employment elasticity between 0.2 and 1; medium grey: employment elasticity higher than 1). Areas with insignificant employment effects are in white.

Source: Author's calculations.

Figure 7

Elasticity of employment with respect to minimum wage changes for the group of workers aged 30–49 across sectors and regions in Poland (on average, 2006–2020)

Note: Dark colors indicate regions with negative employment elasticity across given age groups and sectors (black: employment elasticity lower than –1; dark grey: employment elasticity between –1 and 0.2). Light colors indicate regional labor markets with positive employment responses (light grey: employment elasticity between 0.2 and 1; medium grey: employment elasticity higher than 1). Areas with insignificant employment effects are in white.

Source: Author's calculations.

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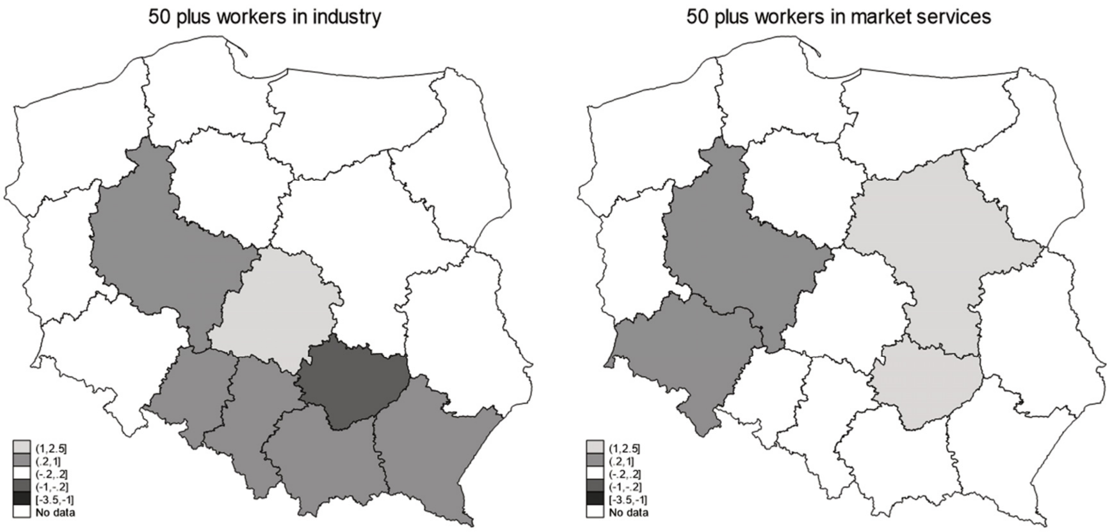
Finally, we observe a completely different picture in the case of workers aged 50 and more. A negative employment reaction in this group was observed only in one, less-developed region. Contrary, in six regions we observe a growth of employment of workers aged 50 and more in the industry sector and in four regions – in the market services. In most of the regions the minimum wage growth did not affect employment among the 50 plus workers either in industry or in market services (see Figure 8).

Figure 8
Elasticity of employment with respect to minimum wage changes for the group of workers aged 30–49 across sectors and regions in Poland (on average, 2006–2020)



Note: Dark colors indicate regions with negative employment elasticity across given age groups and sectors (black: employment elasticity lower than –1; dark grey: employment elasticity between –1 and 0.2). Light colors indicate regional labor markets with positive employment responses (light grey: employment elasticity between 0.2 and 1; medium grey: employment elasticity higher than 1). Areas with insignificant employment effects are in white.

Source: Author's calculations.

The results are in line with theoretical predictions. More negative employment effects are observed in the industry, firms are more exposed to international competition and cannot increase product prices. To maintain profits, they reduce employment among those who are the least costly, i.e., less experienced and less educated workers. We observe the most negative employment effects in underdeveloped regions of Poland, confirming the findings of Majchrowska (2022), who found higher minimum wage pass-through effects on prices in richer, highly developed regions of Poland. For robustness, we estimated cell employment effects in a model with regional trends instead of common time effects included; the results confirm the main findings.[12]

Noteworthy in our approach is that we can see the existing heterogeneity within regions. The differences are undetectable in one- or two-dimensional approaches. Prior studies indicated that, in some regions, regional employment effects were insignificant. Our approach finds that statistically insignificant values of employment elasticity at the regional level very often mask diverse employment effects within the region: across age groups and economic sectors.

5.2. Determinants of differences in employment elasticity

In this part of the study, we aim to ascertain why employment elasticities for groups of workers defined by age and sector differ strongly among regions. In particular, we want to find out why we report negative employment elasticities for young or middle-aged workers in some regions but not in others.

[12] Available upon request.

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To explain the differences in employment elasticity among cells, we performed a k-medoid cluster analysis using the Manhattan distance. We clustered three-element vectors of employment elasticity, the share of workers in manufacturing, and the share of workers employed in the public sector. We chose the best solution according to the Caliński and Harabasz criterion.[13]

Following prior empirical findings, unemployment effects are expected to be more pronounced in industries where it is difficult to pass higher wage costs on to consumers. Therefore, in cells with a larger share of workers employed in the tradable sector, approximated in our study by manufacturing, employment elasticity should also be negative. We expected employment elasticity to be positively correlated with the proportion of workers in the public sector; in less competitive environments – in cells with a higher proportion of workers in the public sector – employment elasticity should be lower than in cells with more private-sector workers.

The cluster analysis results indicate that we should choose the solution with seven clusters. The first comprises two cells with positive employment elasticity, i.e., the cells with middle-aged workers and workers aged 50+ employed in market services in the capital region (Mazowieckie). Another cluster with positive employment elasticity comprises cells mostly of workers aged 50+ in other regions with big agglomerations (Wroclaw, Krakow, Katowice). The cells are characterized by a high share of low-educated workers but also a relatively low share of public sector workers, a low share of workers in manufacturing, and a high share of workers employed in firms with 250 and more employees.

There are also two clusters with strong negative employment elasticity. One group mostly comprises cells for young workers in industry and market services, mostly in underdeveloped regions. The second consists mostly of cells for middle-aged workers in both industry and market services, again mostly in underdeveloped regions. The latter cells are characterized by a high share of employment in manufacturing, a low share of employment in the public sector, and a high share of workers with a low level of education. Unfortunately, the remaining clusters have no clear interpretation.

In the last step of our analysis, we enlarged model (1) by incorporating the labor market characteristics that describe variation in employment elasticity to minimum wage changes. We interacted the relative minimum wage with the share of public sector workers in a given cell. We expected the interaction coefficient to be positive, indicating that elasticity is not as adverse when more public firms are present. If the coefficient of the interaction term is not significant, it may also indicate no heterogeneity in employment elasticity between the public and private sectors. Additionally, we interacted the relative minimum wage with the share of those working in manufacturing and expected the interaction coefficient to be negative. A significant share of firms in the manufacturing sector is exposed to international competition, and if there are increased labor costs, they cannot pass them on to consumers, so they decide to lower their employment.

Table 3 shows results of model (1) with interaction terms included. Each specification contains cell specific effects.[14] The interactions were found to be significant, which is consistent with our predictions. The higher the public sector share, the less negative the employment reaction to minimum wage changes. Conversely, employment elasticity with respect to minimum wage changes is negatively correlated with the share of people employed in manufacturing. The higher the share of workers in manufacturing, the stronger the unemployment effects predicted.

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13 The full results of the cluster analysis are available upon reasonable request.

14 Table 4 presents results of the model with country time effects. For robustness check we estimated also the model with regional trends. The results are very similar and are available upon request.

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Table 3
Results of Model (1) with interactions included

	(1a)	(2a)	(3a)
Minimum to average wage ratio	–0.857***	3.578***	2.800***
	(0.130)	(0.435)	(0.423)
Lagged gross value added	0.518***	0.306***	0.394***
	(0.066)	(0.080)	(0.069)
Population	0.535***	0.975***	0.639***
	(0.045)	(0.045)	(0.046)
Unemployment rate	–0.062***	–0.065***	–0.079***
	(0.021)	(0.023)	(0.020)
Herfindahl-Hirschman index	0.023***	0.032***	0.025***
	(0.006)	(0.007)	(0.006)
Share of public sector	0.084***		0.067**
	(0.030)		(0.030)
*Share public*wrel*	0.462***		0.421***
	(0.041)		(0.043)
Share of manufacturing		–0.267***	–0.423***
		(0.096)	(0.084)
*Share manufacturing*wrel*		–0.893***	–0.972***
		(0.120)	(0.110)
Constant	–2.131**	–3.826***	–0.427
	(0.881)	(1.193)	(0.999)
N	672	672	672
Cell specific effects	Yes	Yes	Yes
Country time effects	Yes	Yes	Yes

Note. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Source: Author's calculations.

6. CONCLUSIONS

Literature shows that negative employment effects are observed among less-skilled and less-experienced workers. Moreover, a growing number of authors underline heterogeneity of the labor markets across regions as a possible source of non-significant employment elasticity at the aggregate level. The paper analyzes which factors determine the size of employment effects with respect to minimum wage increases. We estimate employment elasticities of minimum wage increase for different segments of the labor market in Poland simultaneously and analyze factors behind them.

At first, we assumed homogeneity of employment elasticity of minimum wages. Next, using the Blomquist and Westerlund's (2013) test, we rejected slope homogeneity. We then allowed

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employment elasticity to vary across age groups, economic sectors, and regions simultaneously. It creates a novel four-dimensional approach. Using cluster analysis, we searched for similarities among estimated employment elasticities. Finally, we enlarged our model by incorporating the labor market characteristics that described the obtained clusters of workers. To the best of our knowledge, the study is the first of its kind.

We confirmed regional differences in employment elasticity due to minimum wage changes and discovered latent heterogeneities in the regional employment effects. In many regions, insignificant and close-to-zero overall results include both strongly positive and strongly negative values of employment elasticities due to minimum wage changes for different groups of workers. Finding the heterogeneities would have been impossible without implementing our four-dimensional approach.

Age and sector were found to be the most important determinants of employment elasticity diversity. Negative employment effects were observed mostly among the youngest groups of workers, while positive effects were observed mostly in the groups of workers aged 50+. Among middle-aged workers, both negative and positive reactions were observed. Employment reaction depends also on the economic sector: negative effects are observed more often in industry than in market services. Conversely, positive elasticities are more likely in market services.

We cannot confirm that negative employment effects of minimum wage increases for young workers are observed in all regions. We found out that the employment reaction to changes in the minimum wage is the result of a combination of regional labor market features. In some regions, there are highly intense features that increase probability of negative employment effects; in other regions, the opposite is true.

Negative employment effects are more likely when there is a larger proportion of workers in the private sector, where there are industries in which it is more difficult to increase prices of the goods or services produced, and where small firms are widespread. In the regions, employers act in a highly competitive environment, have more bargaining power, and the probability of unemployment is relatively high for young workers and the middle-aged, and especially for those less educated.

A positive employment effect is more probable in regions with a high share of workers in the public sector and in large enterprises. In the regions, private sector employers have less bargaining power because they have to adjust their wage policy to the public sector wages. Being employed in a big firm also diminishes the probability of dismissal, even among less-educated workers. Significantly, the two completely different labor market segments coexist within a given region, as in Poland, which explains why empirical analyses at the regional level have often resulted in insignificant values for the minimum wage parameter.

The results are important for the minimum wage research. They show that previous analyses at the aggregated (national or regional) level might underestimate employment effects of minimum wage. The small or insignificant employment elasticities obtained might be the result of significant opposing effects across different groups of workers. The multidimensional approach presented in the study enabled us to uncover internal heterogeneities.

The results are also important for minimum wage policies, as they show that minimum wage effects cannot be easily predicted by policymakers. Due to the differences in the characteristics of employers and employees in regional labor markets, the local employment effects of changes in the national minimum wage may substantially differ. Even for workers with similar personal characteristics, the employment reaction may depend on the employer's size, the economic sector, or the degree of local competition. The variety of labor market features that influence employment elasticity makes predicting total effects related to minimum wage changes very difficult.

Our results are also important for policymakers in Poland. They undermine the purposefulness of the regional differentiation of Poland's minimum wage proposal, endorsed by, among others, the OECD, which emphasized: "Consider differentiating the minimum wage across regions depending on local labor market conditions" (OECD 2018). Our results show that finding an

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optimal regional minimum wage rate would be difficult due to large intra-regional heterogeneities in labor markets.

Like most research, our study has some limitations. First, the data used includes only companies with at least ten workers. Unfortunately, individual data on micro-firms in Poland is not available. Small firms are usually found in the market services sector, where labor costs are more important than capital costs, and the firms are probably more intensively affected by minimum wage changes. It may impact results for market services, which we have ascertained can be underestimated. The second limitation stems from the ability to construct only a biannual panel; it does not let us capture very unsuccessful firms that survived for a short time. Third, there is a discrepancy in the data¾since workers are identified in the data through their firms, we inferred the location of a worker's residence as the same as that of the firm. However, workers can commute to work over long distances, and therefore, spatial interactions should be considered. As it is a very broad issue, it could be the subject of future research.

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APPENDIX

Table A1
Descriptive statistics of the variables used in the model

		N = 768 / n = 96 / T = 8			
		Mean	Standard deviation	Minimum	Maximum
Number of employed (number of workers)	overall	49,832	54,390	4,538	514,313
	between		53,000	6,846	370,172
	within		13,224	–111,795	193,973
Minimum-to-average wage ratio (%)	overall	47.1	9.5	20.1	76.0
	between		7.9	24.0	67.5
	within		5.3	28.7	63.4
Gross value added (millions of PLN constant 2010 prices)	overall	35,586	34,350	6,346	235,409
	between		33,503	9,006	184,881
	within		8,225	–11,963	86,113
Population (number of people)	overall	501,426	302,147	125,885	1,721,540
	between		300,166	168,972	1,570,209
	within		44,894	313,801	669,541
Unemployment rate (%)	overall	7.8	3.5	1.7	16.7
	between		1.3	5.5	10.3
	within		3.3	1.8	16.0
Herfindahl-Hirschman index	overall	0.09	0.02	0.05	0.21
	between		0.02	0.06	0.19
	within		0.01	0.05	0.18

Note. The Herfindahl-Hirschman index is calculated for 2-digit occupational groups.
Source: Authors' calculations.

Table A2
Results of the Blomquist and Westerlund (2013) homogeneity test

Dimensions	Number of cells	Delta	p-value
4: Region, Age, NACE, Time	16 x 3 x 2 = 96	4.296	0.000
3: Age, NACE, Time	3 x 2 = 6	–1.323	0.186
3: Region, NACE, Time	16 x 2 = 32	2.274	0.023
3: Region, Age, Time	16 x 3 = 48	3.668	0.000
2: Region, Time	16	0.815	0.415

Note: Null hypothesis: Slope homogeneity.
Source: own calculations.

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