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Editorial introduction

This issue of *Economic and Business Review* delves into pressing economic and social dynamics shaping innovation, inequality, and financial markets amid global uncertainties. Recent trends, such as rising R&D incentives, post-pandemic widening income disparities highlighted by OECD reports, and the surge in fintech services like Buy Now Pay Later (BNPL), underscore the need for targeted policies. Theories like behavioural finance and trust-based social cohesion frameworks further illuminate how investor sentiment, financial inclusion, and algorithmic trading influence growth and stability.

These six articles written by nineteen scholars from five countries (Bangladesh, Belgium, Poland, Thailand, Vietnam) collectively advance this discourse by blending empirical rigour with policy insights across diverse contexts—from European tax designs to emerging Asian markets. They reveal how incentives drive SME innovation, class-specific attitudes affect trust, and digital tools reshape liquidity and consumer behaviour. The following overviews detail their contributions.

The first article, **R&D tax credits, innovative activity and the targeting approach**, by Erik Gjymshana, Annelies Roggeman, and Isabelle Verleyen, investigates the impact of French research and development (R&D) tax credits on the innovative activity of small and medium enterprises (SMEs). The unique feature of this system lies in its application to expenditures incurred during the development phase of R&D projects, instead of all R&D expenditures. The authors employ a regression discontinuity design (RDD) and compare targeted SMEs with larger firms not subject to the tax credit over the period 2014–2018. The findings highlight the positive impact of the system on SMEs' innovative activity and their stronger response to incentives during the growth stage. This effectiveness is not held over time. Thus, the study provides a valuable policy lesson for other countries that are designing both their innovation policies and tax systems supporting R&D.

The second article, **Attitudes towards income inequality and trust: An analysis by income class in Poland**, written by Małgorzata Szczepaniak and Katarzyna Bentkowska, focuses on the under-researched relationships between attitudes toward income inequality and generalised and institutional trust in Poland, thereby contributing to the discussion on social cohesion. The authors apply an economic stratification framework with five income classes, complemented by non-parametric tests and logistic regression (for $N = 1{,}352$). Acceptance of inequality increases with income, with the sharpest contrasts between low- and high-income classes. Generalised trust rises with income, but institutional trust follows more complex and non-linear patterns. Because the links between trust and inequality attitudes are class-specific, the article argues that efforts to reduce inequality should address not only material disparities but also public perceptions of them.

The third article, **Financial inclusion and economic growth in Vietnam: Evidence across provinces and income groups**, by Thi Thuy Huong Luong, Attasuda Lerskullawat, and Thi Anh Nhu Nguyen, highlights the effect of financial inclusion on economic growth in Vietnam. The authors used panel data from 63 provinces during 2014–2020, and researched the full sample and two income groups. The difference-GMM estimation results demonstrate that financial inclusion, as measured by the number of commercial bank branches and the use of bank accounts, saving passbooks, and ATM cards, has a significant positive effect on economic growth in Vietnam. Additionally, for high-income provinces, participating in life and non-life insurance positively affects economic growth. This study suggests that policymakers should prioritise measures to expand access to and use of financial services in Vietnam, in addition to designing targeted programmes to increase the accessibility of insurance products, with a special focus on rural and low-income regions.

The fourth article, **Perceived usefulness, ease of use, risk, and trust: Explaining BNPL user recommendation intention through behavioural models**, by Krzysztof Waliszewski, Ewa Cichowicz, Mateusz Folwarski, Łukasz Gębski, Jakub Kubiczek, Paweł Niedziółka and Małgorzata Solarz, examines the main factors that influence consumers' willingness to recommend Buy Now, Pay Later (BNPL) services to others. Based on survey data from a quota sample of 350 BNPL users, the study analyses consumer recommendation intention using Partial Least Squares Structural Equation Modelling. The findings indicate that perceived usefulness and trust in the BNPL provider have a strong positive effect on recommendation intention, whereas perceived risk exerts a significant negative influence. Overall, the results extend the existing literature by clarifying the behavioural mechanisms driving BNPL adoption, while also offering practical implications for financial service providers seeking to strengthen consumer advocacy and retention.

The subsequent article, entitled **Sentiment and dividend smoothing: Do firms alter dividends during periods of high market activity?** and written by M. Jahir Uddin Palas and M. Adnan Ahmed, explores whether investor sentiment influences dividend policy among publicly listed firms in Bangladesh. Using a balanced panel dataset of 116 firms covering the period 2010–2021, the authors apply panel regression techniques, including random effects models, panel-corrected standard errors, and instrumental variable estimation. The analysis shows that firms tend to increase dividends in response to positive investor sentiment, while simultaneously engaging in dividend smoothing to preserve stable payout signals. The paper contributes to the behavioural finance literature, and its insights are particularly useful for investors and managers of public companies.

The final paper in the issue, **Algorithmic trading, liquidity and volatility: Evidence from Poland**, authored by Henryk Gurgul and Robert Syrek, investigates the causal relationships between algorithmic trading intensity, market liquidity, and volatility for selected blue-chip companies listed on the Warsaw Stock Exchange. Using daily and high-frequency intraday data for firms permanently included in the WIG20 index over the period 2020–2023, the study applies information-theoretic methods, including Shannon and Rényi transfer entropy. The research shows that algorithmic trading significantly affects both liquidity and volatility, with stronger and more widespread causal relationships observed at higher data frequencies, while such effects are less pronounced for extreme market conditions. These findings are of particular importance for regulators and market participants seeking to better assess the implications of algorithmic trading for market stability and efficiency in European equity markets.

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R&D tax credits, innovative activity and the targeting approach

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Abstract

The aim of this study is to investigate whether the French R&D tax credit targeted at small and medium-sized enterprises (SMEs) has a positive impact on innovative activity. The French institutional setting provides a unique research framework as the R&D tax credit targeted at SMEs only applies to expenditures incurred during the development phase of R&D projects instead of all eligible R&D expenditures. In order to explore the effectiveness of the French R&D tax credit, a regression discontinuity design (RDD) is applied by comparing targeted SMEs with larger firms not subject to the tax credit over the period 2014–2018. In general, we find that the French R&D tax credit has a positive impact on innovative activity. Moreover, SMEs react more strongly to this incentive in their growth stage. The findings suggest, however, that this effectiveness in increasing SMEs' innovation does not persist over time.

Keywords

- R&D tax credit
- innovation
- targeted tax incentive
- SMEs
- firm life cycle

JEL codes: H25, H32, O31, O32

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Introduction

Governments have many measures at their disposal in order to stimulate private R&D, among which subsidies and R&D tax credits are the most commonly used (Chen & Yang, 2019; Montmartin et al., 2018). The governments of most Western nations have introduced tax incentives designed to complement subsidies (Mulkay & Mairesse, 2013). However, the complementarity between direct support, in the form of subsidies, and indirect support, in the form of various tax incentives, has only been considered in a limited number of studies (Dumont, 2017; Montmartin et al., 2018). Whereas countries such as Belgium, France and the Netherlands increasingly favour tax incentives, others such as Germany, Sweden and Switzerland up until recently only provided direct support (Appelt et al., 2019; Dumont, 2017).

While tax incentives are essentially a generic policy instrument, targeting specific groups of firms is quite common and many European countries target small and medium-sized enterprises (hereinafter, SMEs) and young firms (European Commission, 2014). The level of R&D performed by SMEs is crucial to a country's technological progress and economic growth: as Kobayashi (2014) points out, dormant R&D by Japanese SMEs contributed to the slowdown of Japan's economic growth and its lost decade. The effects of tax incentives on R&D vary across subgroups of firms, with most studies focusing on whether the impact of tax incentives is related to firm size (e.g., Baghana & Mohnen, 2009; Foreman-Peck, 2013; Kobayashi, 2014; Koga, 2003; Lokshin & Mohnen, 2012).

R&D tax credits are a major public policy instrument geared towards increasing private firms' incentives to invest in R&D activities. The growth in the literature is partly due to the increasing popularity of tax credits, which are adopted in more than 20 OECD countries (Castellacci & Lie, 2015). According to the meta-regression analysis of Castellacci and Lie (2015), the additionality effect of R&D tax credits, i.e. the rate at which R&D investments increase due to the introduction of tax incentives is on average stronger for SMEs. Moreover, the empirical literature investigates the effects of R&D tax credits targeted at SMEs in France, the United Kingdom and Japan (e.g., Bunel & Hadjibeyli, 2021; Dechezleprêtre et al., 2023; Foreman-Peck, 2013; Kobayashi, 2014).

The French R&D tax credit targeted at SMEs, also referred to as the innovation tax credit, offers interesting research opportunities for several reasons. Firstly, in France, SMEs account for 86% of R&D tax relief recipients, while the share of R&D tax support accounted for by SMEs only amounts to 30%. This indicates that larger firms, which represent the remaining 14% of R&D tax relief recipients, account for 70% of R&D tax support (Appelt et al., 2019). Secondly, France offers a policy mix of public support to business R&D, providing R&D tax relief through its R&D tax credits and its regime for young innovative companies, which was introduced back in 2004. This makes France one of the most generous OECD countries in terms of R&D tax incentives (Appelt et al., 2019). Thirdly, the French institutional setting offers a unique research framework as the French R&D tax credit targeted at SMEs only applies to expenditures incurred during the development phase of R&D projects. The R&D tax credits targeted at SMEs in the United Kingdom, Canada and Japan do not share this characteristic. Providing additional tax advantages may be an effective tool to induce SMEs to conduct more R&D, as existing studies find that many SMEs face financial constraints and have limited access to external funding, which in turn hinders their R&D (Kobayashi, 2014).

The aim of our research is to investigate whether the French R&D tax credit targeted at SMEs was successful at promoting R&D for French SMEs. We explore this by comparing French SMEs with larger companies which do not have access to this specific R&D tax credit. We contribute to literature by being, to the best of our knowledge, the first to conduct an in-depth study of the impact of the French R&D tax credit targeted at SMEs on innovation. On behalf of the French Institute of Statistics and Economic Studies, the National Bank of France and the French tax authorities, Bunel and Hadjibeyli (2021) also evaluated this French R&D credit, although they did so in terms of employment and turnover instead of R&D development as measured by the firms' investments in intangible fixed assets. Although the empirical literature has investigated the impact of targeted tax incentives for R&D in the United Kingdom, Canada and Japan, as yet no independent study exists of the French R&D tax credit targeted at SMEs. We also investigate if this impact is more pronounced for SMEs in their growth stage and in high-tech industries. Moreover, we contribute to the research methodology by applying a fuzzy regression discontinuity design and taking into account all SME criteria for determining eligibility for this specific R&D tax credit. By doing so, we reduce the likelihood that SMEs are incorrectly assigned to the control group or large firms to the treatment group. With this study, we provide further evidence regarding the effectiveness of targeted tax incentives, an aspect that might have important policy implications. In general, we find that the French R&D tax credit has a positive impact on innovative activity. Moreover, SMEs react more strongly to this incentive in their growth stage.

The remainder of our paper is structured as follows: Section 1 discusses the French R&D tax credit targeted at SMEs in greater detail. Section 2 gives an overview of the existing literature on the impact of R&D tax credits on corporate innovation and develops our hypotheses. Section 3 delves deeper into our data, research design and econometric model. Section 4 presents the results of our regression analyses and robustness tests. Finally, we state our conclusions.

1. Institutional setting: The French R&D tax credits

In late 2013, a new R&D tax credit targeted at French SMEs was introduced. Called the innovation tax credit (*Crédit d'Impôt Innovation*, CII), this would come into effect on 1 January 2014. The CII allows SMEs an extra reduction of 20% of incurred expenditures, with a ceiling of 400,000 EUR per year, thus resulting in a maximum tax credit of 80,000 EUR per year (Bozio et al., 2014). According to Art. 244 quarter B of the French Tax Code, only the expenditures incurred during the development phase of R&D projects can be taken into account for the CII. The expenditures incurred during the research phase of R&D projects can still be taken into account for the CIR (*Crédit d'Impôt Recherche*), since this R&D tax credit is applicable to all French firms. The CII can be combined with the CIR in order to offer a very generous tax treatment to French SMEs for expenditures incurred during the development phase of R&D projects.

The design and implementation of tax incentive schemes can have an important impact on the capacity of SMEs to benefit from them (Mitchell et al., 2020). In this regard, an immediate refund, which is available in the CII, is beneficial for SMEs, as they might be faced with liquidity constraints. Internal funding is important for making investments in activities with uncertain outcomes, such as R&D. If liquidity-constrained firms have any difficulty raising capital externally, R&D credits might be especially important (Kobayashi, 2014). Table 1 provides an overview of the main characteristics of the CIR and CII during the period 2008–2018.

Whereas all French firms are eligible for the CIR, only French SMEs are eligible for the CII. The definition of an SME in France is based on three firm-specific criteria: the number of employees, the total assets and the turnover. As is the case in most European countries, this definition takes into account the above-mentioned criteria from the last two accounting years. A French firm is considered to be an SME if, during the last two accounting years, its number of employees was less than 250. Moreover, its total assets or its turnover could not exceed 43,000,000 EUR or 50,000,000 EUR, respectively, during the last two accounting years.

Table 1. Overview of the French R&D tax credits (2008–2018)

Characteristics	CIR (2008–2018)	CII (2014–2018)
Target	all	only SMEs
Credit tax rate	30% for R&D exp. <= 100 mln EUR 5% for R&D exp. > 100 mln EUR	20%
Eligible expenditures	all R&D expenditures (fundamental, applied and experimental research)	only the expenditures incurred during the development phase of R&D projects
Ceiling	–	400,000 EUR
Refund	after 3 years (for large firms) or immediately (for SMEs)	immediately

Source: on the basis of (Bozio et al., 2014; Liu, 2013; Mulkay & Mairesse, 2013).

2. Literature and hypotheses

Tax incentives for R&D are one of the most popular innovation policy tools (European Commission, 2014). Over the last two decades, substantial research has been done on the effects of R&D tax credits and their impact on corporate innovation (e.g., Chen & Yang, 2019; Kobayashi, 2014; Koga, 2003; Lokshin & Mohnen, 2012). These studies' findings converge, as they find that R&D tax credits increase firms' innovative activities.

The effects of tax incentives for R&D vary across subgroups of firms, with most studies focusing on firm size (European Commission, 2014). In some countries, SMEs and liquidity constrained firms respond more positively to tax incentives (e.g., Baghana & Mohnen, 2009; Lokshin & Mohnen, 2012). In other countries, however, the opposite conclusion can be drawn (e.g., Koga, 2003). Thus, when investigating how the impact of tax incentives relates to firm size, the results differ across countries (European Commission, 2014). Nonetheless, the meta-regression analysis conducted by Castellacci and Lie (2015) indicates that the additionality effect of R&D tax credits, i.e. the rate at which R&D investments increase due to the introduction of tax incentives, is on average stronger for SMEs.

However, the studies so far mentioned in this section focused on untargeted tax incentives applicable to all firms. When considering targeted tax incentives which offer preferential regimes to any specific subgroup of firms, the results converge. R&D tax credits targeted at SMEs, which offer preferential regimes to SMEs relative to large firms, impact positively on SMEs' decisions to conduct more R&D (e.g., Agrawal et al., 2020; Dechezleprêtre et al., 2023; Kobayashi, 2014). Moreover, recent studies advocate for a more targeted approach of tax incentives for R&D in order to make them more effective (Chen & Yang, 2019; Montmartin et al., 2018).

Montmartin et al. (2018) found evidence of a neutral impact of the French tax credit system prior to the introduction of the CII, and concluded that this result was principally due to the French tax credit policy being untargeted. Furthermore, Chen & Yang (2019) recommend that governments should restrict the scope of industries and mainly allocate tax credits to innovation-driven enterprises, as their results indicate that the Chinese R&D tax credit only facilitated innovative activities in large firms. This result seems in line with the study for Japan conducted by Koga (2003). Hence, evidence suggests that tax incentives which do not target any specific subgroup of firms tend to support the larger incumbent R&D firms. Rao (2016) adds that larger firms have a weaker immediate response to tax incentives for R&D but do not go on to reduce their research spending in future years like smaller firms do.

In terms of a cost-benefit rationale, a key question in our research setting is whether additional tax advantages offered to French SMEs can offset mechanisms such as compliance burdens or liquidity constraints (Mitze & Kreutzer, 2023). As these authors point out in their research, policy enforcement and monitoring cost rise with strategic importance. Moreover, hidden implementation costs that stem from unpredicted task complexities may be associated with the disruption to the cohesion and consistency of a firm's internal (innovative) activity configuration. We therefore formulate our first hypothesis as follows:

H1: The introduction of the CII in France leads to an increase in innovative activity among SMEs.

The most widely studied firm characteristic in the context of heterogeneity of tax incentives is firm size. However, few studies contain evidence on whether the impact of tax incentives on innovative activity is related to firm age and to firm life cycle. Studies that consider whether the impact of tax incentives is related to firm age are still rare, although contributions exist, Coad et al. (2016) and Rao (2016), in particular. Their results indicate that younger firms react more strongly to tax incentives, especially in the short run.

Anthony and Ramesh (1992) pioneered the empirical measures for sorting firms in different life cycle stages by using classification variables such as dividend payout ratio, sales growth rate, capital expenditures and firm age. Economic theory suggests that firms make early investments in their growth stage to gain a competitive edge and early entry into the market. As the firm matures, it will cut new investments, reducing risky and innovative investments (Shahzad et al., 2022). This is in line with the results of Chang et al.

(2017), which indicate that managers of firms in their growth stage tend to increase R&D expenditures. As R&D expenditures possess uncertain benefits as compared to capital expenditures, it is necessary for managers to understand how and when to maximise the benefits from R&D (Chang et al., 2017). We therefore formulate our second hypothesis as follows:

H2: The positive effect of CII on innovation activities is more pronounced for SMEs in their growth stage.

In order to increase the effectiveness of R&D tax credits, Montmartin et al. (2018) conclude that the tax credit scheme should introduce a more targeted approach by implementing different levels of tax credits by regions and industries. Moreover, Castellacci and Lie (2015) argue that tax incentive schemes should be concentrated in industries with high technological opportunities in sectors that lead to strong spillover effects to the rest of the economy.

The OECD has drawn up a series of technology-themed classifications of economic activities. In the study by Hatzichronoglou (1997), the technology classification was created by clustering industries based on a measure of internal R&D intensity. A similar approach was used in Galindo-Rueda and Verger (2016), whose study represents an update and reframing of previous OECD taxonomies that were based on earlier versions of the International Standard Industrial Classification (ISIC). According to Castellacci and Lie (2015), R&D tax incentive schemes should be restructured to better incorporate sector-specific innovation drivers and to allocate a greater proportion of fiscal support to sectors with high opportunities and strong technological dynamism. We therefore formulate our third hypothesis as follows:

H3: The positive effect of CII on innovation activities is more pronounced for SMEs in high-tech industries.

3. Research methodology

3.1. Sample and data

We gather firm-specific information from the Bureau Van Dijk's Orbis Europe database, which contains financial statement, ownership and intellectual property data on European firms. We collect financial statement data from a sample of French SMEs and large firms. We obtain financial statement information of French firms' intangible fixed assets, number of employees, total assets, turnover, number of patents, year of establishment, paid dividends, sales, capital expenditures and industry classification.

These intangible fixed assets are needed as the expenditures incurred during the development phase of R&D projects will be recorded on this item of the balance sheet.[4] According to the generally accepted French accounting principles, firms have the option to capitalise expenditures incurred during the development phase of R&D projects if the following conditions are met: (1) the R&D project has a high probability of being successful, (2) the firm will be able to complete the R&D project, (3) the R&D project will generate future economic benefits. Furthermore, according to Art. 244 quarter B of the French Tax Code, it is only such expenditures that can be taken into account for the CII. The number of employees, total assets and turnover are the criteria that are used in order to determine whether a French firm is an SME. We collect data on the number of patents a firm has registered during its lifetime in order to determine whether the firm has engaged in innovative activities. The year of establishment is needed in order to determine a firm's age. The paid dividends, sales and capital expenditures are the criteria that are used in order to determine a firm's life cycle stage.

Our sample period covers the five-year period 2014–2018, thus going beyond the sample period 2013–2016 previously studied in Bunel and Hadjibeyli (2021). Although the CII was introduced in late 2013, it only came into effect on 1 January 2014. We do not extend the sample period beyond 31 December 2018, since the French definition of an SME was revised in 2019 with the PACTE law. With this law, two separate definitions were introduced for small firms and medium-sized firms. We select French firms whose book value of intangible fixed assets was greater than or equal to 1 euro during the entire sample period, whose registered number of patents during its lifetime was at least equal to one, and whose number of employees in 2012 was not missing. This approach resulted in an unbalanced panel data set of 2,023 firms over 5 years, which, due to 231 missing values for the variable turnover, led to 9,884 firm-year observations.

Table 2 presents the descriptive statistics of our full sample. Due to the distributions of our outcome variable and the criteria for the SME definition being highly skewed, we decided to take their natural logarithms, as working with the natural logarithm of a variable often helps to deal with outliers and heteroskedasticity (Verbeek, 2017).

It can be seen from Table 2 that after the logarithmic transformations, these variables appear to be symmetrically distributed. The criteria for determining a firm's life cycle, i.e. the dividend payout ratio, sales growth ratio and capital expenditures, all in percentages, are highly skewed. The median firm age is 31 years, and its distribution is skewed to the right.

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[4] Intangible fixed assets include more than the development costs incurred during R&D projects. While all firms in our sample are innovative in nature, i.e. they have registered at least one patent application during their lifetime, we lack the data isolating these development costs from the rest of the intangible fixed assets.

Table 2. Descriptive statistics

Panel A: Full sample, all firms Number of observations: 9,884					
	Mean	**Standard deviation**	**Q1**	**Median**	**Q3**
Outcome variable (after logarithmic transformation)					
IFA	13.2854	3.3544	11.0696	12.9244	15.1543
Criteria for the SME definition (after logarithmic transformations)					
Total assets	17.2936	2.2981	15.7520	16.9459	18.5123
Turnover	17.3844	2.1630	16.0178	17.1504	18.5157
Number of employees	4.7901	2.0654	3.4657	4.5109	5.8319
Criteria for the firm life cycle (in %, except for firm age)					
Dividend/ EBIT	0.1082	4.4523	0.0000	0.0000	0.0000
Sales growth	44.8012	1318.9754	–4.5478	2.1181	9.8365
Capex/Assets	50.3541	239.3572	12.2884	30.0424	56.7311
Firm age	36.58	23.83	21.00	31.00	48.00
Panel B: Firms in their growth stage Number of observations: 2,793					
	Mean	**Standard deviation**	**Q1**	**Median**	**Q3**
Outcome variable (after logarithmic transformation)					
IFA	13.0691	2.83617	11.1138	13.0077	14.8637
Criteria for the SME definition (after logarithmic transformations)					
Total assets	16.8200	1.9520	15.5180	16.7126	17.9554
Turnover	16.8501	1.8297	15.6367	16.7076	17.9287
Number of employees	4.4951	1.6727	3.4012	4.3631	5.5175
Criteria for the firm life cycle (in %, except for firm age)					
Dividend/ EBIT	0.0000	0.0000	0.0000	0.0000	0.0000
Sales growth	90.1923	1800.7956	3.9425	8.3674	16.4255
Capex/Assets	86.6764	109.9802	40.9582	58.1018	88.1124
Firm age	20.70	6.84	15.00	21.00	27.00

Table 2 continued

Panel C: Firms in high-tech industries Number of observations: 3,373					
	Mean	**Standard deviation**	**Q1**	**Median**	**Q3**
Outcome variable (after logarithmic transformation)					
IFA	13.6688	3.3344	11.3840	13.2615	15.5581
Criteria for the SME definition (after logarithmic transformations)					
Total assets	17.5733	2.2217	16.0367	17.1442	18.8076
Turnover	17.6319	2.1348	16.1854	17.3057	18.8162
Number of employees	5.2195	1.9873	3.8712	4.9200	6.2634
Criteria for the firm life cycle (in %, except for firm age)					
Dividend/ EBIT	0.2979	2.2233	0.0000	0.0000	0.0000
Sales growth	31.1328	487.0278	–3.4361	3.0965	10.6174
Capex/Assets	56.8648	176.4442	16.5019	34.1455	61.4719
Firm age	36.35	22.54	21.00	31.00	47.00

Note: IFA: intangible fixed assets, EBIT: Earnings before interest and taxes, Capex: Capital expenditures.

Source: own calculations on the basis of Orbis Europe.

3.2. Research design

We test our hypotheses by using a regression discontinuity design (hereinafter, RDD) and by comparing the level of innovative activity of French SMEs and large French firms after the introduction of the CII in late 2013. The RDD has become one of the leading quasi-experimental strategies in economics and many other social and behavioural sciences (Lee & Lemieux, 2010). In RDD, the research units are assigned to the treatment group based on the value of the running variable, i.e. the variable which determines treatment, with the probability of treatment assignment jumping discontinuously at a known cut-off (Angrist & Pischke, 2014). Since the number of employees is the most binding criterion in determining whether a French firm is an SME or not, it becomes our running variable of choice for our RDD.[5]

[5] Sensitivity checks based on totals assets or turnover can be found in the Appendix. See also the robustness checks section.

We test the validity of the running variable by performing the McCrary test, which estimates the discontinuity of the running variable at the cut-off (Dechezleprêtre et al., 2023). Figure 1 shows that the distribution of firms' number of employees in 2012 appears to be continuous around the threshold equal to the natural logarithm of 250.

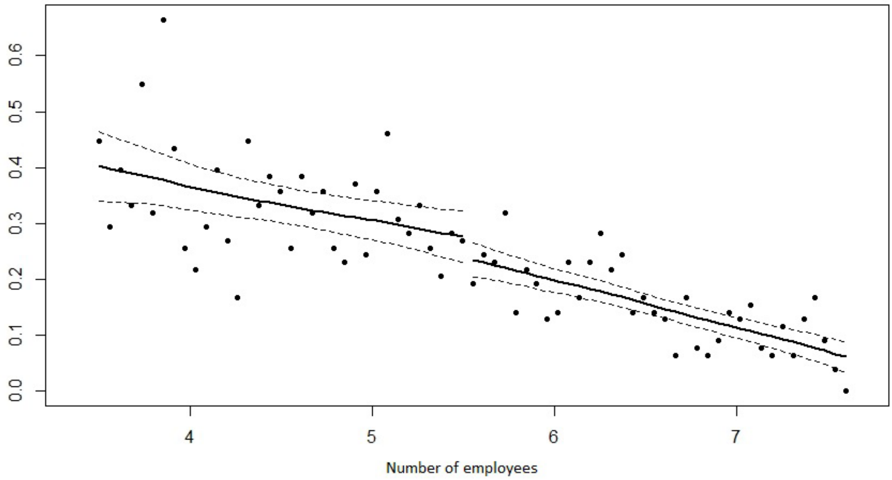


Figure 1. McCrary test at the SME threshold (number of employees) in 2012

Note: The density of number of employees is displayed on the vertical axis. The number of employees is displayed on the horizontal axis, in a natural logarithmic scale.

Source: own calculations on the basis of Orbis Europe.

The McCrary test gives a discontinuity estimate of –0.1511, which is statistically insignificantly different from zero (p-value = 0.2009). We thus cannot reject the null hypothesis of no discontinuity in the density of the running variable at the cut-off and conclude that the firms in our sample did not precisely manipulate their number of employees in the year 2012.

3.3. Econometric model

For our RDD, we estimate the following log-linear model:

$$\ln(IFA_{i,t}) = \alpha_0 + \beta_1 \ln(a_{i,2012}) + \beta_2 SME_{i,t} + \gamma_1 \overline{\ln(IFA_{i,past})} + \varepsilon_{i,t} \qquad (1)$$

The left-hand side of equation (1) contains the outcome variable, which measures innovative activity. In line with previous studies (e.g., Alstadsaeter et al., 2018; Ernst & Spengel, 2011; Karkinsky & Riedel, 2012), the amount of the intangible fixed assets of firm i in year t will be our measure for innovative activity. The intangible fixed assets displayed on the balance sheet contain the expenditures incurred during the development phase of R&D projects and, according to Art. 244 quarter B of the French Tax Code, these expenditures can be taken into account for the CII. We measure the amount of the intangible fixed assets of firm i both in levels and in first differences.

The right-hand side of equation (1) includes a treatment dummy *SME*, which will be equal to 1 if firm i is considered to be an SME in year t and 0 otherwise. The right-hand side of equation (1) also includes the running variable *a*, which is equal to the number of employees firm i had in 2012, i.e. the year prior to the reform. Because of the two-year rule, a firm's SME status in 2014 was partly based on its financial information in 2012 and 2013. Using the number of employees in 2012 as our running variable of choice mitigates the concern that there might have been endogenous sorting of firms across the SME threshold (Dechezleprêtre et al., 2023). Finally, the right-hand side of equation (1) also includes the average value of a firm's intangible fixed assets during the pre-treatment period 2008–2012, as we control for past innovative activity.

As mentioned earlier, a firm's number of employees is not the only criterion for determining whether a French firm is an SME or not. The firm's total assets and turnover also play an important role, albeit a less binding one. In this setting, a *fuzzy* RDD, which is based on an instrumental variable approach, might be a remedy (Stock & Watson, 2015). For our first-stage regression, we estimate the following linear probability model:

$$SME_{i,t} = \alpha_1 + \beta_3 \ln(a_{i,2012}) + \beta_4 D_{i,2012} + \gamma_2 \overline{\ln(IFA_{i,past})} + \mu_{i,t} \tag{2}$$

The right-hand side of equation (2) contains a new dummy variable, D_i, which will be equal to 1 if firm i had less than 250 employees in 2012, i.e. the year prior to the reform. Due to the CII being introduced in late 2013, French firms were unable to precisely manipulate the value of the running variable around the cut-off in 2012. Under this assumption, D_i is as good as randomly assigned at the cut-off (Dechezleprêtre et al., 2023; Lee & Lemieux, 2010). As the number of employees in 2012 does not determine post-reform SME status perfectly, equation (2) represents the reduced form of a fuzzy RDD, in which D_i is the instrument for firm's i actual SME status and its eligibility to the CII. For our second-stage regression, we estimate the following log-linear model:

$$\ln(IFA_{i,t}) = \alpha_2 + \beta_5 \ln(a_{i,2012}) + \beta_6 \widehat{SME}_{i,t} + \gamma_3 \overline{\ln(IFA_{i,past})} + \omega_{i,t} \tag{3}$$

The right-hand side of equation (3) includes a new variable, $\widehat{SME}$, which contains the fitted values of the first-stage regression. In a two-stage least squares (2SLS), the consistency of the second-stage estimates is not based on getting the first-stage functional form right. This means that using a linear regression for the first-stage estimates generates consistent second-stage estimates even with a dummy endogenous variable (Angrist & Pischke, 2014).

4. Results

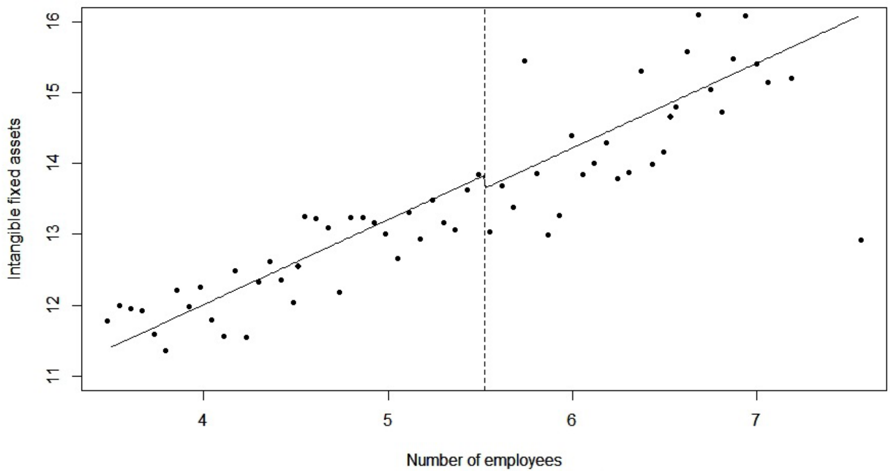
4.1. Main regression analyses

In Figure 2, we present two graphs that show the relationship between the outcome and the running variable, during the post-treatment period (Panel A) and the pre-treatment period (Panel B). In both graphs, our outcome variable is a firm's intangible fixed assets, which is our measure for innovative activity. Our running variable is a firm's number of employees, which is the most binding criterion in determining whether a French firm is an SME or not. A natural logarithmic scale is used for both variables. The cut-off value represented by the dashed line is equal to the natural logarithm of 250, which approximates the value of 5.52.

In Panel A, we observe a jump in the outcome variable during the post-treatment period, i.e. from 1 January 2014 to 31 December 2018. In Panel B, we do not observe a jump in the outcome variable during the pre-treatment period 2008–2012 prior to the introduction of the CII. Thus, only Panel A shows a jump in a firm's intangible fixed assets at the threshold of 250 employees, suggesting that firms below the threshold increase their innovative activity as a result of the treatment and not due to other factors in the pre-treatment period.

When introducing our econometric model in the previous section, we assumed a linear relationship between our running variable and the outcome variable. In general, as in any other setting, there is no particular reason to believe that the true model is linear (Lee & Lemieux, 2010). Angrist and Pischke (2014) point out that the problem of distinguishing jumps from non-linear trends diminishes as we concentrate on observations close to the cut-off, which in our case is the number of employees being equal to 250. This suggests an approach that compares averages in a narrow bandwidth just to the left and just to the right of the cut-off. We balance the reduction in bias near the cut-off against the increased variance suffered by fewer observations, generating an optimal bandwidth. We calculate this optimal bandwidth following the procedure in Imbens and Kalyanaraman (2012) and are left with a sample containing 6,685 firm-year observations.

Panel A: Post-treatment period (2014–2018)



Panel B: Pre-treatment period (2008–2012)

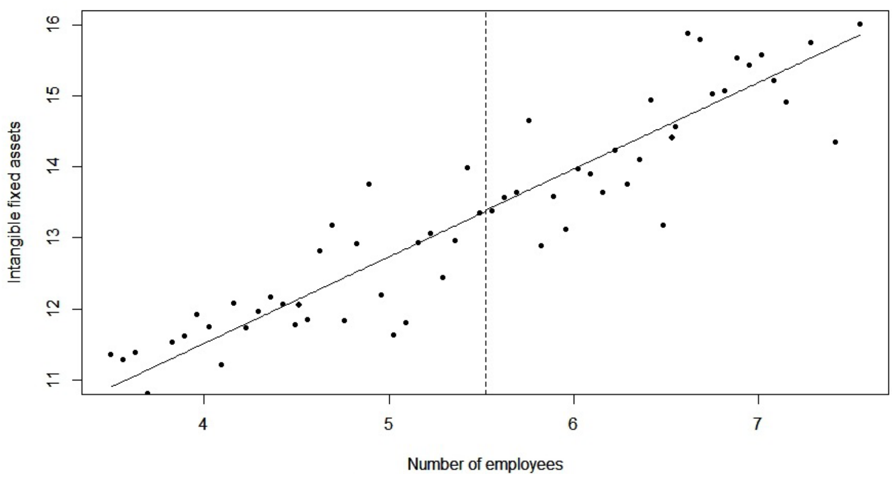


Figure 2. Graphical presentation of the regression discontinuity design

Note: A natural logarithmic scale is used for number of employees and intangible assets. The dashed line represents the cut-off value, which is equal to ln(250).

Source: own calculations on the basis of Orbis Europe.

The first column of Table 3, Panel A contains the results of our main regression analyses for all firms whose number of employees lies within the optimal bandwidth. A firm's intangible fixed assets, measured in levels, is the outcome variable. We regress our outcome variable on the running variable employees, our variable of interest, namely *SME*, and on the average amount of a firm's intangible fixed assets during the five years preceding the reform, i.e. the period 2008–2012.

Table 3. Main regression results

Panel A

Outcome variable: ln(Intangible fixed assets)			
Estimator: Two-stage least squares (2SLS)			
Firms	**All firms (within bandwidth)**	**Growth stage (within bandwidth)**	**High-tech (within bandwidth)**
Intercept	2.7994*** (0.4296)	2.9435*** (0.8625)	2.5889*** (0.7306)
ln(Employees)	0.5872*** (0.0592)	0.6501*** (0.1169)	0.7346*** (0.1011)
SME	0.4211** (0.1662)	0.6119* (0.3306)	0.5000* (0.2747)
ln(Past IFA)	0.5636*** (0.0084)	0.5398*** (0.0171)	0.5139*** (0.0137)
Observations	6,685	1,795	2,475

Panel B

Outcome variable: First difference of ln(Intangible fixed assets)			
Estimator: Two-stage least squares (2SLS)			
Firms	**All firms (within bandwidth)**	**Growth stage (within bandwidth)**	**High-tech (within bandwidth)**
Intercept	0.1344 (0.1693)	0.3182 (0.2867)	0.5279* (0.2764)
ln(Employees)	0.0123 (0.0232)	–0.0017 (0.0387)	–0.0428 (0.0382)
SME	0.0146 (0.0661)	–0.0735 (0.1108)	–0.1052 (0.1051)
ln(Past IFA)	–0.0151*** (0.0032)	–0.0174*** (0.0056)	–0.0185*** (0.0049)
Observations	5,348	1,436	1,980

Note: IFA: intangible fixed assets. We report heteroskedasticity-and-autocorrelation-consistent standard errors in parentheses. ***, **, and * denote a significant difference at the 1%, 5% and 10% levels, respectively.

Source: own calculations on the basis of Orbis Europe.

Our variable of interest, the treatment dummy *SME*, has a statistically significant positive effect at the 5% level on a firm's intangible fixed assets during the post-treatment period. Looking at the economic order of magnitude, it is evident that SMEs invested on average 42% more in intangible fixed assets than large firms in our sample during the post-treatment period. Employees and the average amount of a firm's intangible fixed assets prior to the reform also have a statistically significant positive effect on the outcome variable. These results do support hypothesis 1, stating that the introduction of the CII in France would lead to an increase in innovative activity among SMEs. Although measured in number of patents, Dechezleprêtre et al. (2023) find a comparable large innovation increase (58%) in the UK as a response to changing SME thresholds for R&D tax incentives.

Next, we identify a firm's life cycle stage (Anthony & Ramesh, 1992; Chang et al., 2017). Firms in their growth stage are young and usually exhibit lower dividend payout ratios, higher sales growth rates and have more capital expenditures. We perform similar regression analyses with a subsample containing the firms in their growth stage. We identify firms in their growth stage as firms with below median age, below median dividend payout ratio, above median sales growth rate, and above median capital expenditures using five-year historical data. All four criteria must be fulfilled in order for a firm to be classified as being in the growth stage. Once more, we calculate the optimal bandwidth following the procedure in Imbens and Kalyanaraman (2012) and are left with a subsample containing 1,795 firm-year observations.

The second column of Table 3, Panel A contains the results for the firms in their growth stage whose number of employees lies within the optimal bandwidth. The treatment dummy *SME* has a statistically significant positive effect on the level of a firm's intangible fixed assets during the post-treatment period, albeit at the 10% significance level. Looking at the economic order of magnitude, we can see that SMEs in their growth stage invested on average 61% more in intangible fixed assets than large firms in their growth stage during the post-treatment period. This coefficient is greater than the one reported in the first column of Table 3, Panel A, when looking at all SMEs in our sample (42%), irrespective of their life cycle stage. This result does support hypothesis 2, assuming that the positive effect of the CII in France on innovation activities is more pronounced for SMEs in their growth stage.

Next, we identify firms in high-tech industries based on the OECD's technology-themed classifications of economic activities, as can be found in the study of Hatzichronoglou (1997). Hatzichronoglou (1997) created these technology-themed classifications by clustering industries into four clusters—high-tech, medium-high-tech, medium-low-tech and low-tech—based on a measure of internal R&D intensity.

An updated version of the OECD's technology-themed classifications of economic activities can be found in Galindo-Rueda and Verger (2016). Once more, these technology-themed classifications were created by clustering industries based on a measure of internal R&D intensity. The above study classified industries into five clusters: high-tech, medium-high-tech, medium, medium-low-tech and low-tech.

We perform similar regression analyses with a subsample containing the firms in high-tech industries. We identify firms in high-tech industries as firms in high-tech and medium-high-tech industries, according to the updated OECD's classification. Once more, we calculate the optimal bandwidth following the procedure described in Imbens and Kalyanaraman (2012), and are left with a subsample containing 2,475 firm-year observations.

The third column of Table 3, Panel A contains the results for firms in high-tech industries whose number of employees lies within the optimal bandwidth. In the post-treatment period, *SMEs* show a statistically significant positive effect at the 10% level on a firm's intangible fixed assets. In terms of economic magnitude, we can see that SMEs in high-tech industries invested on average 50% more in intangible fixed assets than large firms in high-tech industries during the post-treatment period. This coefficient is greater than the one reported in the first column (42%), when looking at all SMEs in our sample, irrespective of their industry. This result does support hypothesis 3 that the positive effect of the CII on innovation activities is more pronounced for SMEs in high-tech industries.

In Table 3, Panel B, the outcome variable firm's intangible assets is now measured in first differences. In line with Panel A, the columns contain the results within the optimal bandwidth for all firms, firms in their growth stage, and firms in high-tech industries, respectively. Again, we regress our outcome variable on the running variable employees, our variable of interest *SME*, and on the average amount of a firm's intangible fixed assets during the five years preceding the reform, i.e. the period 2008–2012. When a firm's intangible fixed assets are measured in first differences, however, the treatment dummy *SME* no longer has a statistically significant effect. The factor 'employees' also has no statistically significant effect, while the average amount of a firm's intangible fixed assets prior to the reform now has a statistically significant negative effect on the outcome variable.

In brief, when the outcome variable is measured in levels, the results in Table 3 support hypothesis 1, 2 and 3. However, when intangible fixed assets are measured in first differences, the results indicate that this positive effect for SMEs is a one-time level shift, as it has no effect in terms of growth rates. This is in accordance with the study by Rao (2016), who noted that SMEs have a stronger immediate response to R&D tax incentives but go on to reduce their research spending in future years, unlike larger firms.

4.2. Robustness tests

RDD does not guarantee to produce reliable causal estimates, since one cannot be certain of a linear relationship between the running variable and the outcome variable. There is also the risk of confusing nonlinearities with discontinuities (Angrist & Pischke, 2014). It is therefore essential to explore how RDD estimates are robust to the inclusion of higher-order polynomial terms and to changes in the bandwidth around the cut-off (Lee & Lemieux, 2010). In accordance with the study by Lee and Lemieux (2010), we perform two additional robustness tests by including a smaller bandwidth as well as the squared value of the running variable to account for nonlinearity. The smaller bandwidth increases the comparability between the treatment group and the control group, but reduces the initial sample to 5,705 firm-year observations.

Table 4 presents the findings within the smaller optimal bandwidth, including squared values for our running variable employees. In line with our main analyses, results are reported for all firms, firms in their growth stage, and firms in high-tech industries.

When intangible assets are measured in levels (Panel A), the treatment dummy *SME* has a statistically significant positive effect at the 1% level on this outcome variable during the post-treatment period. Thus, in line with the main analyses, hypothesis 1 is supported. Moreover, hypothesis 2 is supported, as SMEs in their growth stage show a significant positive effect at the 5% level and invest on average 80% more in intangible fixed assets than large firms during the post-treatment period. This coefficient is greater than the one reported for all SMEs in our sample (70%), irrespective of their life cycle stage. The robustness tests for high-tech firms are somewhat weaker compared to our main analyses. The treatment dummy *SME* has a statistically significant positive effect on a firm's intangible fixed assets during the post-treatment period at the 10% level. However, SMEs in high-tech industries invested on average 54% more in intangible fixed assets than large firms in high-tech industries. This economic effect is smaller compared to SMEs active in all kinds of industries (70%). Therefore, when applying robustness tests, the empirical findings do not support hypothesis 3.

In Table 4, Panel B, the outcome variable firm's intangible assets is now measured in first differences. We find similar results as for our main analyses considering the first differences. *SME* no longer has a statistically significant effect.

In brief, after having performed these robustness tests, our conclusion does not change as regards hypotheses 1 and 2. Following the introduction of the CII in France, innovative activity increases for SMEs, relative to large firms. Moreover, the positive effect is more pronounced for SMEs in their growth stage. However, when intangible fixed assets are measured in first differences, the results indicate that this positive effect for SMEs does not persist over time.[6]

[6] We furthermore perform sensitivity checks based on total assets and turnover, which can be found in the Appendix. The results are not in line with our main regression analyses and robustness tests. These divergent results can be explained by the fact that a fuzzy RDD, being underpinned by an instrumental variable approach, requires a strong instrument. In the French setting, a firm's total assets and turnover are weak instruments to predict for a firm's SME status compared to the number of employees.

Table 4. Robustness tests

Panel A

Outcome variable: ln(Intangible fixed assets)			
Estimator: Two-stage least squares (2SLS)			
Firms	**All firms (within bandwidth)**	**Growth stage (within bandwidth)**	**High-tech (within bandwidth)**
Intercept	2.3126*** (0.8968)	4.0745*** (1.2185)	3.5134** (1.4958)
ln(Employees)	0.3759 (0.2887)	0.0688 (0.4197)	0.2842 (0.4879)
ln(Employees)^2	0.0272 (0.0263)	0.0617 (0.0413)	0.0373 (0.0448)
SME	0.7010*** (0.1824)	0.8013** (0.3373)	0.5419* (0.3002)
ln(Past IFA)	0.6080*** (0.0094)	0.5435*** (0.0171)	0.5362*** (0.0148)
Observations	5,705	1,515	2,155

Panel B

Outcome variable: First difference of ln(Intangible fixed assets)			
Estimator: Two-stage least squares (2SLS)			
Firms	**All firms (within bandwidth)**	**Growth stage (within bandwidth)**	**High-tech (within bandwidth)**
Intercept	0.2993 (0.3492)	0.1988 (0.5661)	0.3369 (0.5642)
ln(Employees)	–0.0848 (0.1102)	–0.0604 (0.1948)	0.0089 (0.1789)
ln(Employees)^2	0.0113 (0.0099)	0.0113 (0.0189)	–0.0029 (0.0162)
SME	0.0655 (0.0742)	0.0489 (0.1293)	–0.0715 (0.1187)
ln(Past IFA)	–0.0159*** (0.0036)	–0.0146** (0.0062)	–0.0202*** (0.0054)
Observations	4,564	1,212	1,724

Note: IFA: intangible fixed assets. We report heteroskedasticity-and-autocorrelation-consistent standard errors in parentheses. ***, **, and * denote a significant difference at the 1%, 5% and 10% levels, respectively.

Source: own calculations on the basis of Orbis Europe.

Our findings are in line with previous studies which suggest that R&D tax credits targeted at SMEs or offering preferential regimes to SMEs, relative to large firms, positively influence SMEs' decisions to conduct more R&D (Agrawal et al., 2020; Dechezleprêtre et al., 2023; Kobayashi, 2014). Bunel and Hadjibeyli (2021) find an increase in employment in the short term, along with an increase in turnover in the medium term for the French CII, and our results complement theirs. In particular, our results show an immediate, though not sustained, increase in the amount of intangible fixed assets. Also, our findings are in line with previous studies suggesting that firms innovate more in their growth stage (Chang et al., 2017; Shahzad et al., 2022). Hence, this study demonstrates that, despite being only applicable to expenditures incurred during the development phase of R&D projects, the French targeted tax credit is still effective, at least in the short run. The effect might have been bolstered by the immediate refund available to liquidity-constrained SMEs.

Conclusions

In order to stimulate private R&D, the governments of most Western nations have introduced tax incentives designed to complement subsidies. While tax incentives are essentially a generic policy instrument, targeting specific groups of firms is quite common, and many European countries target SMEs and young firms. The empirical literature has investigated the effects of some tax incentives targeted at SMEs in the United Kingdom, Canada and Japan, as the level of R&D performed by SMEs is crucial to a country's technological progress and economic growth. Introduced in late 2013, the CII serves as an extension to the already existing R&D tax credit available to all French firms. One notable characteristic of the CII that distinguishes it from other R&D tax credits is that it only applies to expenditures incurred during the development phase of R&D projects.

The aim of our research was to investigate whether the CII was successful at promoting R&D for French SMEs. We explored this by implementing a fuzzy regression discontinuity design and by comparing the innovative activity of SMEs with the innovative activity of larger firms in France over the period 2014–2018. We also investigated whether SMEs reacted more strongly to this targeted tax incentive in their growth stage and in high-tech industries.

In general, our results demonstrate that the CII has a positive impact on their level of innovative activity. In particular, SMEs invested on average 42% more in intangible fixed assets than large firms during the post-treatment period. Our results also demonstrate that this impact is more pronounced for SMEs in their growth stage. However, we do not find evidence that this impact is more pronounced for SMEs in high-tech industries. Furthermore, our analysis indicate that the positive increase for SMEs is a one-time level shift. We do not find any effect on growth rates, thus CII does not lead to accumulating effects over time.

The findings might be of interest to policy makers assessing the design and implementation of R&D tax incentives and their capacity to benefit certain target groups. The French R&D tax credit, with its unique features, seems successful at promoting the level of innovative activity of SMEs, albeit only in the short term. Our study has its limitations, as we were unable to exclude the non-R&D items (e.g., goodwill) from the intangible fixed assets. Moreover, we solely focused on investigating French firms. A study of recently introduced targeted R&D tax incentives in other countries might be an important avenue for further research.

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Appendix

Table A1. Sensitivity analyses (total assets)

Panel A

Outcome variable: ln(Intangible fixed assets)			
Estimator: Two-stage least squares (2SLS)			
Firms	**All firms (within bandwidth)**	**Growth stage (within bandwidth)**	**High-tech (within bandwidth)**
Intercept	–2.3289*** (0.7121)	4.4315** (1.7834)	–4.4900*** (1.5408)
ln(Employees)	0.5234*** (0.0357)	0.2600*** (0.0886)	0.6711*** (0.0783)
SME	0.1173 (0.1494)	–1.5037*** (0.3907)	0.3918 (0.3000)
ln(Past IFA)	0.4999*** (0.0081)	0.4328*** (0.0161)	0.4619*** (0.0132)
Observations	6,685	1,795	2,475

Panel B

Outcome variable: First difference of ln(Intangible fixed assets)			
Estimator: Two-stage least squares (2SLS)			
Firms	**All firms (within bandwidth)**	**Growth stage (within bandwidth)**	**High-tech (within bandwidth)**
Intercept	0.0685 (0.2828)	1.2039** (0.6389)	0.4207 (0.5880)
ln(Employees)	0.0092 (0.0142)	–0.0431 (0.0319)	–0.0091 (0.0299)
SME	–0.0026 (0.0595)	–0.2758* (0.1399)	–0.0379 (0.1148)
ln(Past IFA)	–0.0167*** (0.0032)	–0.0192*** (0.0055)	–0.0182*** (0.0048)
Observations	5,348	1,436	1,980

Note: IFA: intangible fixed assets. We report heteroskedasticity-and-autocorrelation-consistent standard errors in parentheses. ***, **, and * denote a significant difference at the 1%, 5% and 10% levels, respectively.

Source: own calculations on the basis of Orbis Europe.

Table A2. Sensitivity analyses (turnover)

Panel A

Outcome variable: ln(Intangible fixed assets)			
Estimator: Two-stage least squares (2SLS)			
Firms	**All firms (within bandwidth)**	**Growth stage (within bandwidth)**	**High-tech (within bandwidth)**
Intercept	3.0586*** (0.8149)	10.3271*** (1.6446)	1.6932 (1.3115)
ln(Employees)	0.2011*** (0.0412)	–0.1389* (0.0839)	0.3025*** (0.0672)
SME	–0.5774*** (0.1437)	–1.6314*** (0.3026)	–0.3370 (0.2259)
ln(Past IFA)	0.5453*** (0.0083)	0.5128*** (0.0167)	0.5064*** (0.0136)
Observations	6,685	1,795	2,475

Panel B

Outcome variable: First difference of ln(Intangible fixed assets)			
Estimator: Two-stage least squares (2SLS)			
Firms	**All firms (within bandwidth)**	**Growth stage (within bandwidth)**	**High-tech (within bandwidth)**
Intercept	0.3719 (0.3107)	0.6082 (0.5474)	1.2368*** (0.4706)
ln(Employees)	–0.0090 (0.0157)	–0.0159 (0.0279)	–0.0512** (0.0241)
SME	–0.0295 (0.0550)	–0.1172 (0.1013)	–0.1647** (0.0815)
ln(Past IFA)	–0.0143*** (0.0031)	–0.0171*** (0.0054)	–0.0186*** (0.0048)
Observations	5,348	1,436	1,980

Note: IFA: intangible fixed assets. We report heteroskedasticity-and-autocorrelation-consistent standard errors in parentheses. ***, **, and * denote a significant difference at the 1%, 5% and 10% levels, respectively.

Source: own calculations on the basis of Orbis Europe.

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Attitudes towards income inequality and trust: An analysis by income class in Poland

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Abstract

Increasing income inequality has raised concerns about social cohesion, yet the subjective dimension of inequality and its relationship to trust remain underexplored. This article examines links between attitudes toward income inequality and generalised and institutional trust in Poland, a post-socialist state characterised by strong anti-inequality sentiment and low trust. Using data from the 5th wave of the European Values Study (N = 1,352), we employ an economic stratification framework with five income classes, complemented by non-parametric tests and logistic regression. The results show that acceptance of inequality increases with income, with the sharpest contrasts between low- and high-income classes, while middle strata remain relatively homogeneous. Generalised trust rises with income, whereas institutional trust follows more complex, non-linear patterns. Crucially, the links between trust and inequality attitudes are class-specific: generalised trust in strangers legitimises inequality overall, while generalised trust in relatives has divergent effects across lower- and upper-middle-income groups.

Keywords

- income inequality
- inequality perception
- income class
- generalised trust
- institutional trust
- economic stratification
- informal institutions
- institutional economics

JEL codes: D02, D31, D63, P36, Z13

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Introduction

In recent decades, income inequality has become a central concern in economic research. Existing studies have predominantly focused on its actual level, sources, and structural consequences (Deaton, 2016; Piketty, 2015; Stiglitz, 2015). The literature consistently links rising inequality to adverse social outcomes, such as diminished quality of life, heightened social tensions, and the erosion of institutional legitimacy (OECD, 2019). Wilkinson and Pickett (2011) further emphasise its broader societal consequences, demonstrating that societies with greater income disparities tend to experience higher levels of anxiety, weaker social cohesion, and lower trust. Evidence increasingly links attitudes toward income inequality to trust, a key mechanism for social interaction. A substantial body of research underscores its pivotal role in facilitating interactions among actors, fostering cooperation, and enabling collective solutions to be developed. Specifically, trust contributes to the reduction of transaction costs, encourages desired individual behaviours, and motivates actors to act by existing rules (e.g., Bjørnskov, 2012, 2022; Bjørnskov & Méon, 2015; Guiso et al., 2006; Knack & Keefer, 1997; Tabellini, 2010; Zak & Knack, 2001). Societies with higher trust tend to exhibit greater equality, yet findings remain ambiguous.

Although the link between trust and income inequality has been studied, the subjective perspective on income inequality remains underexplored. Therefore, we focus on attitudes toward income inequality rather than objective measures such as the Gini index, capturing how individuals perceive economic disparities and how these perceptions relate to trust. There is a growing focus on how people assess their social and economic environment, including their perception of income inequality (Knell & Stix, 2020). Scholars suggest that perceptions of inequality may be a more reliable predictor of responses to economic disparities than objective measures (Uddin, 2025, p. 2). Knell and Stix (2021, p. 801) claim that "an unbiased estimate of the effect of inequality on trust can be obtained with a measure of individual-specific perceptions of inequality." Even if perceptions do not accurately reflect actual inequality levels, tensions in society arise when they are perceived as unfair (Zmerli & Castillo, 2015, p. 180). Subjective income inequality is important to the functioning of societies, for example, it can erode trust in political and economic systems (Deaton, 2016).

Moreover, considering the income class gradient in inequality acceptance observed in Europe (Szczepaniak et al., 2025), we employ an economic stratification perspective to capture how individuals' positions within the income distribution condition both their attitudes toward inequality and their orientations toward trust. This approach highlights relational differences between income classes often obscured in aggregate-level analyses, thereby offering a nuanced understanding of how trust and inequality attitudes are embedded in social hierarchies.

Given the distinct origins and implications of various forms of trust, our study categorises them. Generalised trust extends beyond face-to-face interactions, encompassing individuals with whom one is not personally acquainted (Rothstein & Stolle, 2008, p. 441). Institutional trust, a form of particularised trust, concerns well-identified subjects and arises from interactions between citizens and public institutions (Bjørnskov, 2006, p. 2; Camussi & Mancini, 2019, pp. 489–490). Furthermore, research has yielded weak and inconsistent evidence that individuals with high levels of social trust also display greater political trust (Zmerli & Newton, 2008), suggesting the need to analyse the categories of trust in more detail. Studies on inequality primarily focus on one category of trust, most frequently generalised trust, and occasionally on institutional trust. Meanwhile, recent crises have eroded trust in institutions, undermining the credibility of governments and public administration in Western economies (Algan et al., 2017; Botsman, 2018; Camussi & Mancini, 2019; Hetherington, 1998; Inglehart, 1997). Institutional trust, therefore, remains a vital research topic. Since trust is often used as a proxy for informal institutions (Cruz-García & Peiró-Palomino, 2019; Tabellini, 2010; C. Williamson, 2009), this study contributes to the field of institutional economics by examining how informal institutions, particularly trust, relate to attitudes towards income inequality in terms of economic stratification. This perspective offers valuable insights, especially given the underexplored role of informal institutions (Bentkowska, 2024, p. 28).

In our analysis, we focus on Poland, a post-socialist country, where income inequality rose during the transition to a market economy (Bukowski et al., 2023; Bukowski & Novokmet, 2021). Negative attitudes towards income inequality are now widespread, with the vast majority of the population considering the current level of inequality excessive (Berlingieri et al., 2023; CBOS, 2024a; OECD, 2021). Poland thus represents a post-socialist EU economy characterised by high growth and development (Piątkowski, 2023), yet also exhibiting a strong anti-inequality sentiment (Bukowski et al., 2024). Research indicates that post-socialist societies tend to be less trusting due to past regimes and the upheaval following their collapse (Bjørnskov, 2006, p. 15). Totalitarian regimes fostered fear and distrust, eroding social capital (Paldam & Svendsen, 2001). The institutional changes forced discontinuity and adaptation to unfamiliar structures

(Mishler & Rose, 2001, p. 31). These effects may persist, and levels of trust in society appear relatively stable over time (Bjørnskov, 2006, p. 3). Polish society, like other post-socialist countries (Rothstein & Uslaner, 2005, p. 48), continues to exhibit comparatively low trust. Nearly three-quarters of Poles believe one cannot be too careful with others, a pattern stable for two decades (CBOS, 2024b, pp. 1–2). Only 38% report confidence in the government, compared with 60–80% in countries such as Switzerland and Luxembourg (OECD, 2022). Despite the salience of inequality perceptions and persistently low levels of trust, few studies examine their link within an income-stratified framework in Poland. Existing research examines either perceptions of inequality (Czerniak et al., 2018; Domański, 2010; Litwiński et al., 2023; Szczepaniak, 2025) or trust (in a broader comparative perspective) (Bjørnskov, 2006; Rothstein & Uslaner, 2005), but not their intersection.

This article addresses this gap by examining how attitudes toward income inequality relate to generalised and institutional trust across income classes in Poland. Using data from the European Values Survey, we apply nonparametric tests and logistic regression models to answer: (1) How do attitudes towards income inequality vary across income classes? (2) How do levels of generalised and institutional trust differ across income classes? and (3) To what extent are attitudes towards income inequality associated with different dimensions of trust, and do these relationships vary across income classes?

The remainder of the paper is organised as follows: Section 1 presents the key concepts, focusing on perceived income inequality and its links to different forms of trust. Section 2 describes the data and methods. Next, after presenting the results, we discuss the empirical findings. The final section provides conclusions.

1. Literature review

Subjective income inequality is more relevant for understanding social and political outcomes than objective measures of income inequality (Gimpelson & Treisman, 2018) because of biased perceptions of income inequality (Engelhardt & Wagener, 2014). Szirmai (1988) demonstrated that public opinion on inequality is remarkably persistent over time and deeply rooted in cultural and historical contexts. Consequently, in societies with egalitarian traditions, such as post-socialist countries, negative attitudes toward inequality can persist even after structural reforms. In Poland, in the early post-socialist period, inequality was often regarded as a temporary and acceptable cost of modernisation (Gijsberts, 2002). Yet over time, prolonged exposure to market-based stratification has eroded these legitimising beliefs, particularly among groups facing blocked upward mobility (Grosfeld & Senik, 2008). Drawing on Hirschman's "tunnel effect" hypothesis, Durongkaveroj (2025) argues that tolerance of inequality declines when rising aspirations are not matched by actual mobility, an experience common in Poland during the later stages of the transition period. Since EU accession, survey evidence indicates that about 90% of Poles view the income gap as too wide (CBOS, 2017; 2024a).

In the economics literature, it is noted that an individual's position in the income distribution, specifically their income class, has a profound impact on perceptions of income inequality. Income class gradient is observed in acceptance of income inequality, with higher income groups tending to perceive inequality more positively (Gijsberts, 2002; Szczepaniak et al., 2025) or, more generally, the positive relationship between affluence and acceptance of inequality is identified (Berlingieri et al., 2023; Corneo & Grüner, 2002; Czerniak et al., 2018; Haddon &Wu, 2022; Litwiński et al., 2023; OECD, 2021; Rueda & Stegmueller, 2019). Among other determinants of income inequality acceptance, the following are: higher education levels (Knell & Stix, 2020; Kuziemko et al., 2015), a sense of empowerment, conservative worldview, and support for redistribution (Litwiński et al., 2023), and age (Czerniak et al., 2018; Litwiński et al., 2023; Szczepaniak, 2025).

Additionally, behavioural-economic theories provide further insight into why individuals differ in their attitudes toward income inequality. Inequality aversion models (Fehr & Schmidt, 1999) posit that individuals experience disutility from unfair income differences and thus react not only to absolute income but also to the distributional context. Such preferences help explain why attitudes toward inequality may vary across income classes, with lower-income groups exhibiting a stronger aversion to unequal outcomes. At the same time, subjective evaluations of inequality are shaped by perceptions of relative income, as documented in extensive research showing that individuals assess their satisfaction through social comparisons rather than objective income levels (Cruces et al., 2013). This aligns with Social Comparison Theory (Festinger, 1954; Michalos, 1985), which posits that people evaluate their standing by benchmarking themselves against their peers, often normalising inequality within their reference groups. Moreover, System Justification Theory (Jost & Banaji, 1994) suggests that individuals may rationalise disparities to maintain psychological stability, contributing to the internalisation of inequality. Taken together, these behavioural and psychological frameworks explain why attitudes toward inequality and their links to trust may differ across income strata and how the perceived legitimacy of inequality emerges from cognitive and socio-economic processes.

A substantial body of evidence shows that higher income or social class is positively correlated with generalised trust (e.g., Alesina & La Ferrara, 2002; Ananyev & Guriev, 2018; Brückner et al., 2021; Navarro-Carrillo et al., 2018; Qiang et al., 2021). Qiang et al. (2021) argue that trust entails risk, which disproportionately affects lower-class individuals with fewer resources to absorb losses, whereas wealthier individuals face lower risk and can leverage trust for greater benefits. The pattern is not universal; Hamamura (2012) finds that while social class predicts generalised trust in wealthy countries, the association disappears in less affluent contexts. By contrast, the link between institutional trust and income is less consistent. Some studies report a positive correlation, but only in certain countries (Catterberg & Moreno, 2006) or specific regions within a country (Chen & Wang, 2022). Other research finds no significant association (Kaasa & Parts, 2008).

As previously noted, extensive research confirms a positive relationship between trust and income inequality. Bjørnskov (2006, pp. 5, 8) indicates that a more equal income distribution is reported to be beneficial for generalised trust in virtually all studies. The relationship between trust and income inequality is supported by others (Alesina & La Ferrara, 2002; Knack & Keefer, 1997; Knack & Zak, 2003; Kyriacou & Velásquez, 2015; Rothstein & Uslaner, 2005; Zak & Knack, 2001). Uddin (2025, p. 2) claims that research on the relationship between trust and inequality has produced mixed results; while some studies report a negative correlation, others identify no significant association.

Algan & Cahuc (2014) highlight the need to examine the direction of causality in the relationship. The negative correlation may reflect a high level of trust co-occurring with a greater preference for redistribution, which reduces inequality. Conversely, high inequality may reduce trust, as individuals feel unfairly treated by other social classes, leading them to limit trust to their own class. When societal rewards are inequitably distributed, individuals may perceive themselves as being exploited, leading to undermining trust and reducing social cohesion (Brehm & Rahn, 1997, p. 1009). Bergh & Bjørnskov (2014) indicate that the support for the causality, where inequality determines trust, is ambiguous, arguing that causality is bi-directional, with the effect from trust to inequality substantially stronger. Accordingly, trust remains a deeply embedded informal institution and relatively insensitive to deliberate change (O. E. Williamson, 2000).

Bergh and Bjørnskov (2014) argue that trust facilitates cooperation and, consequently, leads to more equitable outcomes. People trust those they perceive as similar. Lower inequality is associated with greater societal homogeneity, which encourages trust (Rothstein & Uslaner, 2005). In hierarchical societies, trust is often developed within small groups bringing together similar people (Kyriacou & Velásquez, 2015). Such trust may reflect solidarity and the belief that each group shares a common fate (Rothstein & Uslaner, 2005, p. 42). There is no single accepted explanation of the negative relationship, as multiple justifications exist in the literature that receive varied degrees of empirical support (Jordahl, 2007).

Institutional trust also shapes attitudes. Recent decline in public trust appears closely linked to economic factors. Foster and Frieden (2017) attribute this not directly to inequality, but to individuals' positions in the labour market, with those in more favourable positions reporting greater trust in government. Wroe (2016) finds that economic insecurity significantly reduces political trust. While inequality is negatively correlated with political trust, Zmerli and Castillo (2015) show that it operates through both objectively measured and subjectively perceived inequality. Bobzien (2023) highlights the significance of the gap between preferred and perceived inequality, with a widening gap lowering political trust. Similarly, Loveless (2013) links perceived excessive inequality to lower trust and political efficacy. Shuai et al. (2024) emphasise mediating mechanisms in the inequality-political trust relationship. Cross-border evidence further indicates that citizens in more unequal societies tend to express lower confidence in public institutions (Anderson & Singer, 2008). Economic inequality also undermines democratic attitudes, regardless of social class (Krieckhaus et al., 2013).

As discussed above, existing research on trust and inequality typically relies on objective income inequality measures and largely overlooks the relationships between trust and attitudes toward inequality. In the Polish context, given the widespread preference for income equality and low trust levels, studying the subjective inequality in link with trust provides essential insight into how both generalised and institutional trust may shape subjective evaluations of inequality in income distribution.

2. Data and methods

2.1. Data

The analysis is based on data from the 5th wave of the European Values Study (Gedeshi et al., 2020). The EVS is a large-scale, cross-national research programme that provides repeated measures of social attitudes, values, and beliefs across European countries. The 5th wave EVS data collection spanned from 2017 to 2021. The Polish subsample is nationally representative, covering adults aged 18 and older. The final sample used for analysis consists of 1,352 respondents from Poland.

2.2. Economic stratification

To examine patterns by income class, we use the EVS household income groups variable (Q98), in which respondents placed their household's net monthly income on a 10-point scale corresponding to the national income deciles (from the lowest to the highest). For analytical purposes, the original ten categories were consolidated into five income classes. Following approaches commonly applied in research on economic stratification and middle-class dynamics (OECD, 2019; Szczepaniak, 2024; Szczepaniak et al., 2025; Vaughan-Whitehead, 2016), we grouped the deciles as follows: low income (LIC, deciles 1–2), lower-middle (LMIC, 3–4), core middle (CMIC, 5–6), upper-middle (UMIC, 7–8), and high income classes (HIC, 9–10). The five-class structure aligns with stratification thresholds widely used in comparative research, providing sufficient granularity to examine attitudinal heterogeneity across income positions.

2.3. Key variables

Attitudes towards income inequality, as interpreted in our study, refer to how income inequality is perceived and are synonymous with perceptions of inequality (Szczepaniak et al., 2025). Attitudes toward income inequality are measured using variable Q32D, which captures agreement on a 10-point scale between the statements: “Incomes should be made more equal” and “There should be greater incentives for individual effort”. The dichotomised measure of income inequality attitudes used in our study follows the conceptual interpretation of the EVS item, where values 1–5 reflect support for reducing income differences (a negative attitude toward inequality), whereas values 6–10 reflect greater acceptance of income inequality (a positive attitude toward income inequality). Moreover, other studies of the perception of income inequality conceptually use binarisation, collapsing responses into two groups: egalitarians and incentive-oriented respondents (Berlingieri et al., 2023; Knell & Stix, 2020).

Variables related to trust are constructed using factor analysis (presented in Section 2.4), based on responses to two sets of questions: Q8, which measures trust in people from various social groups, and Q38, which assesses confidence in various organisations. Both use a four-point response scale. A detailed list of the trust items included in the factor analysis is presented in Table 1.

2.4. Factor analysis

From the EVS questions mentioned above, we selected those that assess trust, and grouped them into two categories: generalised trust and institutional trust. Data on generalised trust pertain to trust in specific groups of people. Proxies for institutional trust are derived from questions measuring respondents’ confidence in various organisations. Although Q38 comprises trust ratings for various organisations (e.g., the Church, the armed forces, and the press), not all items were included in the analysis. The selection of variables was aligned with the research aim of examining trust in the state's institutional structures in relation to perceptions of inequality. Therefore, only items reflecting key governance-related institutions (government, parliament, political parties, justice system, police, and civil service) were selected for further analysis.

We employed factor analysis to validate the selection of questions for the two identified trust categories. The necessary conditions for factor analysis (determinant value, KMO, and Bartlett's test) are met. This analysis allowed us to assess whether the questions were homogeneous and measured similar constructs. Additionally, it helped identify the internal structure of the scales and extract the underlying factors. The number of factors within each trust category was determined using the Kaiser criterion, which retains factors with eigenvalues greater than one. Varimax orthogonal rotation was applied to define the factors. To ensure the reliability of the scales and the appropriateness of the selected items, we also conducted Cronbach's alpha tests. It is generally accepted that a Cronbach's alpha value of at least 0.7 is necessary for a scale to be considered reliable. In our analysis, the Cronbach's alpha values are 0.777 for generalised trust and 0.786 for institutional trust. Generalised trust consists of two factors, which together explain 67% of the variance. Similarly, institutional trust comprises two factors explaining 66% of the variance.

For generalised trust, the first factor reflects trust in strangers, while the second captures trust in familiar individuals (relatives and acquaintances). For institutional trust, the first factor is primarily associated with political organisations, reflecting confidence in abstract or systemic actors that individuals typically perceive but do not interact with directly. The second factor represents trust in implementing organisations, those responsible for maintaining order and providing public services, perceived as less political (Table 1). This distinction aligns with the approach developed by Rothstein & Stolle (2008), who emphasise that people perceive institutions differently.

Table 1. Generalised and institutional trust

Generalised trust	Institutional trust
• Trust in strangers • Trust in people of another religion • Trust in people of another nationality • Trust in people you meet for the first time • Trust in relatives • Trust in your family • Trust in people in your neighbourhood • Trust in people you know personally	• Trust in the political organisations • Confidence in government • Confidence in parliament • Confidence in political parties • Trust in implementing organisations • Confidence in the justice system • Confidence in the police • Confidence in civil service

Source: based on data from EVS and results of factor analysis.

Based on the factor scores, we divided respondents into two groups. Respondents with factor values below zero were classified as having high trust, while those with values above zero were classified as having low trust. This division reflects the mean-centred nature of factor scores, where the average value is zero.

2.5. Statistical procedures

The analysis proceeds in the steps corresponding to the research questions. Primarily, for the first and second research questions, the procedure was as follows: (1) Descriptive statistics by income class; (2) Normality tests (Shapiro–Wilk); (3) Kruskal–Wallis test for differences across income classes (non-parametric due to ordinal scale and non-normal distribution); (4) *Post-hoc* Dunn's test with Bonferroni correction to identify class pairs with significant differences (Dunn, 1964).

Then, for the third research question procedure was: (1) Cross-tabulations of trust categories and inequality attitudes (overall sample and by income class); (2) Chi-square tests for associations; (3) Binary logistic regression: dependent variable is dichotomised inequality attitude (1–5 = low acceptance, 6–10 = high acceptance). Independent variables are trust subdimensions, entered separately and by income class to examine heterogeneous effects.

All analyses were performed in SPSS 28. Non-parametric tests were chosen due to the ordinal nature of variables and deviations from normality. Logistic regression models were assessed using the Hosmer–Lemeshow test for model fit.

3. Results

3.1. Attitudes towards income inequality by income classes

First, attitudes towards income inequality from the perspective of economic stratification were analysed. The results indicated a clear upward trend in the acceptance of income inequality across income classes. Mean scores on the attitudes toward inequality increase progressively from the lowest to the highest income class, suggesting that individuals in higher income classes are more likely to view income disparities as acceptable (Figure 1).

Following this, a normality test was conducted. The normality of attitudes towards income inequality was assessed using the Shapiro–Wilk test.

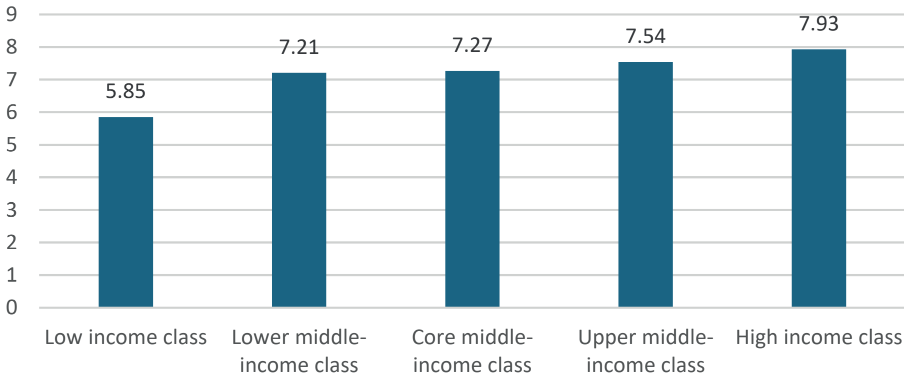

Figure 1. Attitudes towards income inequality by income classes

Source: based on data from EVS.

Significant deviations from normality were detected for attitudes towards income inequality (Shapiro–Wilk test statistic = 0.862; $p < 0.001$). Accordingly, non-parametric statistical methods were used for subsequent analyses.

A Kruskal–Wallis test was conducted to examine whether individuals from different income classes exhibit significantly different attitudes toward income inequality. The test revealed a statistically significant difference in the distribution of attitudes toward income inequality across income classes

Table 2. Pairwise comparisons of income classes regarding attitudes towards income inequalities

Income class 1–Income class 2	Test statistic	Standard error	Significance	Adjusted significance
LIC vs. LMIC	–40.876	26.889	0.128	1.000
LIC vs. CMIC	–55.133	29.022	0.057	0.575
LIC vs. UMIC	–70.998	29.307	0.015	0.154
LIC vs. HIC	–123.355	31.047	<0.001	0.001
LMIC vs. CMIC	–14.257	28.416	0.616	1.000
LMIC vs. UMIC	–30.122	28.707	0.294	1.000
LMIC vs. HIC	–82.479	30.481	0.007	0.068
CMIC vs. UMIC	–15.865	30.714	0.605	1.000
CMIC vs. HIC	–68.222	32.378	0.035	0.351
UMIC vs. HIC	–52.357	32.634	0.109	1.000

Note: Asymptotic significances (2-sided tests) are displayed. The significance level is 0.05.
Each row tests the null hypothesis that the distributions of Samples 1 and 2 are the same.
* Adjusted significance – significance values have been adjusted using the Bonferroni correction for multiple tests.

Source: based on data from EVS.

(H(4) = 16.972, p = 0.002; N = 1076). Moreover, the Chi-square test revealed a statistically significant association between income class and attitudes toward income inequality (χ^2(36) = 68.334, p < 0.001). These findings corroborate descriptive analyses, indicating that income class plays a meaningful role in shaping attitudes toward income inequality. Individuals from different income brackets do not share the same level of acceptance of inequality, with higher income classes showing greater acceptance.

In the next step, post-hoc pairwise comparisons were conducted using Dunn's test with Bonferroni correction to examine the differences between selected income classes and attitudes towards income inequality (Table 2).

There is robust evidence that HIC members exhibit significantly greater acceptance of income inequality than members of the LIC. However, the differences between other income classes remained insignificant, suggesting homogeneity between middle-income class subgroups in attitudes towards income inequalities.

3.2. Relationships between different categories of trust and income classes

Based on the factor analysis (Section 2.4), we distinguished four categories of trust: trust in strangers, relatives, political organisations, and implementing organisations. In this analysis stage, the mean values of trust by income classes were analysed and are presented in Figure 2. The results revealed nuanced patterns depending on the trust categories. Trust in strangers increases

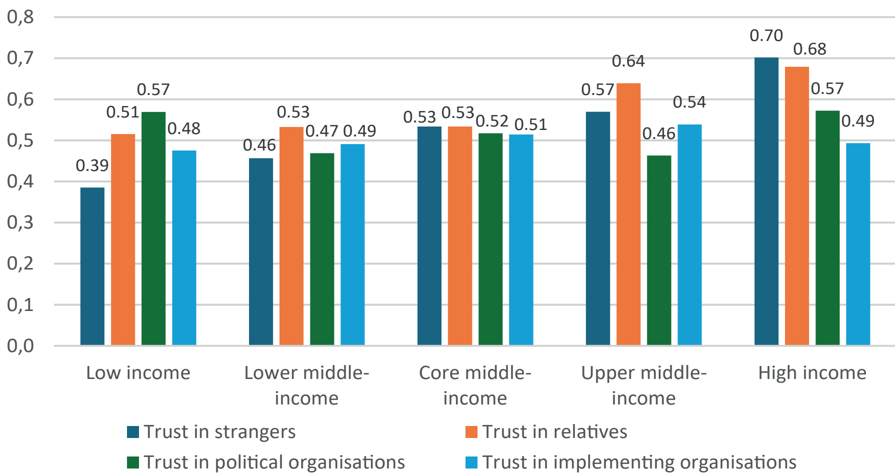

Figure 2. Mean trust (in four categories) by income class

Source: based on data from EVS.

consistently with income class, suggesting that this kind of generalised trust may be strongly tied to economic advantage. Trust in relatives also shows an upward trend, though less steeply. Higher income may facilitate stronger or more secure social ties. Trust in political organisations, in contrast, displays a non-linear pattern: higher in low- and high-income groups, but dips in the middle classes. This may reflect varying institutional scepticism or political alignment across classes. Lastly, trust in implementing organisations increases up to the upper-middle class, then drops in the highest class, indicating a non-linear relationship. This pattern suggests a nuanced or critical stance among the most affluent.

Normality of the trust categories was assessed using the Shapiro–Wilk test. Significant deviations from normality were detected for all categories of trust (Shapiro–Wilk test statistics equalled 0.946 ($p < 0.001$), 0.941 ($p < 0.001$), 0.977 ($p < 0.001$), 0.995 ($p = 0.004$) for trust towards strangers, relatives, political, and implementing organisations, respectively). Accordingly, non-parametric statistical methods were used for subsequent analyses. A Kruskal–Wallis test was conducted to examine whether individuals from different income classes exhibit significantly different levels of trust. The results, presented in Table 3, reveal that a statistically significant difference in the distribution of trust across income classes exists for trust in strangers, trust in others, trust in political organisations, but not trust in implementing organisations.

Table 3. Kruskal–Wallis test results for trust categories across income classes

	Trust categories			
	Generalised trust		Institutional trust	
	trust in strangers	trust in relatives	trust in political organisations	trust in implementing organisations
Test statistic	36.619	14.929	11.041	1.741
p-value	<0.001	0.005	0.026	0.783

Source: Own preparation based on data from EVS.

To conduct an in-depth analysis, a pairwise comparison was conducted to examine the differences between selected income classes and trust categories, where the differences were statistically significant. *Post-hoc* pairwise comparisons were performed using Dunn’s test with Bonferroni correction. The results revealed the significant differences between the low and high-income classes in three categories of trust. Between the HIC and the LMIC, significant differences were identified in trust in strangers and political organisations. The most sensitive to class differences was trust of strangers, with significant differences also between the middle parts of the income distribution (Table 4).

Table 4. Pairwise comparisons of trust towards strangers, relatives, and political organisations by income classes

Income class 1–Income class 2	Trust in strangers		Trust in relatives		Trust in political organisations	
	test statistic	adjusted signifi-cance	test statistic	adjusted signifi-cance	test statistic	adjusted signifi-cance
HIC vs. UMIC	76.814	0.088	10.129	1.000	–44.247	0.848
HIC vs. CMIC	89.166	0.022	50.822	0.634	–68.147	0.175
HIC vs. LMIC	126.521	0.000	63.563	0.290	–83.162	0.024
HIC vs. LIC	165.967	0.000	78.932	0.049	–87.106	0.017
UMIC vs. CMIC	12.352	1.000	40.692	1.000	–23.899	1.000
UMIC vs. LMIC	49.707	0.564	53.434	0.552	–38.915	1.000
UMIC vs. LIC	89.153	0.009	68.803	0.103	–42.859	1.000
CMIC vs. LMIC	37.355	1.000	–12.741	1.000	15.015	1.000
CMIC vs. LIC	76.801	0.039	28.111	1.000	18.959	1.000
LMIC vs. LIC	39.446	1.000	15.369	1.000	–3.944	1.000

Note: Asymptotic significances (2-sided tests) are displayed. The significance level is 0.050.
Each row tests the null hypothesis that the distributions of Samples 1 and 2 are the same.
* Adjusted significance – significance values have been adjusted using the Bonferroni correction for multiple tests.

Source: based on data from EVS.

To conclude, the analysis revealed that trust is not uniform across social strata. Higher-income individuals are generally more trusting of others on a general level (strangers and relatives) but demonstrate greater scepticism toward institutional actors, especially political bodies.

3.3. Relationships between trust and attitudes towards income inequality

To explore this relationship, cross-tabulations and logistic regression models were conducted to assess whether higher levels of trust increase the chance of a more positive attitude toward income inequality, both in the overall sample and within each income class.

In the whole sample, higher trust in strangers was significantly associated with greater acceptance of income inequality. This suggests that trust in strangers may play a legitimising role in shaping more favourable views of inequality. Moreover, other categories of trust, such as trust in relatives, political organisations, or implementing organisations, did not exhibit a significant association with attitudes toward inequality for all classes combined. Therefore, given that both attitudes toward income inequality and most categories of trust differed significantly across income classes, an analysis of the relations between attitudes toward income inequality and trust was conducted in terms of income class stratification (Table 5).

In the LIC, none of the trust dimensions was significantly related to attitudes about inequality. In both LMICs and UMICs, a significant association emerged between trust in relatives and the acceptance of inequality. In the CMIC, none of the trust variables showed a significant association with inequality attitudes. Finally, no significant relationship was found between any trust dimension and attitudes towards inequality in the HIC. To conclude, the relationship between trust and acceptance of inequality is complex and income-class specific, underscoring the need to consider class-specific social orientations when analysing the moral and psychological foundations of inequality's legitimacy.

In the final step of the analysis, to assess whether higher trust influences the chance of accepting income inequality, binary logistic regressions were performed using trust categories as predictors and a dichotomised version of the income inequality attitude variable as the dependent variable. Analyses were conducted on the full sample and disaggregated by income class to capture possible class-specific effects (Table 6).

The results from the full-sample model indicate that higher trust in strangers significantly increases the chance of expressing a more positive attitude toward income inequality. This supports the idea that this kind of generalised trust may function as a legitimising mechanism for the social and economic order. No other trust categories (trust in relatives, political organisations, or implementing organisations) significantly affected the probability of higher acceptance of income inequality (Table 6).

Remarkably, the class-stratified models reveal important heterogeneity in the effect of trust. In the LIC model, only trust in strangers showed a weakly significant positive association with acceptance of inequality; other trust categories were not significant. In the LMIC model, a significant relationship was found between trust in relatives and the acceptance of inequality. Individuals in this group who expressed a high level of trust in their close network were less likely to accept income inequality. However, those who expressed great trust in strangers were more likely to accept inequality to a greater extent (weak significance). This may suggest that interpersonal solidarity and proximity reinforce egalitarian values in this class. In the CMIC model, no trust variable was significantly associated with attitudes toward inequality. This neutrality may reflect a more ambiguous position in the social structure, where both upward and downward identifications are possible. In the UMIC, in con-

Table 5. Attitudes towards income inequality and trust dimension in all classes combined and by income class (Chi-squared)

Relations between income classes and trust categories	Income class											
	All classes combined		LIC		LMIC		CMIC		UMIC		HIC	
	test statistic (χ^2)	p-value	test statistic (χ^2)	p-value	test statistic (χ^2)	p-value	test statistic (χ^2)	p-value	test statistic (χ^2)	p-value	test statistic (χ^2)	p-value
Trust in strangers	26.987	0.001	10.368	0.322	15.220	0.085	10.950	0.279	10.928	0.281	8.320	0.502
Trust in relatives	9.348	0.406	5.119	0.824	20.687	0.014	4.743	0.856	21.580	0.010	4.686	0.861
Trust in political organisations	4.592	0.868	6.895	0.648	15.137	0.087	3.527	0.940	10.393	0.320	9.974	0.353
Trust in implementing organisations	13.611	0.137	13.390	0.146	4.035	0.909	8.334	0.501	4.55	0.871	9.84	0.367

Source: based on data from EVS.

Table 6. Results of the binary logistic regression models of all income classes and by income class

	All income classes combined. N = 1352		Income classes									
			LIC (N = 255)		LMIC (N = 274)		CMIC (N = 202)		UMIC (N = 195)		HIC (N = 160)	
	B (p-value)	Exp (B)	B (p-value)	Exp (B)	B (p-value)	Exp (B)	B (p-value)	Exp (B)	B (p-value)	Exp (B)	B (p-value)	Exp (B)
Trust in strangers (1)	0.524 (<0.001)	1.688	0.698 (0.062)	2.009	0.646 (0.063)	1.908	–0.118 (0.772)	0.889	–0.752 (0.147)	0.472	0.253 (0.669)	1.288
Trust in relatives (1)	–0.100 (0.536)	0.905	0.034 (0.924	1.034	–0.907 (0.012)	0.404	0.326 (0.425)	1.386	1.103 (0.029)	3.012	0.016 (0.978)	1.017
Trust in political organisations (1)	0.024 (0.879)	1.024	–0.037 (0.918)	0.964	0.046 (0.894	1.047	–0.112 (0.783)	0.894	–0.187 (0.688)	0.829	–0.020 (0.972)	0.980
Trust in implementing organisations (1)	0.190 (0.231)	1.209	0.468 (0.189)	1.597	–0.277 (0.422	0.758	0.054 (0.895)	1.055	0.570 (0.229)	1.768	0.206 (0.719)	1.229
Hosmer and Lemeshow Test	Chi-square	**p-value**	Chi-square	**p-value**	Chi-square	**p-value**	Chi-square	**p-value**	Chi-square	**p-value**	Chi-square	**p-value**
	4.349	0.824	10.979	0.203	5.051	0.752	1.240	0.996	2.807	0.946	6.913	0.546

Source: based on data from EVS.

trast to the lower middle class, trust in relatives was positively associated with acceptance of inequality. Individuals with stronger bonds to close networks were more likely to express favourable attitudes toward inequality, possibly reflecting shared values of meritocracy or individual responsibility. Finally, in the HIC model, none of the trust categories were significantly related to attitudes about inequality. This suggests that within the most affluent group, acceptance of inequality may be broadly internalised and no longer contingent on generalised or institutional trust levels.

The internalisation of inequality can be understood through several psychological and behavioural mechanisms. Research on system justification (Jost & Banaji, 1994) shows that individuals are motivated to perceive existing social arrangements as fair and legitimate, which leads them to rationalise income disparities even when these are unfavourable to them. Another mechanism relevant to this process is related to the Social Comparison Theory (Festinger, 1954), which posits that individuals evaluate their socio-economic position by comparing themselves with similar peers. Such intragroup social comparisons often lead to the normalisation of income differences, particularly when people assess their circumstances relative to peers with comparable incomes or social status. As a result, inequalities may come to be perceived as natural or deserved, reflecting a psychological internalisation of inequality.

The results from logistic regressions suggest that trust influences the legitimisation of inequality, but in class-dependent ways. Generalised trust (in strangers) plays a legitimising role in the full sample. However, generalised trust (in relatives) shows a bifurcated effect, discouraging acceptance of inequality among LMIC while promoting it among the UMIC. These findings support a differentiated model of inequality legitimation, in which both the object of trust and the economic context, in terms of income class, shape individuals' attitudes.

4. Discussion

This study contributes to the literature by connecting subjective inequality with multidimensional trust across income classes in a post-socialist context, an approach that has been rarely taken in research thus far. The findings are thought-provoking, which underscores the complexity of these relationships. Three central findings emerge from the analysis.

Firstly, attitudes toward income inequality differ significantly across income classes, with the sharpest contrasts observed between high- and low-income groups, while the middle-income strata display relatively homogeneous views. Higher-income classes tend to be more accepting of inequality, whereas lower-income classes express rather negative perceptions. Our results indicating an income-class gradient in positive attitudes to inequality are consistent with Corneo and Grüner (2002), Haddon and Wu (2022), Litwiński et al. (2023), Szczepaniak (2025), and Szczepaniak et al. (2025). It should be noted, however, that while Corneo and Grüner (2002) primarily analyse preferences for redistribution as an indirect indicator of inequality acceptance, and Haddon and Wu (2022) examine perceptions of inequality in relation to class and contextual levels of actual inequality, their findings nonetheless align with the general pattern of greater acceptance of inequality among higher-income groups.

The behavioural and psychological frameworks outlined in the literature help interpret the class-specific patterns from our results. Inequality aversion models suggest that lower-income groups are more sensitive to unfair distributional outcomes, consistent with their stronger opposition to income inequality. Conversely, the greater acceptance of inequality among upper-income groups can be understood through status-based preferences and system-justifying tendencies that legitimise existing disparities. Together, these mechanisms clarify how economic stratification influences the relationship between trust and attitudes toward inequality, reinforcing the correlational patterns observed empirically.

Secondly, trust varies systematically by income class, but the pattern depends on the type of trust. Generalised trust increases consistently with income, reflecting higher perceived societal security among higher-income groups. This notion is consistent with other research (Alesina & La Ferrara, 2002; Ananyev & Guriev, 2018; Brückner et al., 2021; Navarro-Carrillo et al., 2018; Qiang et al., 2021). In contrast, institutional trust is more nuanced: trust in political organisations dips in middle-income classes and partially recovers among high-income groups, while trust in implementing organisations peaks in the upper-middle class and declines among the wealthiest. Results regarding institutional trust are mixed in other studies as well. Catterberg & Moreno (2006) demonstrate that income increases institutional trust in Eastern Europe and Latin America, but has no effect in the former Soviet republics. Medve-Bálint & Boda (2014) indicate a different pattern for Western and East-Central Europe, with a negative relationship in the latter region. Chen & Wang (2022) show that while public trust is positively associated with income in China, this relationship only holds in some regions.

The study reveals a clear divergence between generalised and institutional trust, underscoring the importance of examining trust from multiple perspectives. Trust should be disaggregated into distinct categories, as they may have unique determinants and consequences. Furthermore, the analyses should move beyond a single concept, such as generalised trust, and account for distinct forms, sources, consequences, and interrelations of trust. The findings have crucial implications for redistribution policy: trust, income inequality, and the welfare state are linked through complex relationships (Bergh &

Bjørnskov, 2014). Research confirms that high levels of generalised and institutional trust foster support for redistributive policies (Bergh & Bjørnskov, 2011; Bjørnskov, 2006; Daniele & Geys, 2015). Therefore, trust differences across income groups may not only shape inequality perceptions but also the political feasibility of policies designed to address it.

Thirdly, the relationship identified between trust and attitudes toward income inequality in Poland is complex and heterogeneous. At the aggregate level, trust in strangers is associated with greater acceptance of inequality, suggesting a legitimising role (Rothstein & Uslaner, 2005; Zmerli & Castillo, 2015). However, class-specific results reveal contrasting effects: in the LMIC, greater trust in relatives corresponds to lower acceptance of inequality, indicating solidarity-driven egalitarian preferences, while in the UMIC, the same trust dimension correlates with greater acceptance, possibly reflecting meritocratic beliefs. In the HIC, no significant relationships emerge, suggesting normalised acceptance of inequality regardless of trust levels. Trust in strangers appears to support the acceptance of inequality, both in the general population and among lower-income groups.

In summary, trust patterns vary significantly across income groups. Policymakers can leverage these trust profiles to design more inclusive and effective policies and to tailor communication to the specific income class's expectations.

Conclusions

This article contributes to the literature on trust and subjective inequality by examining their relationship across income classes in Poland. The results show that acceptance of inequality increases with income strata. Moreover, while generalised trust rises with income classes, institutional trust follows more complex patterns. The link between trust and inequality attitudes is heterogeneous, with different types of trust playing distinct roles in shaping acceptance in specific income classes. These findings underscore that trust attitudes toward inequality relationships are mediated by economic stratification.

Those insights have important policy implications. Efforts to reduce inequality should address not only material disparities but also public perceptions of these disparities. In this context, media and political narratives about inequality matter, as they shape how individuals interpret economic inequality and shape trust. As Fukuyama (1997, p. 180) observed, the economic value of trust becomes most apparent when we consider the dysfunction arising in its absence. This reinforces the urgency of safeguarding trust as a key societal resource amid persistent negative perceptions of inequality. Building trust is not straightforward. Although the origins of trust remain debated (e.g., Bentkowska, 2023; Mishler & Rose, 2001), the literature suggests that trust is strengthened when institutions demonstrate competence, fairness, and responsiveness. Accordingly, policies addressing inequality may be more effective when they combine material measures with credible signals of commitment, including transparency, clear priorities, and consistent implementation. Participatory mechanisms may further reduce psychological distance and reinforce perceptions of institutional reliability, increasing the likelihood that trust is sustained over time.

This study focuses on Poland, a post-socialist country characterised by a common belief about excessive income inequality and relatively low levels of trust. While this context offers insights into the relationship between trust attitudes and inequality, it also limits generalisability due to the single-country design. Although Poland can serve as a benchmark for EU-based post-socialist states, future research would benefit from cross-country analyses and from examining these relationships over time.

Although our findings reveal systematic associations between trust and attitudes toward income inequality across income classes, these relationships are non-causal. The use of EVS data limits our ability to determine the direction of influence between the variables. Our results, therefore, reflect correlational patterns that may arise from multiple underlying mechanisms.

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Financial inclusion and economic growth in Vietnam: Evidence across provinces and income groups

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Abstract

This research aims to examine the effect of financial inclusion on economic growth in Vietnam. Using panel data from 63 provinces during 2014–2020, estimations are conducted for both the full sample and across two income groups. Financial inclusion is measured by indicators capturing geographical penetration and using products and services in commercial banks and insurance. The difference-GMM estimation results demonstrate that financial inclusion captured by higher commercial bank branches and using bank accounts, saving passbooks, and ATM cards present significant positive effects on economic growth in Vietnam. In contrast, participating life and non-life insurance shows a non-significant effect. For high-income provinces, participating in life and non-life insurance positively affects economic growth. In addition, the study indicates robust

Keywords

- financial inclusion
- economic growth
- Vietnam

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effects of commercial bank branch penetration and using ATM cards in enhancing economic growth in both low-income and high-income localities.

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Introduction

The growing body of literature demonstrates the critical role of financial inclusion as a cornerstone for economic growth. Financial inclusion is defined as the provision of affordable financial products and services to all individuals and firms that require them (Sarma, 2008; Sethi & Acharya, 2018). Wider access to financial system fosters greater participation of individuals and firms in savings and credit products that increase the money multiplier and provide a variety of capital resources for technology innovation, thus spurring economic development (Siddiki & Bala-Keffi, 2024). Moreover, through offering a range of financial services and financial applications using web-based portals and mobile phones, financial inclusion makes conducting financial transactions more convenient, thus easing trading activities and creating higher national income (Creane et al., 2004). The benefits of financial inclusion contribute to the economy also in the insurance sector when customers are protected by policies of insurance compensation and risk management that reduce economic loss and promote business activities, stimulating economic expansion (Apergis & Poufinas, 2020). The majority of empirical studies support the claim that financial inclusion positively impacts economic growth (N. Khan et al., 2022; Osuma, 2025; Siddiki & Bala-Keffi, 2024; Singh & Mallick, 2024; Wibowo et al., 2023). However, not only a positive linkage is documented in previous research. Other studies document negative effects (Pal et al., 2025), a U-shaped relation (Sahay et al., 2015), and a bi-directional causality linkage (Sethi & Acharya, 2018).

The existing literature also highlights how financial inclusion might impact on economic growth to different degrees. In particular, the magnitude and direction of this relation in developing countries is more pronounced than developed countries (Hussain et al., 2024; Narain et al., 2022; Sethi & Acharya, 2018). Furthermore, there may be differences between the short-term and long-term results (N. Khan et al., 2022), and the effect may be contingent on the financial system background and the level of financial development (Z. Chen et al., 2023; Siddiki & Bala-Keffi, 2024).

While there is a large cohort of studies examining the linkage in other developing countries (Hussain et al., 2024; Narain et al., 2022; Pal et al., 2025), this research topic has largely been ignored for Vietnam. Due to its remarkable economic transformation over the past decades, Vietnam presents a compelling context for studying financial inclusion. As a transition country, Vietnam has been developing a market-oriented economy that shifted from a centrally planned system since 1986. This has resulted in rapid growth, enabling Vietnam to develop from being one of the world's poorest nations to a lower middle-income country. Financial inclusion is recognised as a key driver of such economic development. The Global Findex Database (2021) indicates that the percentage of people aged above 15 who possess a financial institution account was 56% in 2021, representing a sharp increase from 21% in 2011. However, this level is still low compared to the world average of approximately 74% (Global Findex Database, 2021). Given this context, Vietnam provides a relevant case for examining the impact of financial inclusion on economic growth.

The current literature on Vietnam is limited on analysing how demographic and governance factors determine financial inclusion (H. S. Nguyen et al., 2023; T. T. H. Nguyen & Luong, 2023; H. S. Tran et al., 2019; T. Q. Tran & Dinh, 2021). Also, research on the correlation with economic growth mainly uses national aggregate data by time series (Hye, 2022; Tran, 2023). Meanwhile, it is more relevant to observe this connection at the disaggregate level by provincial units, since panel data estimation can account for the unobservable fixed socio-economic characteristics of each province. In addition, relative differences exist among aspects of financial inclusion in Vietnam when comparing the 56% of respondents possessing a financial institutions account to the 40% of the population with savings in financial institutions and the fraction of population with a mobile money account (16%). Therefore, prior research using an aggregate index cannot examine how different aspects of inclusion in the financial system shape economic growth. Moreover, while provincial income levels have been shown to influence access to finance in Vietnam (N. T. Nguyen et al., 2021), the effect of financial inclusion on economic growth across provinces within different income classes remains open.

Accordingly, a number of issues remain unresolved in the literature in the subject, and this research aims to address these limitations and contribute to current literature in a few respects. Firstly, the study assesses the impact of financial inclusion on economic growth in Vietnam by using panel data from 63 provinces which have not yet been studied. Secondly, we consider financial inclusion in a more nuanced way by employing a range of indicators which capture various aspects of financial inclusion, from providing geographical penetration to using banking and insurance services. By doing so, this study can identify specific channels of financial inclusion's impact on economic growth. Thirdly, we investigate this relationship in two different income groups. In addition to observing the nexus for the whole group, the study classifies 63 provinces into two groups–high-income provinces and low-income provinces. The empirical model is estimated by the difference Generalised Method of Moments (difference-GMM). The robustness check employs system-GMM.

The remainder of this paper is structured as follows: Section 1 presents theoretical concepts and empirical literature. Section 2 focuses on methodology. This is followed by a results and discussion section and the last section presents concluding remarks, policy recommendations and suggestions for further studies.

1. Literature review

Financial inclusion is considered a wide concept with several ways to define it. Focusing on poor individuals, low-income households, and microenterprises, the Asian Development Bank (2000) states that inclusive finance provides these kinds of customers with a broad range of financial services, such as deposits, loans, payment services, and insurance products. The United Nations (2006) captures financial inclusion in three aspects—credits, savings and insurance. These allow access to credit sources, savings, payment services as well as insurance for all individuals and firms. Financial inclusion thus emphasises the ability more than the eligibility in terms of using financial products and services. Meanwhile, the World Bank (2008) highlights the price aspect in capturing financial inclusion, which offers non-price barriers in accessing financial services. Generally, this study defines financial inclusion as a phenomenon that makes financial products and services available to all individuals and firms at an affordable cost (Sarma, 2008; Sethi & Acharya, 2018). Moreover, financial inclusion is expressed in a broad concept including both geographical penetration and using products and services in the financial system.

Theoretical studies have highlighted financial inclusion's influence on economic growth through the activities of the financial system, which focuses on two channels—capital accumulation and technological innovation, represented by functions of financial markets and intermediaries. Levine (1997) states that the existence of market friction, including information costs and transaction costs, motivates the financial system's activities in order to meet business and individual demand. In particular, a greater number of individuals included in the financial system through savings products could mobilise savings. This could create more capital resources and make the production sector more efficient, thereby boosting economic development (Crane et al., 1995; Siddiki & Bala-Keffi, 2024). In addition, financial inclusion is presented by individuals' participation in financial instruments that provide capital for technological innovation. Meanwhile, financial intermediaries and financial markets support this process and play a better role in selecting profitable firms and managers that would induce more efficient capital allocation and boost national outcomes (Kodan & Chhikara, 2013). Savings mobilisation and the consequent support for capital accumulation and technological production, along with inclusive finance, allows individuals and firms to conduct financial transactions at lower cost, which in turn eases the trade of goods, services, and contracts, facilitating economic activities thus increasing economic growth (Ang, 2011). Moreover, when individuals are included in the financial system through hedging products related to the insurance sector, financial inclusion might contribute to economic growth through facilitating risk management and offering corporate controls. That could not only attract capital to the economy but also promote technology innovation projects, which require significant capital and risk protection, thus channelling higher national prosperity (Apergis & Poufinas, 2020).

Financial inclusion is documented as contributing to economic growth in a large number of empirical studies. The majority of these have supported the positive correlation between financial inclusion and economic growth. Focusing on 153 countries from 2011 to 2020, Siddiki and Bala-Keffi (2024) present a consistently positive relationship between inclusive finance and economic development. The magnitude of this relationship might differ across countries. Z. Chen et al. (2023) suggests that greater access to the financial system spurred economic growth in the 10 highest financially inclusive economies. However, the effect would not be same in different quantiles in these countries in the period of 2004–2018. Also, N. Khan et al. (2022) established a positive long-run nexus for 15 developed and emerging economies over the period 2004–2017, along with a non-significant effect in the short run. Additionally, Hussain et al. (2024) not only demonstrate the positive impact of financial inclusion on economic growth in 21 Asian countries over a period from 2004 to 2019, but also posit that developing countries experience more pronounced effects compared to advanced countries. Similarly, studies exist that indicate how financial inclusion positively affects economic growth in other groups of countries, such as the 55 countries of the Organization of Islamic Cooperation (OIC) from 1991 to 2015 (D. W. Kim et al., 2018), transition economies of the European Union (Bayar & Gavriletea, 2018), European Union member states (EU), and OECD members (Huang et al., 2021; J. H. Kim, 2016), as well as middle-income countries (Narain et al., 2022) and highest-emitting countries (Usman et al., 2021). Accordingly, the influence of financial inclusion on economic growth at different levels of the economy might result in distinct effects.

In this sense, some studies have documented a causality nexus, a U-shaped relationship, and negative effects. Sethi and Acharya (2018) used a panel of 31 developed and developing countries from 2004–2010 and identified a long-term positive relation, in addition to a bi-directional connection between financial inclusion and economics growth. Capturing traditional and digital inclusive finance, Pal et al. (2025) observed such a link in 23 emerging countries from 1990 to 2022. This study shows the positive effect of traditional financial inclusion on economic growth, while the effect of digital financial inclusion is found to be negative in the long run. Sahay et al. (2015) pointed out the U-shaped correlation between increasing the level of financial inclusion and boosting economic growth up to a dynamic threshold, after which remains stable or becomes negative in some advanced countries. Moreover, financial inclusion was found to have no effect on economic growth in a sense that it does not reduce the poverty level (Donou-Adonsou & Sylwester, 2016; Sukmana & Ibrahim, 2018).

Several kinds of relations between financial inclusion and economic growth are noted in existing research, they are mainly observed panels of countries. For a single country, China, Ahmad et al. (2021) observed how financial inclusion was linked to economic growth for 31 provinces in the period from 2011 to 2018. This study pointed out that increased use of bank accounts, payment accounts and other services such as savings, credit, insurance, and investment instruments enhances economic growth. Using national aggregate data, Dahiya and Kumar (2020) explored the effect of financial inclusion on economic growth in India form 2005 to 2017. Financial inclusion is captured in three dimensions: accessibility, penetration and usage. However, only use exhibited positive links with economic growth. Similarly, Odame et al. (2024) examined the financial inclusion and economic growth nexus in Ghana from 2005 to 2016. Financial inclusion in this research encompasses financial access, usage of financial services, and penetration of financial services. Only the use of financial services produced a positive effect on economic growth, while a negative impact was found for financial access and penetration of financial services.

In Vietnamese studies, limited aspects of financial inclusion have been explored. Most studies focused on examining factors affecting financial inclusion in Vietnam and used time series data to investigate financial inclusion in the national aggregate level. T. A. N. Nguyen and Luong (2023) used a household survey for the 2014–2018 period to identify key determinants of financial inclusion in Vietnam's households, which included total income, relative income and distance to bank branches. Whereas total income facilitated households' entry into the financial system, relative income reduced the level of financial inclusion and the distance to bank branches presented an unclear effect on financial inclusion. Focusing on how demographic characteristics affect financial inclusion, H. S. Tran et al. (2019) showed that income, age, and education positively influence the use of official accounts, savings while gender was not associated with and education had a negative effect on the use of credit. As regards the effect of financial inclusion on the economy, a few studies have documented how financial inclusion could reduce the level of multidimensional poverty (H. T. T. Tran et al., 2022) as well as increasing customer loyalty in Vietnam's banking sector. In the linkage with economic growth, Hye (2022) utilised time series data spanning from 2001 to 2021. This research calculated an aggregate index for financial inclusion in Vietnam, demonstrating that inclusive finance could promote economic growth. However, the study faced limitations in using time-series data at the national level that ignores heterogeneity in the socio-economic conditions of Vietnam's provinces. Also, financial inclusion was only measured by the overall index, and this cannot enable particular aspects of financial system to be explored in relation to economic growth.

In the light of the above-mentioned, studies on the financial inclusion-economic growth nexus have attracted considerable attention among researchers. Yet ambiguous effects remain and their different magnitudes depend on region / country-specific circumstances. Moreover, there are limitations in current studies regarding Vietnam, with few studies using provincial units to focus on the financial inclusion and economic growth relation. This study aims to fill the gap in the previous literature by studying whether financial inclusion affects economic growth in Vietnam using data at the provincial level.

2. Data and methodology

The study uses annual data for 63 provinces in Vietnam from 2014 to 2020. The full list of provinces is provided in the Appendix. The time period chosen here represents the stage during which Vietnam experienced the most significant expansion of inclusive finance. It is also determined by data availability. Additionally, this study classifies 63 provinces into two groups: high-income and low-income provinces. Specifically, the 20 high-income provinces are those with a gross provincial product (GPP) per capita above the national average, while the remaining 43 low-income provinces have a GPP per capita below the national average.

Economic growth is measured using GPP per capita expressed in US dollars. The research proxies financial inclusion through five indicators which cover inclusive finance in both the banking sector and insurance sector to present a broad concept of financial inclusion. Instead of presenting financial inclusion as an aggregate index as in previous studies (Pal et al., 2025; Sharma, 2016), financial inclusion is captured through five separate variables to assess the impact of each dimension on economic growth. The first variable is the number of commercial bank branches (FI1). It captures the geographical penetration of commercial bank branches and is widely used in the literature. A higher number of commercial bank branches provides greater accessibility to banking services (N. Khan et al., 2022; Pal et al., 2025). This study also employs micro-level data from a household survey in Vietnam. This survey captures inclusive finance by asking whether households currently use financial products and services such as banking accounts, savings, ATMs, and insurance. Based on the survey data, we calculate the percentage of households included in the financial system in each province. The second variable (FI2) is the share of households with a bank account. This indicator presents the inclusion of households in the financial system by means of a formal account, facilitating the utilisation of other financial services (Siddiki & Bala-Keffi, 2024). The third variable, financial inclusion is measured as having a savings passbook (FI3) that shows the involvement of households in capital mobilisation (Sharma, 2016). Fourth, the study employs the percentage of households using ATM cards as a measure of financial inclusion (FI4). Conducting financial transactions through the use of ATM cards demonstrates engagement in financial products and services (Singh & Mallick, 2024). Also, financial inclusion in this paper is measured by a household's participation in life insurance and non-life insurance (FI5). Overall, this study measures financial inclusion by indicators referring to two main sectors of the financial system in Vietnam: commercial banks (indicators FI1–FI4) and the insurance sector (FI5).

In addition to financial inclusion, economic growth is driven by numerous other factors. This study observes macroeconomic factors within traditional economic growth theory as control variables in the model: inflation, labour, foreign direct investment, urban population, and poverty. Inflation presents the effects of price stability on investment incentives and consumption patterns (Fischer, 1993; Sequeira, 2021), while labour reflects the availability of human capital, which is a fundamental determinant of productive capacity and innovation (Haudi et al., 2020; Lucas, 1988). FDI serves as an indicator of capital inflows, technology transfer, and integration into global value chains (Borensztein et al., 1998; Mehic et al., 2013). Urban population is included as a proxy for the level of urbanisation, often associated with improved infrastructure, market expansion, and productivity gains (H. Chen et al., 2017). Poverty accounts for socio-economic constraints that can suppress aggregate demand, human capital accumulation, and growth (Breunig & Majeed, 2020). It is expected that labour, FDI, and urban population positively impact on economic growth, while inflation and poverty rate have negative effects.

A description of all variables is presented in Table 1. Table 2 shows the descriptive statistics for 63 provinces in Vietnam, and the characteristics across the two income groups (low-income and high-income provinces) are included in Table 3. Table 2 indicates the variation of all economic variables across low-income and high-income provinces. The GPP per capita in high-income provinces ($4,024) is more than double that in low-income provinces ($1,778). A substantial gap is also observed in the financial inclusion indicators (FI1–FI4). Similarly, the percentages of FDI inflows and urban population in high-income provinces are almost double those of low-income provinces.

Table 1. Variable description

Variables	Identifier	Measurement	Source
Dependent variable			
Economic growth	*EG*	The Gross Provincial Product per capita (GPP per capita) in $US[4]	Statistical yearbook of each province
Financial inclusion			
Bank branches	*FI1*	Number of commercial bank branches	Annual statistics of The State Bank of Vietnam in provincial branches
Bank account	*FI2*	% Households having a bank account	Censuses of the General Statistics Office of Vietnam
Saving passbook	*FI3*	% Households having a savings passbook	
ATM cards	*FI4*	% Households having ATM cards	
Insurance	*FI5*	% Households participating in life and non-life insurance	
Control variables			
Inflation rate	*INF*	CPI (in %)	Statistical yearbook of each province
Labour force	*LABOUR*	% Population ageing 15+/ total population	
Foreign direct investment inflow	*FDI*	The ratio of foreign direct investment inflow to gross provincial product (in %)	
Urban population	*URBAN*	% Inhabitants living in urban regions	
Multi-dimensional poverty rate	*POVERTY*	% Poor households	

Source: own work.

[4] The study uses the logarithm of GPP per capita in estimations.

Table 2. Summary statistics for all Vietnamese provinces

Variable	Observations	Mean	Standard deviation	Min	Max
EG	441	2,491.302	1,786.970	789.630	17,277.860
FI1	441	174.379	373.847	15	3029
FI2	441	22.129	14.763	1.550	79.887
FI3	441	10.693	6.650	1.333	36.238
FI4	441	32.307	14.410	6.250	79.209
FI5	441	6.227	6.099	0.191	74.561
INF	441	2.626	1.918	–2.590	10.070
LABOUR	441	59.027	5.217	48.819	76.998
FDI	441	62.752	79.875	–8.790	464.067
URBAN	441	28.573	17.264	9.800	87.356
POVERTY	441	9.132	9.099	0.090	44.820

Source: own calculations.

Table 3. Summary statistics for two income groups of Vietnamese provinces

Variable	Low-income provinces			High-income provinces		
	mean	min	max	mean	min	max
EG	1,778.476	789.63	3,381.000	4,023.877	1,404.761	17,277.860
FI1	83.731	15	250	369.271	66	3029
FI2	16.260	1.550	51.667	34.749	2.899	79.887
FI3	8.745	1.333	35.238	14.880	2.469	36.238
FI4	26.380	6.250	62.262	45.051	11.728	79.209
FI5	6.044	0.190	74.561	6.619	0.328	40.108
INF	2.586	–2.590	10.070	2.712	–1.838	9.690
LABOUR	59.538	49.510	76.998	57.243	48.819	71.410
FDI	47.691	–8.790	464.067	95.133	0.025	227.221
URBAN	22.056	9.800	49.880	42.585	12.155	87.356
POVERTY	11.939	0.470	44.820	3.098	0.090	13.400
Number of provinces	43			20		
Number of observations	301			140		

Source: own calculations.

This paper aims to examine the impact of financial inclusion on economic growth in Vietnam using data from Vietnamese provinces. The empirical model is as follows:

$$EG_{i,t} = \alpha + \beta_1 EG_{i,t-1} + \beta_2 FI_{i,t} + \beta CV_{i,t} + \varepsilon_{i,t} \quad (1)$$

where: $CV_{i,t}$ is vector of control variables; i represents a provincial unit, $i = 1, 2, ..., 63$; t indicates the time period from 2014 to 2020; $\varepsilon_{i,t}$ is error term.

We constructed the panel dataset for 63 Vietnamese provinces in 2014–2020. As the panel has $N = 63 > T = 7$, it is suitable to apply the Generalised Method of Moments (GMM) estimation. GMM is meant for estimating dynamic panel models because it can effectively deal with endogeneity of the lagged dependent variable. Moreover, GMM employs internal instruments instead of using additional instrumental variables to mitigate bias (Arellano & Bond, 1991). We first apply the difference-GMM. While difference-GMM can address endogeneity, it may suffer from finite-sample bias because of instrument bias or weak instruments of the level variables (Bond, 2002). In this case, the system-GMM can control for finite sample bias and improve the consistency of estimated parameters (Blundell & Bond, 1998; Bond, 2002). Therefore, we also use the system-GMM estimation for a robustness check. To avoid spurious regressions arising from the presence of unit roots, all variables in the model are tested for panel unit root using the Augmented Dicky-Fuller test (ADF test). The results point to the rejection of the null hypothesis, suggesting the stationarity of the panel data series (see Appendix B).

3. Results and discussion

3.1. Financial inclusion and economic growth in Vietnam

Table 4 presents the difference-GMM results for the effect of financial inclusion on economic growth in Vietnam with the positive and significant coefficients for four indicators (FI1, FI2, FI3, and FI4), while there is no statistically significant effect for FI5.

Column (1) and column (2) demonstrate the positive link that means a higher number of commercial bank branches (FI1) and a higher percentage of households having bank accounts (FI2) results in bigger economic growth in Vietnamese provinces. These indicators capture geographical penetration of bank branches as well as the availability and accessibility of bank services. More commercial bank branches could serve a wider range of customers' banking services, reduce time costs, commuting costs for customers and, con-

Table 4. The impact of financial inclusion on economic growth in Vietnam (difference-GMM)

Variables	FI1	FI2	FI3	FI4	FI5
EG_{t-1}	0.135	1.069***	1.182***	0.994***	0.303***
	(0.097)	(0.180)	(0.189)	(0.144)	(0.112)
$FI_{i,t}$	0.711***	0.096*	0.188*	0.236***	0.007
	(0.121)	(0.055)	(0.107)	(0.089)	(0.035)
$INF_{i,t}$	0.006***	0.005	0.006	0.004	0.046***
	(0.002)	(0.004)	(0.005)	(0.003)	(0.008)
$LABOUR_{i,t}$	0.066***	0.095***	0.091***	0.080***	0.153***
	(0.018)	(0.032)	(0.025)	(0.027)	(0.035)
$FDI_{i,t}$	0.016***	0.009*	0.005	0.015***	–0.003
	(0.005)	(0.005)	(0.006)	(0.005)	(0.011)
$URBAN_{i,t}$	–0.044	–0.054	0.098	–0.041	0.239*
	(0.061)	(0.058)	(0.255)	(0.065)	(0.137)
$POVERTY_{i,t}$	–0.070*	0.088	–0.134*	0.088	–0.162***
	(0.042)	(0.059)	(0.078)	(0.058)	(0.046)
Number of observations	313	313	313	313	313
AR(1)	0.054	0.001	0.001	0.002	0.001
AR(2)	0.513	0.218	0.312	0.386	0.183
Hansen test for overidentification	0.289	0.245	0.603	0.164	0.182
Hansen test of exogeneity	0.437	0.547	0.308	0.455	0.324

Note: Robust standard errors reported in parenthesis. ***, **, *, significant at the 1%, 5%, and 10% levels, respectively. The results of AR(1), AR(2) and Hansen test are presented as p-values.

Source: own calculations.

sequently, demand for financial products and services which help promote investment and economic growth (Maity & Sahu, 2023). Also, the expansion of commercial bank branches could promote the integration of digital financial services, as banks are able to leverage platforms such as mobile and internet banking to deliver low-cost and diverse financial products. The combination of physical presence and digital delivery addresses both the access and usage dimensions of financial inclusion. This integrated approach enhances capital mobilisation, facilitates investment, and fosters entrepreneurial activity. In addition, having a bank account is considered as a gateway to facilitating customers' access to other financial services. Hence, expanding the chain of commercial bank branches and the percentage of households with bank accounts could also enhance borrowing and saving procedures; having more commercial bank branches would increase the availability of credit sources to rural areas, in turn facilitating bankers' ability to process payments and evaluate debts. Moreover, local customers can easily find bank branches in their area as a place to save money in their current account (Ghosh, 2011). This can increase the capital supply and boost economic growth in the country. This positive nexus is also documented in other developing countries (Pal et al., 2025; Wibowo et al., 2023).

Column (3) of Table 2 reports that financial inclusion expressed as the percentage of households with saving passbooks (FI3) could facilitate credit availability. This in turn promotes business activities, enhancing economic growth (Beck et al., 2009). Similarly, column (4) shows that financial inclusion measured by the percentage of households having ATM cards (FI4) positively affects economic growth. By owning an ATM card, households can conduct simple transactions such as making deposits, transferring money, withdrawing cash, and paying bills. This can ease the trade of goods and services and accelerate economic growth (Ehiedu et al., 2021; Gehrung, 2020). The positive effect of financial inclusion proxied by savings and ATM cards is consistent with previous studies for India (Singh & Mallick, 2024), Asia-Pacific countries (Basnayake et al., 2024), and sub-Saharan Africa countries (Osuma, 2025).

The effect of financial inclusion proxied by the usage of insurance (FI5) shown in column (5) is non-significant. Though participation in life and non-life insurance could provide assurance for economic activities, Table 2 reports a low level of households' insurance participation compared to other products and services in the financial system. In addition, the financial literacy of citizens in Vietnam remains low, as does basic knowledge risk management, therefore limiting inclusion into the financial insurance system, which may lead to a non-significant effect on economic growth in Vietnam (Barcellos & Zamarro, 2019; T. A. N. Nguyen & K. M. Nguyen, 2020). This result suggests a need for further analysis of the insurance aspects of financial inclusion in Vietnam.

The results for control variables also deserve a comment. The share of urban population (URBAN) exhibits a positive effect, while the poverty rate (POVERTY) negatively affects economic growth, as expected. The positive impact of inflation (INF), although appearing to be inconsistent with standard macroeconomic theory, can be explained by the non-linear relationship documented in prior research. M. S. Khan and Senhadji (2000) found that inflation promotes growth in developing countries when maintained within 7%–11%, while Sarel (1995) identified a structural break at around 8%, with positive effects below and negative effects above this threshold. These findings suggest that Vietnam's moderate inflation during the study period likely fell within a range that stimulates production, job creation, and consumption, consistent with Dorrance's (1964) mechanism.

3.2. Financial inclusion and economic growth across income groups

Table 5 illustrates the association between financial inclusion and economic growth in Vietnam across two income groups: low-income provinces and high-income provinces. Consistent with the result of the whole sample shown in Table 4, the number of commercial bank branches, the share of households owning banking account, saving passbook and ATM cards, displayed in columns (1)–(4) and (6)–(9), all present significantly positive effects on economic growth in both groups. For participation in life and non-life insurance, there are different results across two groups. Provinces with a GPP below the country's average present a non-significant effect of insurance participation on economic growth, similar to the result for the whole group. Meanwhile, insurance participation has a significantly positive influences on economic growth in upper-average GPP provinces (column (10)). The explanation for this discrepancy is that high-income provinces have a higher percentage of households participating in insurance, together with a higher income and better background related to investment, urban population and the poverty rate (see Table 3). These conditions seem to facilitate insurance in the economy, spurring significant impacts on economic growth in high-income provinces.

Table 5 also presents the coefficients for control variables, pointing to the positive effects of inflation, foreign direct investment, urban population on economic growth, while the poverty rate is found to negatively affect economic growth. Notably, in some regressions for low-income provinces, we found negative effects of labour on economic growth. This is due to the economy in these localities being underdeveloped and focused on seasonal agriculture activities, with a shortage of manufacturing industry. The workforce tends to migrate to work for companies in urban areas of richer provinces or as export labour (N. A. Nguyen et al., 2018; Pham et al., 2018).

3.3. Robustness test

To verify the robustness of the results estimated using the difference-GMM method, we apply the system-GMM estimation. The relevant estimation results are presented in Table 6. Similarly to the baseline findings, various aspects of financial inclusion maintain their positive and statistically significant impact on economic growth for the whole group of Vietnamese provinces. Although the magnitude of the coefficients exhibits minor changes, it retains the concept of intuition in the previous section. Therefore, the main findings estimated using the difference-GMM approach appear robust and stable, which improves the reliability of the results.

Table 5. The impact of financial inclusion on economic growth in Vietnam across income groups (difference-GMM)

Groups	Low-income provinces					High-income provinces				
Variables	**FI1**	**FI2**	**FI3**	**FI4**	**FI5**	**FI1**	**FI2**	**FI3**	**FI4**	**FI5**
EG_{t-1}	0.026	0.418***	0.814***	0.854***	1.156***	0.501**	0.623**	1.034***	0.697**	0.669*
	(0.097)	(0.094)	(0.134)	(0.168)	(0.155)	(0.194)	(0.307)	(0.198)	(0.287)	(0.377)
$FI_{i,t}$	0.719***	0.123***	0.236*	0.145	0.011	0.767***	0.530***	0.201**	0.653***	0.015**
	(0.169)	(0.039)	(0.125)	(0.087)	(0.026)	(0.136)	(0.170)	(0.080)	(0.248)	(0.007)
$INF_{i,t}$	0.006***	0.022***	0.010**	0.007**	0.007*	0.004*	–0.008	0.004	0.006*	0.003
	(0.002)	(0.004)	(0.004)	(0.003)	(0.004)	(0.002)	(0.005)	(0.006)	(0.003)	(0.004)
$LABOUR_{i,t}$	–0.091	–0.164	–0.187*	–0.112*	–0.126**	–0.520	0.001*	–0.517	–0.179	0.578*
	(0.141)	(1.130)	(0.091)	(0.006)	(0.061)	(0.488)	(0.000)	(0.537)	(0.540)	(0.262)
$FDI_{i,t}$	0.010**	–0.004	0.013**	0.015***	0.003	0.022***	0.022**	0.003	0.034**	0.040***
	(0.005)	(0.007)	(0.006)	(0.005)	(0.005)	(0.006)	(0.010)	(0.001)	(0.009)	(0.012)
$URBAN_{i,t}$	0.008	0.144*	0.184	0.073	0.019	0.223*	–0.176	0.201	0.274	0.127*
	(0.079)	(0.070)	(0.217)	(0.067)	(0.072)	(0.122)	(0.157)	(0.316)	(0.206)	(0.065)
$POVERTY_{i,t}$	–0.010*	–0.080**	–0.004	0.001	0.075	0.058	0.101	–0.081*	0.067	–0.081*
	(0.061)	(0.036)	(0.037)	(0.067)	(0.059)	(0.051)	(0.124)	(0.045)	(0.074)	(0.041)
Number of observations	213	213	213	213	213	100	100	100	100	100
AR(1)	0.030	0.005	0.004	0.001	0.002	0.831	0.099	0.087	0.074	0.049
AR(2)	0.667	0.115	0.808	0.904	0.805	0.114	0.689	0.130	0.470	0.643
Hansen test for overidentification	0.248	0.318	0.207	0.212	0.458	0.706	0.347	0.761	0.318	0.224
Hansen test of exogeneity	0.280	0.765	0.697	0.250	0.690	0.555	0.689	0.706	0.493	0.691

Note: Robust standard errors reported in parenthesis. ***, **, *, significant at the 1%, 5%, and 10% levels, respectively. The results of AR(1), AR(2) and Hansen test are presented as *p*-value.

Source: own calculations.

Table 6. The impact of financial inclusion on economic growth in Vietnam (system-GMM)

Sample	Variables	FI1		FI2		FI3		FI4		FI5	
All provinces	EG_{t-1}	1.108***	(0.020)	0.001***	(0.000)	0.001***	(0.000)	0.001***	(0.000)	0.952***	(0.057)
	$FI_{i,t}$	0.068*	(0.030)	0.137**	(0.063)	0.108**	(0.052)	0.195*	(0.105)	0.016	(0.021)
	$INF_{i,t}$	0.001	(0.004)	0.008***	(0.003)	0.010*	(0.005)	0.007*	(0.003)	0.008**	(0.003)
	$LABOUR_{i,t}$	0.077	(0.09)	0.010	(0.213)	1.345*	(0.751)	0.016	(0.426)	0.372	(0.336)
	$FDI_{i,t}$	0.010***	(0.003)	0.010*	(0.005)	0.006	(0.008)	0.010*	(0.006)	0.010**	(0.004)
	$URBAN_{i,t}$	0.096**	(0.038)	0.089	(0.094)	0.045	(0.137)	0.160	(0.156)	0.073	(0.058)
	$POVERTY_{i,t}$	0.063	(0.041)	–0.012	(0.912)	0.050	(0.042)	0.010	(0.046)	–0.027*	(0.015)
	AR(1); AR(2)	0.004;	0.233	0.565;	0.182	0.397;	0.588	0.546;	0.225	0.001;	0.222
	Hansen test for overidentification; exogeneity	0.454;	0.543	0.352;	0.602	0.503;	0.224	0.259;	0.428	0.656;	0.153
Low-income group	EG_{t-1}	0.854***	(0.064)	0.986***	(0.050)	0.840***	(0.099)	0.787***	(0.084)	1.084***	(0.033)
	$FI_{i,t}$	0.100***	(0.038)	0.003**	(0.001)	0.074*	(0.043)	0.109*	(0.060)	0.016	(0.016)
	$INF_{i,t}$	0.011***	(0.004)	0.002	(0.004)	0.002*	(0.001)	0.007*	(0.003)	0.007*	(0.004)
	$LABOUR_{i,t}$	–0.575*	(0.027)	–0.131	(0.004)	0.115	(0.374)	–0.059*	(0.030)	–0.584*	(0.03)
	$FDI_{i,t}$	0.001	(0.005)	0.011***	(0.003)	0.007*	(0.004)	0.004	(0.004)	0.011***	(0.003)
	$URBAN_{i,t}$	0.032	(0.119)	–0.031	(0.067)	0.075*	(0.035)	0.023	(0.048)	0.036**	(0.015)
	$POVERTY_{i,t}$	–0.037**	(0.019)	0.014	(0.013)	–0.024	(0.020)	–0.018	(0.015)	–0.050**	(0.020)
	AR(1); AR(2)	0.008;	0.361	0.010;	0.871	0.009;	0.996	0.002;	0.648	0.005;	0.729
	Hansen test for overidentification; exogeneity	0.334;	0.254	0.372;	0.666	0.338;	0.763	0.249;	0.326	0.281;	0.404

Table 6. continued

Sample	Variables	FI1		FI2		FI3		FI4		FI5	
High-income group	EG_{t-1}	1.110***	(0.110)	0.970***	(0.036)	1.059***	(0.062)	0.876***	(0.095)	0.901***	(0.131)
	$FI_{i,t}$	0.130**	(0.066)	0.129***	(0.047)	0.144*	(0.074)	0.132**	(0.063)	0.041**	(0.017)
	$INF_{i,t}$	0.006*	(0.003)	0.002	(0.004)	0.011	(0.009)	0.013**	(0.006)	0.008	(0.009)
	$LABOUR_{i,t}$	0.019	(0.395)	0.683*	(0.422)	0.105	(0.177)	0.081	(0.151)	0.067	(0.214)
	$FDI_{i,t}$	0.009***	(0.014)	0.019***	(0.007)	0.014**	(0.006)	0.017***	(0.007)	0.021**	(0.009)
	$URBAN_{i,t}$	0.162**	(0.069)	0.048	(0.062)	0.082**	(0.073)	0.022	(0.630)	0.034*	(0.015)
	$POVERTY_{i,t}$	–0.081*	(0.042)	–0.081**	(0.037)	0.040*	(0.021)	0.010	(0.030)	–0.06	(0.041)
	AR(1); AR(2)	0.282;	0.177	0.101;	0.148	0.054;	0.162	0.052;	0.132	0.035;	0.108
	Hansen test for overidentification; exogeneity	0.753;	0.472	0.602;	0.375	0.625;	0.724	0.479;	0.726	0.578;	0.511

Note: Robust standard errors reported in parenthesis. ***, **, *, significant at the 1%, 5%, and 10% levels, respectively. The results of AR(1), AR(2) and Hansen test are presented as p-value.

Source: own calculations.

Conclusions

The main objective of this paper is to examine the relationship between financial inclusion and economic growth in Vietnamese provinces during 2014–2020. The study used difference-GMM to estimate empirical results and system-GMM for the sake of robustness check. The results suggest the positive linkage for the following measures of financial inclusion: the penetration of commercial bank branches and the usage of banking accounts, savings, and ATM cards. Meanwhile, financial inclusion proxied by insurance participation shows a non-significant effect on economic growth. Similar findings are documented in provinces which have GPP per capita lower than the country's average. Notably, in high-income provinces, it is found that a higher percentage of households participating in insurance could spur economic development. In addition, both high-income and low-income provinces in Vietnam exhibit a significant effect of financial inclusion, measured as commercial bank branches penetration and using ATM cards.

The findings illustrate the crucial role of financial inclusion in accelerating economic growth in Vietnam. The positive linkage suggests that policymakers should prioritise measures to expand access to and use of financial services. The study also raises concerns that insurance participation in Vietnam is still in low level and has non-significant effect on economic growth in low-income provinces. To address this issue, policymakers should design targeted programmes to increase the accessibility of insurance products, particularly in rural and low-income regions. Such policies could include incentives for insurers to expand their reach and foster public–private partnerships to develop affordable products tailored to local needs. In addition, improving financial literacy, especially financial knowledge about the role and benefits of insurance, will help households and businesses understand how insurance can protect assets, stabilise income, and support economic resilience, thereby enhancing its potential contribution to growth.

The statistically significant positive effect of the actual use of financial products, services and penetration highlights the necessity of targeted policies. Policymakers should endorse initiatives that actively encourage individuals and businesses to engage with the available financial products and services. Such initiatives may encompass targeted awareness campaigns, incentives for use, and programmes to strengthen financial literacy, particularly practical skills in managing savings, credit, and digital financial tools. By enhancing both access and capability, these policies have the potential to amplify the growth-enhancing impact of financial inclusion.

Furthermore, the empirical results indicate that the positive impact of financial inclusion, as measured by the number of commercial bank branches, exists in both low-income provinces and high-income provinces. Given that the financial infrastructure in low-income provinces remains underdeveloped, targeted investments in branch networks, mobile banking services, and related infrastructure are likely to stimulate local economic activity and narrow disparities in growth.

Also, enhancing financial inclusion in Vietnam should progress in parallel with the implementation of complementary policies that strengthen the regulatory framework, improve the financial infrastructure, particularly in underdeveloped regions, and advance financial literacy. These conditions would provide a strong foundation and help ensure that expanded access to financial resources translates into substantive economic gains.

The current study has some limitations, which can be addressed in future research. Firstly, provincial data in Vietnam in terms of digital applications in the financial system, such as mobile banking, internet banking as well as credit from financial institutions, is currently not available. Future works can include them to measure financial inclusion. Secondly, this study ignores the contribution of microfinance institutions. Future studies on the impact of financial inclusion on economic growth should include this type of institution. While the study illustrates that financial inclusion promotes economic growth in Vietnam, the inverse relationship should also be discussed in future studies. In addition, this study has focused solely on economic growth; further research can observe other aspects such as poverty, women's empowerment, and well-being.

Appendix A

List of 63 provinces in Vietnam

An Giang	Dak Nong	Kon Tum	Quang Tri
Ba Ria–Vung Tau	Dien Bien	Lai Chau	Soc Trang
Bac Lieu	Dong Nai	Lang Son	Son La
Bac Giang	Dong Thap	Lao Cai	Tay Ninh
Bac Kan	Gia Lai	Lam Dong	Thai Binh
Bac Ninh	Ha Giang	Long An	Thai Nguyen
Ben Tre	Ha Nam	Nam Dinh	Thanh Hoa
Binh Duong	Hanoi	Nghe An	Ho Chi Minh City
Binh Dinh	Ha Tinh	Ninh Binh	Thua Thien Hue
Binh Phuoc	Hai Duong	Ninh Thuan	Tien Giang
Binh Thuan	Hai Phong	Phu Tho	Tra Vinh
Ca Mau	Hau Giang	Phu Yen	Tuyen Quang
Cao Bang	Hoa Binh	Quang Binh	Vinh Long
Can Tho	Hung Yen	Quang Nam	Vinh Phuc
Da Nang	Khanh Hoa	Quang Ngai	Yen Bai
Dak Lak	Kien Giang	Quang Ninh	

Appendix B

Table B1. Unit root test results

Variable	ADF			
	inverse chi--squared	inverse normal	inverse logit	modified inverse chi-squared
EG	681.613***	–12.204***	–21.737***	35.000***
FI1	304.359***	–4.019***	–6.461***	11.236***
FI2	235.776***	–2.547***	–4.816***	6.915***
FI3	386.440***	–2.615***	–8.065***	16.406***
FI4	2501.015***	–39.926***	–88.419***	149.612***
FI5	203.384***	–1.324*	–3.212***	4.875***
INF	419.908***	–10.700***	–12.896***	18.515***
LABOUR	522.299***	–9.961***	–15.476***	24.965***
FDI	400.993***	–2.096**	–8.428***	17.323***
URBAN	432.627***	–4.664***	–9.702***	19.253***
POVERTY	1420.691***	–22.399***	–47.458***	81.558***

Note: *,**,*** indicates the significance level of the 10%, 5%, and 1%, respectively.

Source: own calculations.

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Perceived usefulness, ease of use, risk, and trust: Explaining BNPL user recommendation intention through behavioural models

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Abstract

The main objective of the article is to identify and evaluate the key determinants of consumer recommendation intention with Buy Now, Pay Later (BNPL) services, operationalised

Keywords

- BNPL
- SEM
- consumer finance

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through the intention to recommend such services to others. The study investigates the influence of five core constructs: perceived usefulness, perceived ease of use, perceived trust, and perceived risk. Data are collected from a quota sample of 350 users of deferred payment services, selected in accordance with the demographic profile of BNPL users. The study employs PLS-SEM. The results show that perceived usefulness and perceived trust in the BNPL provider significantly boost recommendation intention. Perceived risk negatively impacts recommendation intention, while perceived ease of use has only a marginal effect. These results contribute to the existing literature by elucidating the behavioural mechanisms underlying BNPL usage and provide actionable insights for financial service providers aiming to enhance consumer recommendation intention and retention.

• consumer recommendation intention

JEL codes: G23, G41, G51

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Introduction

The Buy Now, Pay Later (BNPL) payment model has emerged as one of the most prominent financial innovations within the digital economy over the past decade. Initially introduced as a niche credit solution, BNPL has rapidly evolved into a mainstream consumer financing tool, particularly within e-commerce platforms. It allows consumers to defer payment—typically without interest and in flexible instalments—thus offering a combination of accessibility, convenience, and short-term liquidity. Scientific research has shown that this effectively encourages younger generations to use BNPL (Blue et al., 2024). The model has proven especially appealing to younger, digitally literate users who prioritise speed, control, and budget flexibility. BNPL has gained considerable momentum globally, including in the Polish market, where its adoption continues to accelerate (Waliszewski et al., 2024). This dynamic growth highlights the increasing significance of BNPL from both a market development perspective and a consumer behaviour standpoint.

According to Statista data in 2024, BNPL's share of Poland's e-commerce payments stood at 3%, indicating that while this payment method is present, it is still in the initial stage. Poland's result is slightly below the global average of 5%, ranking it among the countries with a moderate level of BNPL adoption, similar to India, Ireland, Indonesia, Singapore, and Spain and relatively low if compared to Sweden or Germany (Statista, 2025).

In light of this rapid expansion, a deeper understanding of the mechanisms driving consumer engagement with BNPL services is becoming increasingly important. While previous studies (Kinanthi et al., 2024) have investigated the initial adoption of financial technologies—including BNPL—drawing on established theoretical frameworks such as the Theory of Planned Behaviour (TPB), the Technology Acceptance Model (TAM), and the Theory of Perceived Risk (TPR), relatively little is known about the post-adoption phase; in particular, user recommendation intention remain underexplored. Existing research primarily focuses on the decision to adopt or reject BNPL services (Akana, 2022; Dadra et al., 2024), leaving open the question of what drives continued usage and advocacy among users. Addressing this gap is essential for two key reasons. Firstly, identifying the drivers of recommendation intention provides valuable insight into whether BNPL services deliver on their perceived value proposition. Secondly, recommendation intention can serve as a behavioural proxy for satisfaction and a predictor of market diffusion through word-of-mouth recommendations.

This study addresses this gap by focusing on the determinants of consumer recommendation intention as regards BNPL services, using recommendation intention as the key dependent variable. Specifically, the article investigates five conceptual constructs identified in the literature as influential in shaping user attitudes towards FinTech services: perceived ease of use, perceived usefulness, perceived trust in the BNPL provider, perceived trust in the BNPL service itself, and perceived risk.

The main objective of this study is to investigate how these factors influence consumers' willingness to recommend BNPL services to others—interpreted here as a manifestation of a positive user experience and the fulfilment of prior expectations. This analytical perspective is grounded in the logic of Expectation-Confirmation Theory (ECT) and the Net Promoter Score (NPS) framework, both of which conceptualise recommendation behaviour as a proxy for consumer satisfaction and loyalty.

This study offers several contributions to the existing body of knowledge. Firstly, it enriches the academic discourse by shifting the analytical focus from one-time adoption to sustained usage and consumer satisfaction. Secondly, it provides empirical insights from the Polish BNPL market, which remains underrepresented in the international literature. Thirdly, it advances the conceptual understanding of perceived trust by disaggregating it into perceived trust in the provider versus perceived trust in the product, thereby enabling a more nuanced interpretation of the user experience. Finally, the findings yield practical implications for BNPL providers (including FinTech firms) and policymakers, particularly in the areas of user-centred service design and the mitigation of consumer risks.

The article is structured as follows: Section 1 presents the theoretical foundations of the study and outlines the development of research hypotheses, drawing on TPB, TAM, and TPR. Section 2 details the methodology, including the research instrument, data collection procedure, and measurement model construction. Section 3 reports the empirical results and evaluates the structural model. Section 4 discusses the findings in the context of prior research and identifies theoretical and practical implications. The limitations of the study and directions for future research are presented in section 5. The final section concludes the article with a summary of key contributions.

1. Literature review and hypothesis development

The intention to use a new product (in this case, BNPL) is derived from the consumer's attitude toward the innovation. This attitude is shaped by perceived benefits (e.g., usefulness and ease of use), perceived trust and perceived risk. Therefore, three fundamental theories support the explanation of consumer behaviour toward BNPL, which is primarily offered in the context of online sales: the Theory of Reasoned Action (TRA) and its derivatives, i.e. the Theory of Planned Behaviour (TPB) and the Reasoned Action Approach (RAA), assuming that behaviour is determined by the combination of attitude, subjective norms, and perceived behavioural control; the Technology Acceptance Model (TAM), consisting in perceived usefulness and perceived ease of use as factors affecting behaviour; and the Theory of Perceived Risk (TPR), according to which behaviour is conditioned by the consumer's perception of the risk.

The starting point for TRA is the Expectancy-Value Theory (EVT), the foundations of which were formulated by Atkinson (1953). EVT explains the motivation behind behaviour. According to this theory, motivation is a function of expectations (beliefs) regarding success (the degree of certainty that success can be achieved or that a specific benefit can be obtained) and the subjectively perceived value of the outcome (i.e. how valuable the benefit of using a given good or service appears to be). The extent to which individuals believe that achieving the goal is both possible and meaningful influences their engagement in a given task. Other factors—such as stereotypes, prior experiences, and the experiences of others—are already reflected within expectations and value. Expectations depend on beliefs about one's own abilities and the perceived effectiveness of their application. These beliefs also include the anticipated consequences of the behaviour. In turn, the subjective value of the task is closely tied to the importance of the task for identity maintenance, intrinsic value (e.g., pleasure or interest), utility value (e.g., usefulness or relevance), or perceived cost (e.g., time, effort, stress). In other words, the EVT model consists of three components: a response to new information in the form of belief formation, the assignment of value to each belief attribute, and the development of expectations based on beliefs and the values (weights) assigned to them. Fishbein and Ajzen (1975a, 1975b), as well as Fishbein (1967, 1980), extend EVT by formulating the TRA. This theory, originating from social psychology, explains the impact of attitudes (determined by motivation) on actual behaviour and identifies the variables that account for such behaviour (Adamek & Solarz, 2024).

TRA enables the prediction of behaviour based on behavioural intentions, defined as the extent to which a consumer plans to act in a given situation. These intentions are determined by two independent constructs: attitudes, which are a function of beliefs and their associated values (i.e. how (un)favourable the behaviour is perceived to be), and subjective norms, which reflect perceived social pressure (e.g., from parents, peers) to perform or refrain from a specific behaviour. Attitudes toward a given good or service may be positive, negative, or neutral. According to TRA, the intention to perform a particular behaviour precedes the actual behaviour. This intention is shaped by the belief that performing the behaviour will lead to a specific outcome. TRA posits that stronger behavioural intentions lead to greater effort to enact the behaviour. If an individual evaluates the behaviour positively (attitude) and believes it aligns with the values of their social environment (subjective norm), then their behavioural intention—or motivation—to act will be stronger. This intention is highly correlated with actual behaviour. According to TRA, if a person believes that a particular behaviour will not bring any benefit, they form a negative attitude. (Adamek & Solarz, 2024). Due to criticism of theories that explain behaviour primarily through attitudes, and in an effort to increase the model's predictive power, TRA underwent modifications over subsequent decades. One of them is TPB, proposed by Ajzen (1985). By linking beliefs with behaviour, TPB assumes that behavioural intentions are shaped not only by attitudes and subjective norms but also by perceived behavioural control. Perceived behavioural control refers to the extent to which an individual believes it is capable of performing a given behaviour. At the same time, TPB suggests that people are much more likely to engage in certain behaviours when they feel confident in their ability to carry them out successfully. TPB becomes a foundational framework for empirical research on household financial decision-making. For instance, East (1993) demonstrates that subjective norms (the influence of friends and relatives) and perceived control (such as easy access to financial resources) affect individual investment decisions. In turn, Xiao and Wu (2008) argue that customer satisfaction impacts behavioural intentions and indirectly influences actual behaviour, while Bansal and Taylor (2002) suggest that significant interactions exist between perceived control and intention, perceived control and attitude, as well as between attitude and subjective norms.

Another modification, presented by Fishbein and Ajzen (2010), is referred to as the Reasoned Action Approach (RAA). The RAA model assumes that behaviour is a function of intention, moderated by actual control. Intention is determined by attitude, perceived norm, and perceived behavioural control. The latter affects behaviour both directly and indirectly through intention, while actual control influences perceived control.

Due to its innovative nature and advanced technological dimension TAM—an information systems theory—is also included among the theoretical frameworks that explain the tendency to adopt BNPL services. TAM is based on the concepts of Davis (1989) as well as Davis et al. (1989). According to this theory, as with TRA, behavioural intention is the primary antecedent of actual use of a new solution. This intention directly depends on the individual's beliefs regarding the use of the given technology. Beliefs, in turn, are shaped by two factors: perceived usefulness and perceived ease of use. Moreover, perceived usefulness itself is influenced by perceived ease of use. TAM may be regarded as one of the most important developments of TRA, in which many of the original attitude measures are replaced with technology acceptance constructs, namely ease of use and usefulness. TAM corresponds with Rogers' (1962) Diffusion of Innovations Theory, according to which the adoption of innovation is conditioned by factors such as relative advantage (leading to increased efficiency) and accessibility (understood as clarity and ease of use).

The foundations of TPR were established by Bauer (1960), who defined perceived consumer risk as an undesirable outcome resulting from the consumer's actions. Cunningham (1967) proposed using two dimensions to determine perceived risk: the magnitude of risk and the probability of its occurrence. Florea and Munteanu (2012) added a third component—the risk horizon, which is inversely proportional to perceived risk.

Perceived risk is a multidimensional concept, consisting of components, each of which contributes to the consumer's overall perception of risk (Mitchell, 1999). Mulino et al. (2009) argue that risk aversion is not a stable individual trait but varies for the same person depending on the decision-making context. Perceived risk has a direct influence on consumers' purchase intentions (Wei et al., 2018). Perceived trust, in turn, was first studied in the context of interpersonal relations and only later became recognised as an important factor in the adoption of FinTech innovations (Whitman & Mattord, 2009). Cao et al. (2018) identify it as one of the key factors capable of shaping human behaviour toward technology. Consumers' perceived trust in FinTech solutions has a strong impact on their adoption intentions (Moon & Kim, 2016).

The theoretical framework discussed above allows for the identification of the following constructs used in the empirical investigation: perceived risk, perceived ease of use, perceived usefulness, and perceived trust. Together, these models offer valuable insights into the mechanisms underlying the dynamic development of modern technologies in society, particularly by highlighting key determinants that shape the diffusion of deferred payment services. Figure 1 presents the links between the theories discussed in this subsection and the constructs used in the empirical study.

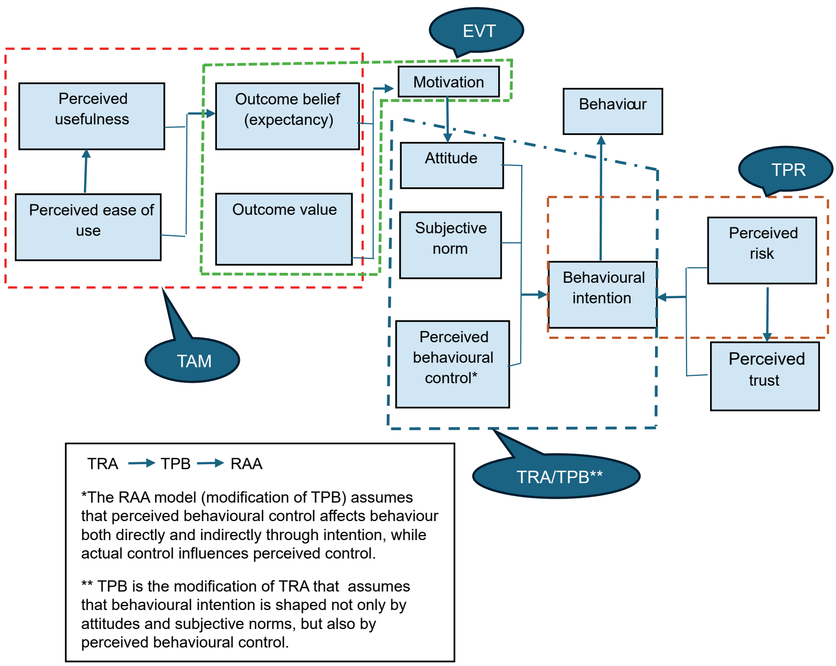

Figure 1. Theoretical framework and constructs used in the empirical study

Source: own work.

Molina-Castillo et al. (2016) argued that wireless and mobile technologies used in m-commerce facilitate social interactions and enable consumers to experience new forms of engagement. To support this claim, they drew upon motivation theory, self-efficacy theory, and TAM to develop a framework explaining consumer behaviour toward mobile payment services. Their extended TAM incorporated intrinsic motivations and extrinsic motivations, both adapted from motivation theory, as well as self-efficacy. Importantly, their model distinguished between mobile pre-payment and post-payment services, and the results varied across these two categories.

Raj et al. (2024b), employing an Extended TPB, demonstrated a relationship between privacy concerns and perceived trust in BNPL services. They found that attitude, subjective norms, and perceived behavioural control are key explanatory variables for consumer intentions regarding BNPL usage. Hidayat et al. (2024) used the TAM and TPB to analyse user intentions related to BNPL adoption in Indonesia. They identified perceived trust, subjective norms, and perceived risk as determinants influencing the intention to use BNPL. However, no significant relationship was found between perceived usefulness and usage intention, nor between perceived ease of use and usage intention.

Akana (2022), on the other hand, argued that for US consumers, a key factor influencing the use of deferred payment services is user-friendliness. A study by Jagadhita and Tjhin (2023) also investigated factors influencing user intentions to adopt deferred payment services in Indonesia. Drawing on TAM variables and integrating them with perceived risk and perceived trust as a mediating variable, they demonstrated that perceived ease of use, perceived usefulness, and perceived trust positively affects the intention to continue using BNPL. Moreover, perceived trust was identified as a partial mediator between both perceived ease of use and perceived usefulness. Behera and Dadra (2024), in their study on consumer attitudes toward deferred payments, found that the key determinants of those attitudes are structural assurance, flexibility, affordability, and perceived usefulness.

Van Tuan et al. (2024) aimed to identify determinants of BNPL usage, incorporated perspectives from the TRA, TPB, TAM, and TPR, among other theories. Based on this theoretical synthesis, they proposed a model comprising five key factors influencing BNPL usage intention: social influence, digital financial knowledge, perceived ease of use, perceived benefits, and perceived risks. According to their findings, social influence had the strongest effect on consumer attitudes toward BNPL usage. Supporting this conclusion, Aisjah (2024) also emphasised the significance of social influence, particularly highlighting the role of social media intensity in shaping consumers' intention to use deferred payment services.

Schomburgk and Hoffmann (2022), in their study of Australian consumers, observed that younger, tech-savvy individuals with a positive attitude toward innovation are more likely to adopt BNPL services. Osman et al. (2024) applied the TPB in a study focusing on Generation Z and observed that consumer choices regarding deferred payment options may be shaped by multiple factors. Among the most important, they identified attitude, subjective norms, perceived behavioural control, and, notably, perceived moral obligation to repay debt. In turn, materialism, money management skills, and self-efficacy were found to positively influence these four variables, reinforcing them and leading to greater use of BNPL services. Materialism reflects the influence of the desire to acquire goods on financial attitudes and behaviour, money management skills represent practical financial constraints, and self-efficacy reflects an individual's confidence in financial decision-making.

In their study, Raj et al. (2024a) emphasised the crucial role of perceived risks and perceived benefits in shaping consumer intentions to use deferred payment services. They argued that it is essential to examine how perceived benefits influence the adoption of BNPL solutions, as such benefits are among the primary drivers of technology acceptance. Conversely, understanding perceived risk is critical for predicting consumers' adoption intentions, as it negatively affects users' willingness to accept and utilise modern financial services. Their empirical findings revealed a significant negative effect of perceived risk on consumers' behavioural intention, underscoring the relevance of privacy concerns in the digital environment. On the other hand, perceived benefits—particularly convenience, financial flexibility, increased purchasing power, budgeting capabilities, and potential cost reductions—were found to enhance consumers' intention to use BNPL. Furthermore, performance expectancy (understood as ease of use) and social influence were also observed to positively affect behavioural intention. Finally, the study showed that consumers' behavioural intention increases their intention to continue using deferred payment services.

The review of theoretical frameworks and empirical findings provides the basis for formulating the study's research hypotheses. These hypotheses focus on assessing the effects of constructs identified in the literature—namely, usage intention, attitudes toward BNPL, perceived usefulness, and perceived risk—on consumer satisfaction. It is further assumed that satisfaction with BNPL services translates into a willingness to recommend them, and more broadly, into the diffusion of use through consumer advocacy.

Based on the assumptions of TAM and empirical findings of Akana (2022), as well as Jagadhita and Tjhin (2023), we propose the following hypothesis:

H1: Perceived ease of use of BNPL services positively affect consumer recommendation intention.

The literature highlights the performance expectancy and perceived usefulness of BNPL as a personal finance management tool (Jagadhita & Tjhin, 2023; Min & Cheng, 2023; Relja et al., 2024). The ability to postpone or reschedule payments increases consumer satisfaction and encourages recommendation behaviour. This leads to the following hypothesis:

H2: The primary reason consumers recommend BNPL services is their perceived usefulness in managing personal and household finances.

TRA, supported by TPR and respective empirical studies, suggests that perceived trust can either strengthen (Rouibah et al., 2016), weaken or have no impact (Hoo et al., 2024) on consumers' intention to use BNPL. Regardless of direction, perceived trust remains a critical determinant of recommendation intention. Therefore, we hypothesise:

H3: Consumer recommendation intention is determined by perceived trust.

Given the potential distinction between perceived trust in the BNPL provider and perceived trust in the BNPL product, we further specify this relationship:

H3A: Consumer recommendation intention is determined by perceived trust in the provider.
H3B: Consumer recommendation intention is determined by perceived trust in the product.

According to TPR and supporting empirical studies (e.g., Raj et al., 2024a; Rouibah et al., 2016), perceived risk weakens both perceived trust and behavioural intention. Assuming that consumers who are uncertain about a service are unlikely to recommend it, we posit:

H4: Perceived risk associated with BNPL services negatively affects consumers' intention to recommend them.

2. Methods

Expectation fulfilment is inherently multidimensional; however, in this study, it is operationalised through the self-declared likelihood of recommending BNPL to family or friends. This approach is grounded in the widely accepted premise that the greater the satisfaction, the stronger the consumer's intention to recommend a given solution. Accordingly, recommendation intention is used as a proxy for consumer satisfaction. This conceptualisation aligns with the logic of the Net Promoter Score (NPS) framework proposed by Reichheld (2003), which has since been widely adopted in customer satisfaction research (Dawes, 2024).

A graphical representation of the research hypotheses and the empirical model is presented in Figure 2. Additionally, for validation purposes, an alternative model was estimated (Figure 3), based on a simplified construct structure. Due to the cognitive similarity of the concepts, *perceived trust in the BNPL provider* and *perceived trust in the BNPL services* were merged into a single composite construct—*perceived trust in BNPL*. This approach aimed to assess the robustness of the findings with respect to a simplified conceptualisation of perceived trust in the model.

The proposed model includes five fundamental constructs that may shape recommendation intention: (1) perceived ease of use, (2) perceived usefulness, (3) perceived trust in the BNPL provider, (4) perceived trust in BNPL services, and (5) perceived risk. These constructs reflect the assumption that consum-

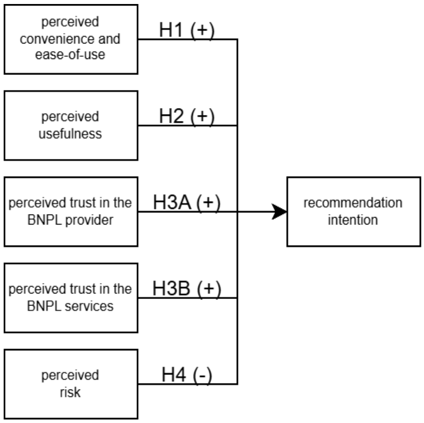

Figure 2. Baseline framework

Source: own work.

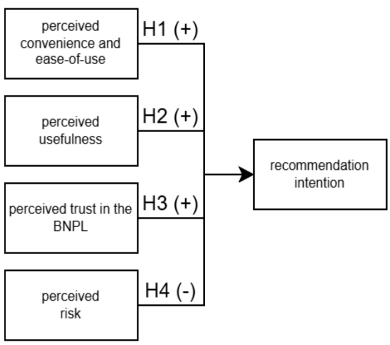

Figure 3. Alternative framework

Source: own work.

er expectation fulfilment depends not only on the core functionality of the instrument—namely, its ability to facilitate intertemporal decision-making—but also on perceived risk, and ease of use.

The data used in this analysis were obtained from a broader study examining the use of BNPL services among Polish consumers. The questionnaire included a cover letter informing respondents of the anonymity and voluntary nature of their participation, a demographic section, and a series of closed-ended, open-ended, and Likert-scale-based items.

For the purposes of the present analysis, responses provided on a 7-point Likert scale were used, where 1 indicated "strongly disagree" and 7 indicated "strongly agree." The model's constructs were operationalised using items measured on this scale. A detailed list of the measurement items assigned to each construct is presented in Table 1.

Table 1. Measurement items and construct definitions

Construct	Code	Measurement statement
Recommendation intention	RECOM_INT	I would recommend BNPL to my family and friends as a convenient payment method.
Perceived ease of use	PEOU_1	I believe using BNPL is convenient.
	PEOU_2	Using BNPL is intuitive and easy to understand.
	PEOU_3	BNPL transactions are completed quickly and efficiently, without unnecessary complications.
Perceived usefulness	PU_1	I believe that using BNPL is beneficial for my finances.
	PU_2	BNPL helps me manage my budget more effectively by allowing me to split payments into smaller parts.
	PU_3	BNPL allows me to avoid fees or interest typically associated with traditional consumer credit, making it a more advantageous option.
Perceived trust in the BNPL provider	PTRUST_P_1	I believe the company offering BNPL is transparent and honest about all fees and service terms.
	PTRUST_P_2	I consider the BNPL service to be reliable and dependable in all purchase situations.
	PTRUST_P_3	I believe companies offering BNPL are trustworthy.
Perceived trust in BNPL services	PTRUST_S_1	I am confident that using BNPL does not expose me to unexpected fees.
	PTRUST_S_2	I am confident that third parties will not gain access to my personal data when using BNPL.
	PTRUST_S_3	I believe that BNPL services offered by providers are secure.
	PTRUST_S_4	It is easy to find information about the terms and conditions of BNPL payments.
Perceived risk	PRISK_1	The ease of going into debt and the risk of falling into a debt trap.
	PRISK_2	Compulsive shopping and purchasing unnecessary or unneeded products.
	PRISK_3	Lack of understanding of BNPL rules, including the fee and interest calculation mechanisms.
	PRISK_4	Loss of personal data.

Note: PTRUST is a composite variable combining PTRUST_P and PTRUST_S.

Source: own work.

The survey was conducted using the CAWI (Computer-Assisted Web Interviewing) technique in August 2024. It was commissioned by PayPo Sp. z o.o.—one of the leading BNPL service providers in Poland—and carried out by the BioStat® Research and Development Centre. Data were collected from a quota sample of 350 users of deferred payment services. Respondents were selected based on two socio-demographic criteria: gender and age, in accordance with the profile of BNPL users reported by the Polish Credit Information Bureau (BIK) as of 31 December 2023. The sample was structured to reflect the age and gender distribution of BNPL users in Poland, as documented in the BIK dataset. Since official statistics in Poland do not disaggregate data by non-binary gender identities—and the sample was intended to mirror the demographic profile of users recorded in BIK—gender was used as a qualification criterion during the sampling process.

Model parameters were estimated using Partial Least Squares Structural Equation Modelling (PLS-SEM). This method is particularly suitable when (1) the sample size is moderate, (2) the assumption of normality is not met, (3) the model includes both reflective and formative constructs, and (4) the primary aim is to maximise the explained variance of the endogenous constructs. PLS-SEM has also been widely employed in recent studies on BNPL services (Adamek & Solarz, 2024; Dadra et al., 2024; Kumar & Nayak, 2024). Model estimation was performed in R (version 4.4.2) using the SEMinR package (version 2.3.4) for structural equation modelling and the psych package (version 2.4.12) for assessing reliability and validity of the measurement model. The criteria for model assessment are presented in Table 2.

Table 2. Model assessment summary

Assessment category	Criterion	Threshold
Indicator reliability	Factor loadings (λ)	$\lambda \geq 0.70$
Internal consistency	Composite reliability (CR)	$0.70 \leq CR \leq 0.95$
	Cronbach's alpha (α)	$\alpha \geq 0.70$
	rho_C, rho_A	$\rho \geq 0.70$
Convergent validity	Average variance extracted (AVE)	$AVE \geq 0.50$
Discriminant validity	Heterotrait–Monotrait ratio ($HTMT$)	$HTMT < 0.90$
Coefficient of determination	R^2 values for endogenous constructs	$R^2 \geq 0.30$ (satisfactory); ≥ 0.20 (moderate); ≥ 0.10 (weak)
Path significance	Bootstrapping (10,000 resamples)	$p < 0.05$

Source: based on (Hair et al., 2021; Streukens & Leroi-Werelds, 2016).

The presented SEM evaluation criteria are widely applied in the literature, which ensures their methodological credibility and validity. By meeting the conceptual assumptions of the SEM framework, it becomes possible to reliably estimate the relationships between variables and confirm the correctness of the proposed theoretical structure.

3. Results

A total of 350 respondents participated in the study. After verifying the completeness and consistency of the responses, all questionnaires were deemed valid and included in the analysis. Table 3 presents the socio-demographic characteristics of the respondents.

Table 3. Respondents' socio-demographic characteristics

Characteristic of the respondent	$N = 350$	
gender		
female	208	(59%)
male	142	(41%)
age		
18-24 y.o.	36	(10%)
25-34 y.o.	111	(32%)
35-44 y.o.	115	(33%)
45-54 y.o.	59	(17%)
55-64 y.o.	19	(5.4%)
65 y.o. or more	10	(2.9%)
education level		
below high school level	16	(5%)
high school level	130	(37%)
university level	204	(58%)
place of residence		
village	67	(19%)
small city (up to 20k inhabitants)	34	(10%)
medium-sized city (20k–99k inhabitants)	74	(21%)
big city (100k–200k inhabitants)	38	(11%)
large city (200k–500k inhabitants)	54	(15%)
very large city (above 500k inhabitants)	83	(24%)

Source: own work.

Women accounted for 59% of the sample. The most numerous age group was from 35 to 44 years, accounting for 33% of the sample, while the over 65 years old group constituted nearly 3%. Educationally, 58% had university-level education, and 37% had completed high school. Most respondents (89%) lived in an urban area, with around 24% residing in very large cities (above 500k inhabitants).

The model converged in a single iteration, indicating stability and proper specification. A graphical representation of the estimated model is provided in Figure 4.

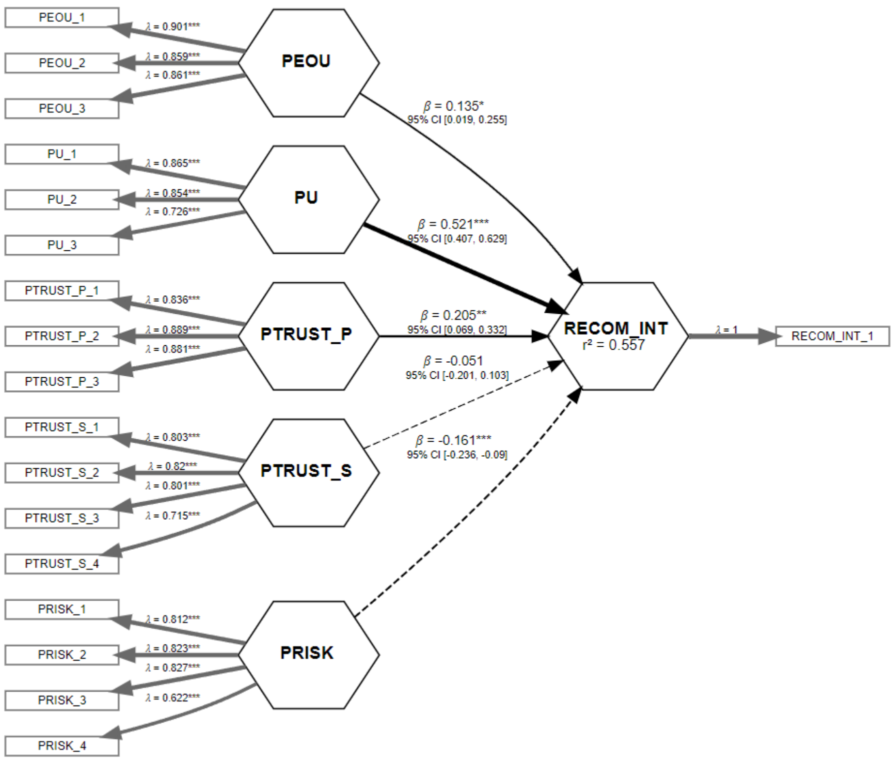

Figure 4. Estimated structural model

Source: own work.

Factor loadings for the items are generally above the recommended threshold of 0.70, with the exception of item PRISK_4, which reached a value of 0.622. However, removing this item did not improve the model fit or overall quality; therefore, it was retained in the model. The R^2 value of 0.557 indicates a satisfactory level of variance explained by the model. A detailed assessment of the model's reliability and validity is presented in Table 4.

All values of Cronbach's alpha, rho_C, and rho_A exceed the threshold of 0.70 while remaining below 0.95, indicating an appropriate level of internal

Table 4. Reliability and validity of model

	Alpha	rhoC	rhoA	AVE
PRISK	0.779	0.857	0.820	0.602
PEOU	0.851	0.907	0.917	0.764
PU	0.757	0.857	0.801	0.668
PTRUST_P	0.838	0.903	0.844	0.755
PTRUST_S	0.794	0.866	0.806	0.617

Note: As RECOM_INT is measured with a single item, a separate presentation is redundant.

Source: own work.

consistency without suggesting item redundancy. Additionally, all composite reliability (CR) coefficients are above 0.80, and the average variance extracted (AVE) values surpass the recommended threshold of 0.50, confirming both the reliability and convergent validity of the model.

Furthermore, all HTMT values remained below the conservative threshold of 0.90, except for the theoretically expected relationship between perceived trust in the BNPL provider (PTRUST_P) and perceived trust in BNPL services (PTRUST_S), which reached a value of 0.903. Given the conceptual proximity of these constructs, this minor exceedance is not considered problematic and justifies their treatment as analytically distinct variables. Overall, the validation results confirm that the model is suitable for testing the research hypotheses presented in Table 5.

Table 5. Structural model results: Path estimates and hypothesis support

Hypothesis	Structural path	Path coefficient	95% CI	*t*–Statistic	Significance	Result
H1	PEOU -> RECOM_INT	0.135	[0.021,0.255]	2.250	significant	confirmed
H2	PU -> RECOM_INT	0.521	[0.406,0.627]	9.264	significant	confirmed
H3A	PTRUST_P -> RECOM_INT	0.205	[0.067,0.334]	3.021	significant	confirmed
H3B	PTRUST_S-> RECOM_INT	–0.051	[–0.197,0.105]	–0.663	not significant	rejected
H4	PRISK-> RECOM_INT	–0.161	[–0.237, –0.091]	–4.355	significant	confirmed
H3*	TRUST -> RECOM_INT	0.162	[0.017,0.307]	2.203	significant	confirmed

* The alternative model yielded highly comparable results for the remaining path coefficients and construct reliabilities.

Source: own work.

Hypotheses H1, H2, and H4 are supported, as indicated by statistically significant effects in the expected direction. However, the interpretation of hypotheses H3, H3A, and H3B requires a more nuanced discussion. Hypothesis H3B, which assumed a significant effect of perceived trust in the product on recommendation intention, was not empirically supported. The results suggest that factors related to perceived trust in the institution providing the BNPL service play a more substantial role than perceived trust in the product itself. This finding confirms hypothesis H3A. The alternative model confirms that composite perceived trust significantly enhances recommendation intention. Thus, hypothesis H3 is supported.

4. Discussion

From the perspective of BNPL's growth potential, a key issue is whether this financial innovation contributes to increased consumer satisfaction. If so, its broader adoption can be expected—especially since the added value BNPL delivers to consumers may translate into benefits for merchants and lenders (Alcazar & Bradford, 2021). In this article, based on previous research exploring the determinants of BNPL adoption (Hoo et al., 2024; Raj et al., 2024a; Relja et al., 2024), consumer recommendation intention is modelled as a result of consumer perceptions in five key dimensions, represented by the following constructs: perceived risk, perceived ease of use, perceived usefulness, perceived trust in the BNPL provider, and perceived trust in the BNPL service itself. To deepen the discussion and derive more nuanced conclusions, it is essential to analyse the internal response structure within each construct, as it reveals which specific aspects of consumer perception drive the overall evaluations.

The results of this study confirm that BNPL contributes positively to consumer recommendation intention, due to its perceived usefulness. This finding supports H2 and is consistent with previous insights on perceived benefits reported by BNPL users (Relja et al., 2024; Schomburgk & Hoffmann, 2022; Tan, 2022). Deferred payments allow consumers to obtain and use products immediately, which increases satisfaction (Banerjee & Tóth, 2025). In this study, 52.5% of respondents agreed that BNPL is beneficial for their finances, over 60% stated it helps them manage their budget by distributing payments over time, and nearly 69% emphasised its financial attractiveness due to the absence of fees and interest. These results indicate that when consumers consider specific financial benefits, their satisfaction with BNPL increases. The findings are encouraging, as they reflect financial responsibility, pragmatism, and foresight among respondents. In their eyes, BNPL primarily serves as a tool to optimise household cash flow rather than a way to spend beyond their means. The absence of an association between BNPL and excessive debt, as well as the intention to use it solely for deferring payments, can be interpreted as a form of responsible consumption (Aalders, 2022). This optimism also extends to users' willingness to promote this financially rational approach to BNPL within their social circles.

Over 80% of respondents agreed with statements indicating that BNPL is an attractive service due to its accessibility, intuitive use, and ability to facilitate quick transactions. This finding aligns with previous research by Kiran and Mishra (2024), as well as Kiran et al. (2024). Therefore, perceived ease of use emerges as important factors influencing user satisfaction and their willingness to recommend BNPL services to family and friends. These results provide empirical support for H1. A substantial proportion of the surveyed sample also demonstrated awareness of the risks associated with BNPL usage. The study examined perceived risk from the user's perspective—specifically, the potential consequences that could negatively impact their well-being. Among the risk dimensions assessed, financial concerns such as falling into a debt trap, were the most frequently cited, confirming the emphasis placed on this issue by Krisnawati and Sam (2024), Bagniewski et al. (2024), and Gilbert and Scott (2023). Behavioural concerns, such as compulsive buying, highlighted by Bergheim et al. (2024), were also acknowledged, with over 65% of respondents agreeing that BNPL may contribute to such outcomes.

Importantly, perceived risk was shown to reduce user satisfaction and discourage recommendations of BNPL to others. These findings support H4. If, as Raj et al. (2024a) argue, perceived risk reduces consumers' intention to use BNPL, and if, as noted by Bagniewski et al. (2024), BNPL—like traditional credit—can lead to over-indebtedness, it follows that individuals are unlikely to recommend a service they themselves approach with caution. This reasoning is particularly relevant given that making a recommendation implies a degree of perceived trust and a positive orientation towards the recipient.

The findings concerning the role of perceived trust in shaping the intention to use BNPL services remain inconclusive. While Hidayat et al. (2024) confirmed the significance of this variable, Hoo et al. (2024) reported no significant effect of perceived trust on usage intention. In light of these discrepancies, perceived trust was examined at two distinct levels: perceived trust in the BNPL service itself and perceived trust in the institution offering it. Perceived trust in the BNPL service was relatively high, with at least 58% of respondents expressing no concerns about using it—and as much as 70% when the issue of personal data was excluded. Furthermore, no fewer than 63% of participants indicated perceived trust in the BNPL provider. The results suggest that perceived trust in the institution plays a more substantial role than perceived trust in the payment mechanism itself. Perceived trust in BNPL as a financial instrument does not exert a significant influence on consumers' willingness to recommend BNPL to friends or family. This supports H3A while leading to the rejection of H3B. These findings imply that consumers are more likely to base their evaluations on the perceived credibility and reputation of the BNPL provider than on the structure of the BNPL mechanism per se. Although the BNPL mechanism is relatively standardised and easy to comprehend—features that tend to be positively evaluated and contribute to overall satisfaction—the ultimate impact of these perceptions may be mediated by the consumer's view of the provider.

As shown by Dong et al. (2025), introducing mandatory reporting requirements for BNPL providers can alter consumer behaviour, lowering the rate at which BNPL transactions convert into long-term credit and reducing service usage among consumers with low creditworthiness. This emphasises the need to investigate whether consumer perceived trust in the provider is derived solely from the BNPL offer (e.g., transparency, communication, convenience, availability, data protection) or whether it also reflects the provider's broader image within the consumer finance landscape. When perceived trust is treated as a composite construct—without distinguishing whether it relates to the provider or the product—it significantly enhances the fulfilment of consumer expectations and increases their willingness to recommend BNPL services. This confirms H3. These findings are consistent with the prior literature (Raj et al., 2024b), which highlights the pivotal role of perceived trust in driving BNPL adoption.

This finding is particularly noteworthy due to the coexistence of two opposing forces: perceived risks associated with BNPL negatively affect consumer recommendation intention, yet perceived trust serves as a compensating factor that reinforces positive behavioural outcomes. This suggests that BNPL users, despite being aware of the inherent risks of the product, continue to perceive the benefits of deferred payment as outweighing the potential downsides. This aligns with previous research indicating that both perceived benefits and perceived risks influence consumers' intention to continue using mobile payment solutions (Adamek & Solarz, 2024). A similar dynamic appears to hold true for BNPL, where satisfaction and recommendation intention are shaped by a constellation of factors including perceived risk, perceived ease of use, and perceived usefulness.

5. Limitations and future research

Despite the rigour of this research, several limitations can be identified. Firstly, the study is constrained by the structure of its sample, which reflects only the distribution of age and gender, without incorporating other demographic characteristics such as income, education, or household composition. Another limitation is that the relationships that can be identified, particularly those related to attitudes and behaviours, are complex. Therefore, it is not always possible to clearly indicate their direction.

The conducted research demonstrates several possible directions for further research in the area addressed in this paper. One possibility would be to examine how perceived risk modifies the constructs derived from the TAM, and to what extent. Another approach would be to expand the models with new psychological constructs. Future research may also focus on extending the current model by examining how consumer satisfaction translates into repeat usage and increased engagement with BNPL services.

Conclusions

The findings confirm that the most significant factor affecting user satisfaction and their willingness to recommend BNPL is the perceived usefulness. Users predominantly perceive BNPL as a practical tool for managing household budgets and improving financial liquidity. Moreover, the absence of interest charges and the flexibility of repayment make BNPL an attractive alternative to traditional forms of purchase financing. In contrast, ease of use plays only a marginal role in shaping overall satisfaction, likely because attributes such as intuitiveness or transaction speed are now considered standard expectations for digital services.

Perceived trust proves to be a multidimensional construct. While perceived trust in the BNPL mechanism itself did not significantly influence the willingness to recommend BNPL service, perceived trust in the institution offering BNPL proved to be a decisive factor. In this study, respondents placed strong emphasis on the reputation, transparency, and credibility of financial service providers, suggesting that the brand and communication strategy of the provider may heavily influence the overall evaluation of BNPL service.

At the same time, perceived risk significantly reduced satisfaction—especially in relation to concerns about excessive debt, overspending, and a lack of understanding of how BNPL works. This indicates that even if BNPL service offers clear benefits, the accompanying concerns may effectively discourage users from recommending it to others. Perceived risk may induce cognitive dissonance, weakening the positive emotional response associated with using BNPL, thereby reducing overall satisfaction.

The study's results point to some theoretical and practical implications. Though its findings are in line with the key assumptions of TAM theory, some emphases are shifted. While perceived usefulness is a primary predictor of satisfaction and recommendation intention, ease of use plays a marginal role. This may mean that in a mature digital services ecosystem, usefulness is more important than the technical aspects of service, which influences how ease of use is perceived. The marginal impact of ease of use suggests that technology acceptance theory should consider a new phenomenon: a situation in which users take high service standards for granted, meaning this factor will not necessarily translate into their recommendation intention. This suggests a reinterpretation of TAM to adapt it to FinTech services.

The results here confirm the significant role of risk in the consumer decision-making process. As expected, consistent with TPR theory, perceived risk has a negative impact on satisfaction and subsequent BNPL recommendation. However, due to the specific risks (including behavioural and informational risks) inherent to this service, it hints that TPR may need to be adapted to more complex services, including BNPL.

The article also contributes to the development of trust theory in perceived financial services. The study reveals that perceived trust should be treated as a multidimensional construct, with individual dimensions having different weights in consumer behaviour. Therefore, it should not be treated as a single category. This reinforces the thesis that perceived security and reputation of a financial institution are stronger determinants of behaviour than the characteristics of the financial tool itself.

The study's findings also point to several practical implications. Above all, BNPL service providers should focus on building perceived trust and minimising perceived risk, for example, through transparent communication, implementing safeguards against over-indebtedness, and enhancing data protection. Therefore, information about payment terms should be published in a clear and easily understandable form. It would be undesirable to focus on information describing BNPL solely as a tool for improving budget liquidity or leading to more flexible budget management. The debt nature of this service should be clearly defined. Furthermore, a transparent fee and penalty policy is essential. Communication should be based on credibility and reputation, and building trust in the BNPL brand is as important as the user experience itself.

Taken together, the findings highlight that while BNPL offers considerable market potential, its continued success depends on the careful balancing of functional benefits with a responsible approach to trust and risk management. Ultimately, it is the perceived usefulness and the reputation of the provider that shape the consumer's positive experience—not merely technological ease of use.

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Sentiment and dividend smoothing: Do firms alter dividends during periods of high market activity?

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Abstract

This study investigates whether investor sentiment shapes dividend policy among publicly listed firms in Bangladesh by testing the hypothesis that firms alter their dividend smoothing practices in response to market optimism. We utilise a balanced panel of 116 firms from 2010 to 2021, applying robust panel regression techniques, including random effects, panel-corrected standard errors, and instrumental variable estimation to address model imperfections and potential endogeneity. Our findings show that, on average, firms increase dividends during periods of heightened investor optimism. However, this effect is moderated by prior dividend levels, indicating a tendency toward dividend smoothing. Firms appear to balance market sentiment with the need to maintain consistent payout signals. The findings contribute to the behavioural finance literature by highlighting sentiment as a key determinant of dividend behaviour within the Bangladesh context, where market volatility and retail participation are pronounced.

Keywords

- investor sentiment
- dividend policy
- dividend smoothing
- behavioural finance
- emerging markets

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Introduction

Dividend policy continues to be central to corporate finance because payouts perform multiples roles such as allocating cash, signalling inflation rate formation to the market, and mitigating agency conflicts. Classical perspectives emphasise fundamentals such as profitability, investment opportunities and financing frictions, while behavioural approaches argue that investor psychology can materially distort corporate incentives and market signals. Foundational empirical studies demonstrate that investor sentiment affects financing and payout choices across countries (Baker & Wurgler, 2006; Brav et al., 2005). More recent evidence refines this insight through new sentiment measures and transnational tests that reveal heterogeneity in how firms respond to market mood (Byun et al., 2021; Kumar & Sinha, 2024). Literature reviews and methodological overviews provide a comprehensive synthesis of themes and trends in the sentiment literature (Kamath et al., 2022, 2024; Maurya et al., 2025). Scholarly discourse on behavioural financial implications for macro and monetary analysis also highlight the broader relevance of sentiment for financial policy and firm behaviour (Willett, 2024). This prompts our core question: Does investor sentiment reshape the dynamics of dividend payouts and, specifically, the practice of dividend smoothing?

Empirical findings on the sentiment–dividend link are mixed and appear to be context-dependent. Baker and Wurgler (2006) demonstrate how sentiment alters capital-market incentives, which can translate into changes in payout policy, while catering theories posit that firms sometimes alter payouts to meet investor demand (Baker & Wurgler, 2004). Other studies find that firms conserve cash in booms to exploit favourable market valuations or to finance acquisitions instead of raising payouts (Ferris et al., 2009; Hoberg & Prabhala, 2008). Recent contributions also show that alternative sentiment proxies, such as online search intensity and news-based indices, predict corporate actions and market reactions (Belhoula et al., 2024; Duc et al., 2024; Qureshi, 2025). Bibliometric and systematic reviews further document the rapid growth of studies using social media, search intensity and machine-driven sentiment measures (Nyakurukwa & Patnaik, 2023; Prasad et al., 2023). Taken together, this evidence points to a non-uniform effect of sentiment on dividends that depends on investor composition, country institutions and the sentiment measure used (Santi & Zwinkels, 2023; Z. Wang, 2023).

A specific and under-tested question is whether dividend smoothing moderates sentiment's effect on payouts. The dividend smoothing literature, initiated by Lintner (1956) and reinforced by empirical studies (Brav et al., 2005; Leary & Michaely, 2011), documents managers' preference for gradual adjustments to preserve reputational credibility. Yet it is unclear whether smoothing serves to counteract sentiment-driven pressures or whether it relaxes under volatile sentiment when dividend signals lose informational value. Empirical work offers competing predictions: firms might cater and increase payouts, or they might withhold distributions to retain flexibility (Dong et al., 2005; Ferris et al., 2009). Studies on momentum, return predictability and the masking of fundamentals provide related evidence that dividend signals may be less informative in momentum-driven markets (Novy-Marx, 2012). Country-specific investigations into ownership structure and monitoring also suggest that shareholder concentration and state support can influence dividend choices under market stress (Kluzek & Schmidt-Jessa, 2022; Pieloch-Babiarz, 2021). Our paper empirically tests the moderating role of prior dividend commitments by interacting sentiment with lagged dividends, a direct test of whether smoothing attenuates sentiment-induced payout changes.

In markets like Bangladesh, retail investor participation dominates, and price movements often reflect sentiment rather than fundamentals. This characteristic elevates the need to explore whether investor sentiment affects firm-level decisions, including dividend payouts. While prior research has mostly focused on developed markets, the emerging market context presents a unique opportunity to examine the behavioural underpinnings of financial policy. Recent regional evidence and practitioner-led studies find evolving dividend practices and strong sentiment effects in South Asia and other emerging markets (Abor & Bokpin, 2010; Kumar & Sinha, 2024; Lubis et al., 2024). Systematic comparative research also shows that emerging economies display greater sensitivity to sentiment because of weaker institutional frameworks and limited institutional investor oversight (Aivazian et al., 2003; Mampouya, 2024). Together, these studies underscore why Bangladesh merits focused analysis.

The Dhaka Stock Exchange (DSE) is a highly informative setting to examine these mechanisms. The DSE features a pronounced retail investor presence, episodic sentiment-driven swings, and a regulatory framework that has historically lagged major developed markets; all these factors amplify behavioural channels and make managerial responses to sentiment more salient. Regional and comparative evidence suggests that emerging markets display stronger sentiment effects on firm behaviour because of weaker investor protections and lower institutional ownership (Abor & Bokpin, 2010; Aivazian et al., 2003; Kumar & Sinha, 2024; Lubis et al., 2024). Recent regional studies and practitioner reports also document evolving dividend practices in South Asia and similar markets, reinforcing the need for evidence from Bangladesh (Kumar & Sinha, 2024; Lubis et al., 2024). These institutional features motivate our empirical focus on the DSE and on how smoothing operates when sentiment is elevated.

The empirical design uses a panel dataset of 116 publicly listed firms across 14 sectors spanning the period 2010 to 2021. We construct a balanced panel with 1,392 firm-year observations. The primary dependent variable is cash dividend payout, while investor sentiment is measured using the Trading Volume Ratio (TVR). The model also includes an interaction term between TVR and lagged dividend payments to capture the moderating role of dividend history. Control variables include key firm characteristics such as size, profitability, leverage, and age, alongside macroeconomic factors like GDP growth rate, inflation rate, unemployment rate, and real interest rates. This comprehensive specification helps isolate the unique role of sentiment while controlling for standard determinants of dividend policy. The design and variable choices are informed by recent methodological overviews and empirical studies on sentiment measurement and dividend modelling (Kamath et al., 2022; Maurya et al., 2025; Prasad et al., 2023).

To ensure methodological rigour, we begin with diagnostic tests that assess multicollinearity, heteroskedasticity, autocorrelation, and cross-sectional dependence. We estimate Random Effects models, as both are appropriate under these conditions (Baltagi, 2005). Following the recommendations made by Beck and Katz (1995) for dealing with panel data exhibiting both cross-sectional dependence and heteroskedasticity, we further validate our findings using the Panel-Corrected Standard Errors (PCSE) estimator. These empirical techniques are essential for controlling for latent firm-level heterogeneity and macroeconomic shocks, factors especially relevant in frontier markets like Bangladesh (Bissoondoyal-Bheenick et al., 2022; W. Wang et al., 2022).

Our findings indicate that investor sentiment exerts a significant influence over firms' dividend payout policy. Specifically, we find a positive and statistically significant relationship between Trading Volume Ratio and cash dividends, suggesting that during high-sentiment periods, firms tend to increase dividend payouts, at least on average. This behaviour aligns with the notion that optimistic market environments encourage firms to distribute their earnings. However, the inclusion of the interaction term between TVR and lagged dividends implies that this effect does not apply to firms that made no dividend payments in the previous period. In their case, it is even less probable that they will distribute their earnings during a period of market optimism. This duality reflects a nuanced strategy: while many firms respond to buoyant sentiment with increased payouts, some are even more inclined to remain conservative. These patterns echo previous findings on sentiment-driven corporate behaviour and its implications for volatility and firm policies (Berger, 2022; Gao et al., 2022; Huynh et al., 2021).

To reinforce the robustness of our conclusions, we estimate a 2SLS model using lagged trading volume ratio as an instrumental variable to account for potential endogeneity in the sentiment-dividend nexus. Diagnostic tests such as the Kleibergen-Paap LM statistic and Hansen's *J* test confirm the validity and robustness of our instrument. The 2SLS results reaffirm our earlier findings: investor sentiment exerts a positive effect on dividend payouts, but some firms moderate this effect through dividend smoothing practices. These findings enhance the causal interpretation of our model and reinforce the argument that even in sentiment-driven markets, firms employ conservative payout strategies to manage long-term investor expectations, mitigating the potential volatility induced by transitory market moods (Gaies et al., 2022; Goel & Dash, 2022).

This paper contributes to the dividend literature by situating payout policy within the broader framework of investor sentiment and financial decision-making in frontier markets. Our findings align with growing evidence that sentiment measures shape financial decisions, including dividends (Belhoula et al., 2024; Duc et al., 2024; Kamath et al., 2024; Maurya et al., 2025; Qureshi, 2025; Z. Wang, 2023; Willett, 2024); we extend this literature by demonstrating how payout policy in frontier markets functions as both a behavioural response and a governance tool. While earlier works highlight the importance of sentiment dynamics in driving stock returns and volatility (Bissoondoyal--Bheenick et al., 2022; Nyakurukwa & Patnaik, 2023; Prasad et al., 2023; Santi & Zwinkels, 2023; W. Wang et al., 2022), our evidence emphasises that dividends themselves can transmit and absorb sentiment shocks. Moreover, we show that in weaker institutional contexts, dividends act as instruments to mitigate agency conflicts and enhance credibility (Kumar & Sinha, 2024; La Porta et al., 2000; Lubis et al., 2024; Pieloch-Babiarz, 2021), complementing studies on how corporate payout behaviour adapts under extraordinary conditions, such as policy interventions and crises (Goel & Dash, 2022; Huynh et al., 2021; Kluzek & Schmidt-Jessa, 2022). By focusing on frontier economies, this study diversifies the geographical scope of dividend research, which remains dominated by advanced and large emerging markets (Bekaert & Harvey, 2003; Kumar & Sinha, 2024). Moreover, we bridge gaps identified in bibliometric surveys by showing that integrating behavioural finance perspectives into dividend policy reveals dynamics obscured in traditional models (Berger, 2022; Gaies et al., 2022; Gao et al., 2022; Kamath et al., 2022).

The remainder of the paper is organised to present the study's arguments and evidence in a clear and logical sequence. Section 1 reviews the existing literature on investor sentiment, dividend policy, and dividend smoothing, and uses these insights to develop the research hypotheses. Section 2 describes the data and methodology, outlining the construction of key variables and explaining the econometric approach used to estimate the interaction effects. Section 3 reports the empirical results, beginning with baseline models and then examining how sentiment interacts with firms' dividend histories. Section 4 discusses these findings in relation to prior research, noting where our results align with or depart from earlier studies. Last section concludes by summarising the contributions of the paper and briefly noting the study's limitations and implications for future research.

1. Literature review and hypothesis

Investor sentiment is commonly understood as a non-fundamental component of asset pricing that reflects investors' collective beliefs, moods, and demand pressures rather than firms' cash-flow fundamentals. Recent systematic reviews and bibliometric overviews treat sentiment as a multi-dimensional construct measured by trading-based proxies, attention and search intensity, news- and social-media text indices, survey-based mood measures, or composite indices that combine several signals (Belhoula et al., 2024; Duc et al., 2024; Maurya et al., 2025; Nyakurukwa & Patnaik, 2023; Qureshi, 2025). Empirically, scholars use relative Trading Volume Ratios and turnover to capture short-term enthusiasm (a liquidity/attention channel); Google Trends or search intensity serve as attention proxies, and text-based sentiment extracted from news or social media captures information flow and affect. These approaches reflect different theoretical channels: attention-driven price pressure, noise trading and mispricing, and information-processing limits, each of which can shape corporate decisions differently (Baker & Wurgler, 2006; Santi & Zwinkels, 2023; Z. Wang, 2023).

A rapidly growing body of literature links investor sentiment to firm-level decisions beyond returns, including financing, investment, and payout choices. Recent frontier- and emerging-market studies show that sentiment measures predict dividend and payout behaviour (Kumar & Sinha, 2024; Lubis et al., 2024), while cross-country work documents strong heterogeneity in sentiment effects, depending on investor protection frameworks and institutional settings (Belhoula et al., 2024; Byun et al., 2021). Methodologically, researchers employ panel estimators with fixed effects, event-study frameworks, attention-based regressions using search intensity, and machine-text sentiment regressions; several recent papers also use instrumental variables and Panel-Corrected Standard Errors to address endogeneity and cross-sectional dependence (Maurya et al., 2025; W. Wang et al., 2022). Collectively, these studies indicate that sentiment often matters for corporate choices but that the direction, magnitude, and persistence of effects vary by market, investor composition, and measurement approach (Bissoondoyal-Bheenick et al., 2022; Gao et al., 2022; Huynh et al., 2021). This broader behavioural perspective connects directly with modern developments in dividend catering theory, which also seeks to explain how market sentiment shapes managers' payout decisions.

Dividend catering theory has evolved from the original catering intuition to more sophisticated empirical tests of when and how managers cater to investor demand. Recent multi-country tests find that catering incentives vary with investor sentiment, legal protections and market liquidity, and may be stronger when sentiment is low or when catering reduces mispricing (Byun et al., 2021; Ferris et al., 2009; Kluzek & Schmidt-Jessa, 2022). Empirical extensions incorporate determinants such as investor demand captured by search intensity, ownership concentration, institutional investor presence, and macroeconomic uncertainty (Belhoula et al., 2024; Pieloch-Babiarz, 2021; Qureshi, 2025). These studies show mixed results: some document that firms increase payouts to satisfy investor preference for dividends, while others find that firms conserve cash during booms to exploit favourable valuations rather than increasing payouts (Dong et al., 2005; Hoberg & Prabhala, 2008). The diverging evidence suggests that catering behaviour is conditional on governance, investor base, and financing opportunities, which necessitates testing both catering and conservation channels in a single specification.

The dividend smoothing literature emphasises that managers prefer stable payouts to signal firm quality and reduce investor uncertainty (Brav et al., 2005; Lintner, 1956). More recent work quantifies determinants of smoothing and conditions under which smoothing weakens or strengthens (Leary & Michaely, 2011; Michaely & Roberts, 2012). Studies linking sentiment to smoothing find two competing mechanisms. On one hand, heightened sentiment may reduce the informational value of dividends, weakening smoothing incentives when investors rely more on mood and momentum than on fundamental signals (Gao et al., 2022; Novy-Marx, 2012). On the other, managers operating in sentiment-rich environments may smooth more aggressively to reassure investors and preserve reputations, particularly when retail participation is high (Kumar & Sinha, 2024; Lubis et al., 2024). This tension motivates our explicit test of whether smoothing moderates the direct effect of sentiment on payouts.

Recent empirical work has diversified the analytical toolkit for studying sentiment and corporate policy. Scholars combine traditional panel estimators with robust corrections for cross-sectional dependence (PCSE, GLS), employ two-stage least squares using lagged or external instruments, and exploit high-frequency attention proxies such as Google Trends or Twitter-derived sentiment (Duc et al., 2024; Maurya et al., 2025; Qureshi, 2025). Bibliometric and systematic reviews summarise these methodological trends, noting the rise of machine-learning text sentiment, network-connectedness approaches, and cross-market spillover analyses (Kamath et al., 2024; Nyakurukwa & Patnaik, 2023; Prasad et al., 2023). This methodological plurality suggests that robust inference requires triangulation across estimators and proxies, a point we adopt in our own empirical design.

Emerging-market studies emphasise the role of institutions and ownership structure in shaping payout responses to sentiment. Evidence from Asia and Africa indicates that weaker investor protections and higher retail participation amplify sentiment effects on corporate policies (Abor & Bokpin, 2010; Aivazian et al., 2003; Kumar & Sinha, 2024; Lubis et al., 2024). Country-specific studies using Google Search and social media proxies find short-lived but economically meaningful attention-driven price pressure in frontier markets such as Vietnam and other South-East Asian exchanges (Duc et al., 2024). State aid and pandemic-era interventions also altered dividend choices, illustrating how policy regimes interact with sentiment to shape payout outcomes (Kluzek & Schmidt-Jessa, 2022). These lines of evidence demonstrate that institutional context is central to interpreting sentiment–dividend linkages.

Despite many new studies, the literature still exhibits three key gaps that our paper addresses. Firstly, while catering and conservation channels are both proposed, few papers jointly estimate the direct sentiment effect and the moderating role of ex-post smoothing within one coherent framework, especially in frontier markets (Byun et al., 2021; Kumar & Sinha, 2024). Secondly, few studies triangulate across estimators that correct for cross-sectional dependence and apply instrumental variables to probe causality in panel settings with episodic sentiment shocks (Belhoula et al., 2024; Maurya et al., 2025). Thirdly, frontier markets such as Bangladesh remain understudied despite evidence that Google-Search-, trading-volume- and local news sentiment effects are strong in comparable markets (Duc et al., 2024; Qureshi, 2025). By testing whether prior dividend commitments attenuate sentiment's effect on payouts and by applying a suite of robust estimators, this paper fills these gaps and helps reconcile divergent empirical findings.

Motivated by the reviewed evidence, we adopt two precise hypotheses:

H 1: An increase in investor sentiment reduces contemporaneous cash dividend payouts, ceteris paribus.

H 2: The negative effect of sentiment on dividends is attenuated for firms with stronger prior dividend commitments.

2. Methodology

We investigate whether heightened investor sentiment influences firms' dividend declaration decisions. To address this question, we draw on a sample of 116 listed companies covering 14 distinct sectors of the Dhaka Stock

Exchange. We selected these companies based on a convenient sampling technique. For our study, we eliminate companies in the financial industry, adhering to market conventions that often differentiate them regarding dividend distribution and stability. This omission helps preserve consistency when evaluating broader industrial sectors, where dividend smoothing is generally more evident. We consider companies listed on the Dhaka Stock Exchange, since this is Bangladesh's primary exchange, which covers the majority of the publicly listed companies. Moreover, companies listed on the DSE are the most representative of the broader corporate landscape in Bangladesh. Our sample period covers the years 2010 to 2021 and is based on data availability and its relevance in understanding the dynamic of dividend policy under significant economic and market developments, including the post-global financial crisis and COVID pandemic. This strategy yields a balanced panel comprising a total of 1,392 firm-year observations. We collect company-specific and macroeconomic data from the annual reports of the respective companies and World Bank indicators, respectively. Outliers were addressed through winsorisation, where observations below the 1st percentile and above the 99th percentile were replaced with the corresponding percentile values. This approach mitigates the impact of extreme observations without excluding data.

2.1. Variables

We incorporate one dependent variable, one key independent variable, and one moderating variable, along with four firm-specific control variables and four macroeconomic control variables. Trading Volume Ratio (TVR) was selected as the proxy of investor sentiment due to its widespread use in the literature. This proxy also covers the individual investor sentiment. Several studies (Baker & Wurgler, 2007; Haritha & Rishad, 2020; Schmeling, 2009) found that TVR can serve as a reliable indicator of the sentiment, as increased trading volume often correlates with heightened investor optimism. Though alternative sentiment proxies such as Consumer Confidence Index, Volatility Index can be used, they may not capture the specific behaviour of the individual investors in emerging countries like Bangladesh. In our analysis, we include several control variables identified in the literature as significant determinants of dividend policy. These variables encompass firm-specific characteristics such as size, profitability, leverage, and age, as well as macroeconomic factors including GDP growth, inflation rate, unemployment rate, and real interest rates. The inclusion of these variables allows for a comprehensive examination of the factors influencing dividend smoothing practices among firms listed on the Dhaka Stock Exchange.

In Table 1 we provide detailed definitions, operationalisation and references of each variable included in the analysis:

Table 1. Description of variables

Role	Variable	Description	Measurement	References
Dependent variable	cash dividend	measures the firm's actual dividend paid;	cash dividend percentage disbursed by the company	
Independent variable	trading volume ratio	proxy for investor sentiment — trading volume ratio	dividing trading volume by number of shares outstanding	Baker & Wurgler (2006); Schmeling (2009)
Moderator / interaction term	trading volume ratio × lagged cash dividend	captures how sentiment affects the tendency to smooth or alter dividends		
Control variables	firm size	proxy of firm size as measured by the assets	natural logarithm of total assets	Bon & Hartoko (2024)
	profitability	the profit that company generates from their core operation	ratio of operating profit to total assets	Wahyuni & Peride (2021)
	leverage	the degree to which firm relies to debt capital	ratio of total debt to total assets	Ali et al. (2015)
	age	age of the firm since inception to date	natural logarithm of age of the firm since inception to date	Benjamin & Tenai (2018)
	GDP growth	annual GDP growth rate	percentage changes of GDP growth rate	Romus et al. (2020)
	inflation rate	inflation rate	annual inflation rate	Osman et al. (2024
	unemployment rate	unemployment rate	annual average unemployment rate	Mahirun et al. (2023)
	real interest rate	interest rate after accounting for inflation	interest rate minus inflation rate	Hasan et al. (2022)

Source: own work.

2.2. Model of the study

We specify the following empirical model to examine the impact of investor sentiment on firms' dividend declaration decisions. We delineate both of the following equations based on our two hypotheses.

$$\begin{aligned}\text{CASH DIVIDEND}_{i,t} &= \beta_0 + \beta_1 TVR_{i,t} + \beta_2 FIRMSIZE_{i,t} + \beta_3 OPPROFIT\_TA_{i,t} + \\ &+ \beta_4 DEBT\_ASSETS_{i,t} + \beta_5 LNAGE_{i,t} + \beta_6 GDP_{i,t} + \beta_7 INF_{i,t} + \beta_8 UNEMP_{i,t} + \\ &+ \beta_9 REALINT_{i,t} + \varepsilon_{i,t} \end{aligned} \tag{1}$$

$$\begin{aligned}\text{CASH DIVIDEND}_{i,t} &= \beta_0 + \beta_1 TVR_{i,t} + \beta_2 \text{CASH DIVIDEND}_{i,t-1} + \\ &+ \beta_3 TVR_{i,t} \times \text{CASH DIVIDEND}_{i,t-1} + \beta_4 FIRMSIZE_{i,t} + \beta_5 OPPROFIT\_TA_{i,t} + \\ &+ \beta_6 DEBT\_ASSETS_{i,t} + \beta_7 LNAGE_{i,t} + \beta_8 GDP_{i,t} + \\ &+ \beta_9 INF_{i,t} + \beta_{10} UNEMP_{i,t} + \beta_{11} REALINT_{i,t} + \varepsilon_{i,t} \end{aligned} \tag{2}$$

In these equations, the subscripts i and t denote firm i in year t, respectively. All variable definitions and notations are provided in Table 1. The term ε captures the error component.

2.3. Diagnostic tests

The study uses a series of diagnostic tests to ensure the robustness of our panel data analysis and to determine the appropriate model specification. These tests are presented in the Appendix. We assess the presence of multicollinearity among the explanatory variables using the Variance Inflation Factor (VIF). The mean VIF value is 1.65, which falls below the commonly accepted threshold of 10. We test for heteroskedasticity using the Breusch-Pagan/Cook-Weisberg test. The test yields a Chi-square statistic of 169.14 with a p-value near 0.000, indicating the presence of heteroskedasticity. We employ the Wooldridge test for autocorrelation in panel data to examine the presence of serial correlation in the error terms. The test produces an F-statistic of 0.014 with a p-value of 0.9055. Thus, we find no evidence of autocorrelation. We apply the Pesaran test to examine the presence of cross-sectional dependence. The test returns a p-value close to 0.0000, suggesting that cross-sectional dependence is present in the dataset. To determine the appropriate estimation technique between Fixed Effects and Random Effects Models, we conduct the Hausman specification test. The test yields a Chi-square statistic of 5.69 with 9 degrees of freedom and a p-value of 0.7732, which justifies the choice of the Random Effects Model.

2.4. Model specification

To ensure the reliability of our regression estimates, we employ both the Random Effects (RE) Model and Panel Corrected Standard Errors (PCSE) approach. The choice of these models is guided by the results of diagnostic tests, which confirm the presence of cross-sectional dependence and heteroskedasticity in the dataset. We employ the PCSE model following the recommendations made by Beck and Katz (1995), who advocate for the use of Panel Corrected Standard Errors (PCSE) in the presence of heteroskedasticity and cross-sectional dependence.

3. Results

3.1. Descriptive statistics

Table 2 represents the descriptive statistics of the study, highlighting the mean, median, first quartile, third quartile and standard deviation of each

Table 2. Descriptive statistics

Variable	Mean	Median	1st quartile (Q1)	3rd quartile (Q3)	Standard deviation
Cash dividend	0.3986	0.1000	0.0000	0.2200	1.2186
Trading Volume Ratio	0.2682	0.0037	0.0012	0.0105	6.3866
Firm size	9.3770	9.3446	8.8393	9.8733	0.7811
Profitability	0.0656	0.0440	–0.0002	0.0934	2.2720
Leverage	1.9759	0.2317	0.0570	0.4357	29.3797
Age	27.2400	2.5200	19.0000	40.0900	0.2408
GDP growth	0.0634	0.0650	0.0605	0.0697	0.0108
Inflation rate	0.0661	0.0615	0.0560	0.0730	0.0119
Unemployment rate	0.0449	0.0444	0.0415	0.0470	0.0047

Note: This table presents the descriptive statistics for the variables used in the analysis. The statistics reported include the mean, median, first quartile (Q1), third quartile (Q3), and standard deviation. Cash dividend and Trading Volume Ratio represents firm-level payout and market activity, respectively. Firm size, profitability, leverage, and age are firm-specific characteristics. GDP growth rate, inflation rate, and unemployment rate capture macroeconomic conditions. The values reflect the distribution of observations across the full sample period. The quartile measures indicate the spread of the data around the central tendency, while the standard deviation captures variability within each variable.

Source: own calculation.

variable. The average cash dividend paid by the firms is around 39.86%, with a median of 10%. The standard deviation of the variable is very high, which indicates that many firms either pay no dividends or maintain low dividends, consistent with dividend-smoothing behaviour. The Trading Volume Ratio, serving as an indicator of investor sentiment, registers a mean of 0.2682 and a very high standard deviation of 6.3866, indicating irregular surges in market activity characteristic of sentiment-driven trading in emerging markets. The distribution of firm size is somewhat symmetrical around the mean. Profitability (mean 0.0656; SD 2.2720) and leverage (mean 1.9759; SD 29.3797) exhibit significant variability. The average firm age is 27 years, implying a relatively mature sample of listed companies. Macroeconomic variables such as GDP growth (6.34%), inflation (6.61%), and unemployment (4.49%) are stable with low variability, representing the broader macro environment during the period.

3.2. Regression results

Table 3 shows the results based on Random Effect Model (RE) and Panel Corrected Standard Error Model under hypothesis 1. The results show that the Trading Volume Ratio has a positive and significant effect on cash dividends in the PCSE model, which indicates that firms generally pay higher dividends when market activity is high. It contradicts our first hypothesis. Additionally, firm size consistently shows a strong positive effect, meaning larger firms pay more dividends. Profitability, leverage, and firm age also register positive and significant relationships in the PCSE model, suggesting that financially stable and mature firms distribute higher dividends. GDP growth shows a negative and significant effect, indicating that firms reduce dividends when economic conditions improve, possibly to retain earnings for investment. Inflation, unemployment, and real interest rates do not show significant effects, which means these macroeconomic factors do not strongly influence dividend decisions. Overall, the PCSE model provides better explanatory power, and the results confirm that both market sentiment and firm characteristics shape dividend policy.

Table 4 shows the results based on both models for hypothesis 2. We follow Brambor et al. (2006) and add an interaction term between the Trading Volume Ratio and lagged dividends. After including the interaction term, the parameter close to trading volume ratio become negative. This indicates that companies that paid no dividends in the previous period are even less likely to pay them when the Trading Volume Ratio increases in the current period. At the same time, the interaction term is positive and significant in the Random Effects Model and positive but weaker in the PCSE model. This suggests that the negative effect of sentiment on dividend payouts becomes less negative for firms with higher prior dividends. Therefore, when firms have historical

Table 3. Assessment of the impact of trading volume ratio on the cash dividend

	Random Effects Model		PCSE Model	
Trading Volume Ratio	0.0015 (0.0029)	0.0008 (0.0030)	0.0009*** (0.0003)	0.0010*** (0.0004)
Firm size		0.2233*** (0.0834)		0.2166*** (0.0714)
Profitability		0.0100 (0.0098)		0.0042** (0.0021)
Leverage		0.0004 (0.0009)		0.0007** (0.0004)
Age		0.5319* (0.3083)		1.0324*** (0.3077)
GDP growth		−2.5465 (1.8976)		−2.4613*** (0.6958)
Inflation rate		−2.2828 (2.4993)		−1.7499 (1.1771)
Unemployment rate		1.9339 (7.0097)		−2.8270 (5.2156)
Real interest rate		0.2181 (0.3676)		0.2208 (0.1356)
Intercept	0.3982 (0.0962)	−2.2420 (0.8936)	0.4070 (0.1083)	−2.7110 (0.5428)
R-Square (%)	0.8	17.11	5.6	22.11
Prob > chi2	0.6120	0.0000	0.0004	0.0000
Companies	116	116	116	116
Observations	1,392	1,392	1,392	1,392

Note: This table presents panel regression estimates assessing the influence of investor sentiment, proxied by the Trading Volume Ratio, on firms' cash dividend payments. Coefficients are reported with standard errors in parentheses; ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. The Random Effects Model provides baseline estimates accounting for unobserved firm heterogeneity, while the PCSE specification corrects for heteroskedasticity and contemporaneous correlation across firms. Overall model fit improves noticeably under PCSE, and Chi-square statistics confirm strong joint significance of the explanatory variables.

Source: own calculation.

dividend commitments, they are less likely to reduce dividends even when market sentiment is high. This directly supports our hypothesis 2, as the results show that the adverse effect of investor sentiment on dividends is limited to companies with no dividends history and is attenuated by stronger prior dividend commitments.

For the control variables, firm size shows a weak positive effect, while profitability, leverage, and age show no consistent influence on dividends. GDP growth has a negative and significant effect, meaning firms reduce dividends when economic conditions improve. Inflation, unemployment, and real interest rates do not show significant effects. Overall, the models indicate that sentiment may reduce dividends, but firms with stronger historical payout commitments moderate this response, aligning closely with the proposed hypothesis.

Table 4. Results for the interaction between Trading Volume Ratio and cash dividend

	Random Effects Model	**PCSE Model**
Trading Volume Ratio	–0.0094*** (0.0031)	–0.0087* (0.1121)
Cash dividend$_{t-1}$	0.9200*** (0.0163)	0.9509*** (0.0553)
Trading Volume Ratio × Cash dividend$_{t-1}$	0.1021*** (0.0180)	0.0945 (0.1065)
Firm size	0.0297* (0.0220)	0.0218 (0.0539)
Profitability	–0.0003 (0.0079)	–0.0007 (0.0021)
Leverage	0.0000 (0.0006)	0.0000 (0.0003)
Age	0.0340* (0.0746)	0.0074 (0.1551)
GDP growth	–4.1407* (1.7705)	–4.1592*** (0.9759)
Inflation rate	–2.6966 (2.1872)	–2.6287 (1.4699)
Unemployment rate	–7.2179 (6.3427)	–6.9015 (5.9369)
Real interest rate	0.2889 (0.6084)	0.3472 (0.1410)
Intercept	0.4889 (0.93)	0.5735 (0.5417)
R^2	78.03%	81.45%
Prob > chi2	0.0000	0.0000
Companies	116	116
Observations	1,392	1,392

Note: This table reports the results of Random Effects (RE) and Panel-Corrected Standard Errors (PCSE) regressions examining how investor sentiment influences firms' cash dividend payments. Investor sentiment is proxied by the Trading Volume Ratio. Coefficients are reported with standard errors in parentheses, and ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively. The PCSE Model demonstrates higher explanatory power, and Chi-square statistics confirm overall model significance.

Source: own calculation.

Following Brambor et al. (2006), we create Figure 1 that shows the marginal effect of the Trading Volume Ratio on cash dividends across the observed range of lagged cash dividends. The solid blue line denotes this marginal effect, which is consistently positive and increases steadily. The shaded area represents the 95% confidence interval. In the case of the Random Effects Model, the entire interval lies above zero for all positive values of the moderator, which means that the marginal effect of the trading volume ratio is statistically significant and positive across the full range. This is consistent with results presented in Table 3.

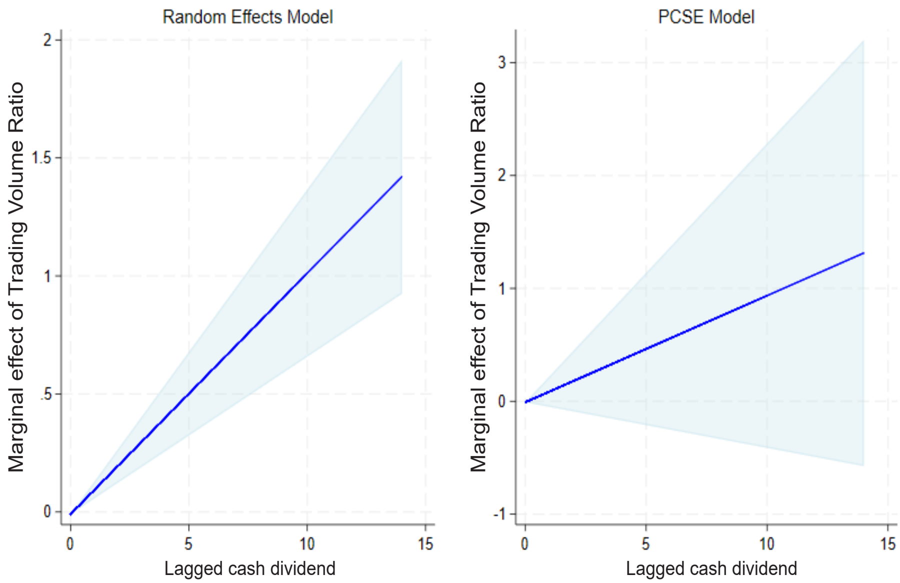

Figure 1. Marginal effect of Trading Volume Ratio

Note: The graphs use the random effects and PCSE models from Table 4. Lagged cash dividend ranges from 0 to 14 (although the 3rd quartile is equal to 0.2200). The confidence interval is 95% in both graphs.

Source: own work.

3.3. Endogeneity correction

To address potential endogeneity concerns related to investor sentiment in shaping dividend policy, we implement a two-stage least squares (2SLS) estimation strategy. Specifically, we employ Trading Volume Ratio using its one-period lag, Trading Volume Ratio$_{t-1}$, under the assumption that past sentiment influences current sentiment but is unlikely to be directly correlated with contemporaneous dividend shocks. This instrumental variable approach helps mitigate reverse causality and omitted variable bias that may arise in the baseline regressions. The model retains the same set of firm-specific and macroeconomic control variables as specified in Equation (2).

Table 5 reports the diagnostic tests confirming the validity and strength of the chosen instrument. The Kleibergen-Paap LM statistic rejects the null of under identification, indicating that the instrument is relevant and meaningfully correlated with the endogenous regressor. The Cragg-Donald F-statistic exceeds the 10% Stock-Yogo critical value, suggesting that the instrument is not weak and that the estimates are unlikely to suffer from weak instrument bias. Additionally, the Hansen J test for overidentifying restrictions yields a p-value of 0.35, indicating that the instrument is valid and exogenous, as we fail to reject the null hypothesis of instrument orthogonality. Taken together, these results confirm that the lag of trading volume ratio is a statistically sound instrument for addressing potential endogeneity in the sentiment-dividend relationship.

Table 5. Instrument Validity Test

Test	Statistic	p-value / critical value	Interpretation
Under identification (Kleibergen-Paap LM)	15.732	0.0001	instrument is relevant
Weak Identification (Cragg-Donald F)	25.814	10% critical value = 16.38	strong instrument
Hansen J (Overidentification)	0.872	0.35	instruments are valid

Note: This table reports the diagnostic tests used to evaluate the validity of the instrumental variables employed in the regression analysis. The tests presented include the Kleibergen-Paap LM statistic for under identification, the Cragg-Donald F-statistic for weak identification, and the Hansen J test for overidentification. For each test, the corresponding test statistic and either the p-value or the relevant critical value is reported.

Source: own calculation.

Table 6 presents the findings from the Two-Stage Least Squares (2SLS) regression, indicating substantial connections among investor sentiment, dividend smoothing, and other firm-specific and macroeconomic variables. The Trading Volume Ratio demonstrates a negative and statistically significant impact on dividend payment behaviour, suggesting that firms with no prior dividends are inclined to reduce payments during times of increased market optimism. The interaction term between the Trading Volume Ratio and lagged cash dividend is both positive and significant, indicating that firms with high prior dividends are more inclined to increase dividends during periods of high sentiment. The lagged cash dividend variable has significant persistence, confirming the influence of previous payment on present dividend determinations.

Table 6. Results of 2SLS Regression

Variable	Coefficient	Robust standard error	z-value	p-value
Trading Volume Ratio	–0.0125	0.0050	–2.50	0.012
Cash dividend$_{t-1}$	0.9163	0.0881	10.40	0.000
Trading Volume Ratio × Cash dividend$_{t-1}$	0.1132	0.0400	2.83	0.005
Firm size	0.0308	0.0226	1.36	0.174
Profitability	–0.0002	0.0025	–0.08	0.933
Leverage	0.0000	0.0002	0.26	0.792
Age	0.0356	0.1199	0.30	0.767
GDP growth	–4.5650	2.6547	–1.72	0.086
Inflation rate	–3.9958	2.1119	–1.89	0.058
Unemployment rate	–9.9214	8.5527	–1.16	0.246
Real interest rate	0.4861	0.1952	2.49	0.013
Intercept	0.6971	0.4808	1.45	0.147
Centred R^2 (%)	78.05			
Prob > F	0.0000			
Companies	116			
Observations	1,392			

Note: This table reports the results of the Two-Stage Least Squares (2SLS) regression estimation. For each explanatory variable, the table presents the coefficient estimate, the corresponding robust standard error, the z-value, and the p-value. The interaction term between the Trading Volume Ratio and lagged cash dividend is included. Control variables consist of firm-specific characteristics and macroeconomic indicators, and an intercept term is also reported. Model fit is summarised using the centred R-squared, while the overall significance of the model is indicated by the Prob > F statistic.

Source: own calculation.

4. Discussion

Our empirical results show that elevated investor sentiment, proxied by the Trading Volume Ratio, is associated, on average, with an increase in contemporaneous cash dividends. However, the interaction between sentiment and lagged dividend indicates that this positive effect applies only to firms with a history of a previous payout. Conversely, firms that did not pay dividends in a preceding period exhibit even lower propensity to pay dividends during periods of elevated investor sentiment. These patterns resonate with recent empirical work in emerging and frontier markets that reports a conservation or reallocation motive during sentiment booms rather than unconditional catering. For example, Kumar and Sinha (2024) and Lubis et al. (2024) document that firms in India and comparable South Asian markets often conserve cash or reallocate funds when market optimism is high, rather than increasing immediate payouts. Our finding therefore aligns with the conservation channel these authors emphasise, and it expands their scope by showing that smoothing weakens the immediate impact of sentiment on payouts.

At the same time, our results reconcile seemingly divergent findings in the literature. Several papers argue that managers cater to investor demand and increase dividends when sentiment is favourable (Baker & Wurgler, 2004; Ferris et al., 2009), while others document that firms exploit market booms by issuing equity or retaining cash for investment (Gao et al., 2022; Hoberg & Prabhala, 2008). By explicitly modelling the interaction between sentiment and prior dividend levels, we show both patterns can coexist. Managers increase payouts on average during sentiment spikes, yet firms without a payout reputation do not change their behaviour during the period of heightened sentiment. This reconciliatory view helps explain why cross-study comparisons sometimes produce conflicting conclusions: differences in investor composition, governance, and the prevalence of firms with established payout reputations can tip the observed net effect toward catering or conserving.

Methodologically, our use of panel estimators robust to cross-sectional dependence (GLS, PCSE) and the 2SLS check with lagged-TVR instruments aligns with the best practice emerging in the recent sentiment literature (Belhoula et al., 2024; Maurya et al., 2025). Many prior studies rely on single-estimator approaches or do not fully account for cross-country common shocks; by triangulating estimators, we reduce the chance that our results are driven by estimator choice or common-factor omitted variables. This methodological robustness is especially important in frontier markets, where episodic shocks and connectedness can bias naive panel estimates (Bissoondoyal-Bheenick et al., 2022; Z. Wang et al., 2022).

Institutional context matters for interpreting our findings. The Dhaka Stock Exchange, with high retail participation and weaker institutional monitoring relative to developed exchanges, is predisposed to stronger short-term sentiment effects (Abor & Bokpin, 2010; Aivazian et al., 2003). Our evidence complements country-specific work showing attention-driven price pressure in frontier exchanges (Duc et al., 2024; Qureshi, 2025), and pandemic-era and policy-distortion studies that altered payout policies in other emerging contexts (Kluzek & Schmidt-Jessa, 2022). Thus, the signs of dividend smoothing observed in our sample should be read alongside these institutional features: smoothing appears to be a deliberate stabilising strategy adopted by managers facing volatile retail-driven sentiment and limited institutional disciplining.

Beyond empirical alignment, our study contributes conceptually by showing how smoothing functions as a moderating mechanism rather than a rival explanation to catering. While catering theory suggests firms may increase payouts to satisfy investor tastes, our results show that smoothing can dampen such responses when managers prioritise long-term reputational capital and financial discipline. This complements recent theoretical and empirical contributions that emphasise the conditional nature of catering incentives, along with the role of governance and market structure in determining payout responses (Byun et al., 2021; Lubis et al., 2024; Pieloch-Babiarz, 2021).

Finally, our findings have observable implications for market participants. Investors interpreting dividend announcements should incorporate prevailing sentiment measures and firms' payout histories into their valuation models, because a dividend cut during a sentiment spike may reflect either a prudent smoothing response or a signal of reallocation that implies different valuation adjustments. Regulators could likewise use disclosure requirements to reduce noise trading and to highlight management's explanations of payout changes during volatile sentiment episodes. These practical implications echo calls in recent studies for improved transparency and investor education in sentiment-prone markets (Maurya et al., 2025; Willett, 2024).

Conclusions

We examine whether investor sentiment shapes corporate dividend policy in a frontier-market setting and whether dividend smoothing moderates that relationship. Using a panel of 116 firms listed on the Dhaka Stock Exchange from 2010 to 2021, we proxy sentiment with the Trading Volume Ratio and estimate Random Effects, Panel-Corrected Standard Errors, and two-stage least squares to address cross-sectional dependence and endogeneity. We test two hypotheses within this framework: firstly, that higher investor sentiment reduces, on average, contemporaneous cash dividends; and secondly, that this negative effect is attenuated for firms with stronger prior dividend commitments. However, empirically, we only find support for the second one.

Our results contribute to the literature in three ways. Firstly, they extend behavioural finance evidence to a frontier market by documenting that sentiment materially affects payout choices where retail participation and episodic volatility are large. Secondly, they refine the catering versus conservation debate by showing that the average response to sentiment is one of catering. However, this does not apply to firms with no payout reputation. Thirdly, they strengthen empirical practice in this literature by triangulating estimators that correct for panel data imperfections and by applying an instrumental-variable check to bolster causal interpretation.

Our findings are of practical relevance for investors, regulators, and policymakers. Understanding how sentiment shapes corporate decisions can help institutional investors refine their valuation models, particularly in contexts where fundamental analysis is often overshadowed by sentiment-driven trading (Kluzek & Schmidt-Jessa, 2022; Pieloch-Babiarz, 2021). Regulators can benefit by designing disclosure standards that address informational asymmetries and limit the impact of sentiment-driven trading. Finally, for policymakers, the observed evidence of conservatism despite heightened sentiment speaks to the strategic rationality of firms operating in volatile environments; rather than overreacting to market optimism, some firms appear to signal long-term stability, thereby preserving reputational capital and investor trust.

We acknowledge several limitations. Firstly, while the Trading Volume Ratio captures an attention and liquidity dimension, it does not encompass all dimensions of sentiment; complementary proxies such as Google Trends, news-based text indices, or survey measures may reveal additional channels. Secondly, the lagged-Trading Volume Ratio instrument improves causal inferences but cannot eliminate all dynamic endogeneity or contemporaneous common shocks. Thirdly, as our sample is limited to the Dhaka Stock Exchange, multi-country validation is required to establish external generality. Fourthly, richer shareholder- and board-level governance data would better explain heterogeneous firm responses. These caveats frame the opportunities for future research.

In sum, dividend policy in frontier markets reflects a strategic interplay between market mood and managerial commitment to conservatism. Firms do not simply conform to sentiment-driven pressures, they balance short-term conservation with long-term reputational concerns. We invite future studies to test these mechanisms using alternative sentiment proxies, multi-country samples, and natural experiments to further clarify how behavioural forces interact with corporate payout policy.

Appendix

Table A.1. Variance Inflation Factor (VIF) test

Variable	VIF	1/VIF
Unemployment rate	3.11	0.3219
Inflation rate	3.06	0.3264
Trading volume ratio	1.47	0.6797
GDP growth	1.44	0.6953
Leverage	1.35	0.7409
Profitability	1.29	0.7747
Real interest rate	1.16	0.8605
Age	1.06	0.9402
Firm size	1.05	0.9505
Mean VIF	1.65	

Note: The VIF values presented in this table are based on a simple OLS (pooled) regression model for Equation (1).

Source: own calculation.

Table A.2. Breusch–Pagan / Cook–Weisberg test for heteroskedasticity

Model	Chi2	Prob > chi2	Presence of heteroscedasticity
1	169.14	0.0000	yes

Note: The value of Breusch–Pagan / Cook–Weisberg test for heteroskedasticity presented in this table are based on a simple OLS (pooled) regression model for Equation (1).

Source: own calculation.

Table A.3. Wooldridge test for autocorrelation

Model	F value	Prob > F	Presence of autocorrelation
1	0.014	0.9055	no

Note: The value of Wooldridge test for autocorrelation presented in this table are based on a simple OLS (pooled) regression model for Equation (1).

Source: own calculation.

Table A.4. Pesaran test of cross-sectional independence

Model	Pesaran value	Probability value	Presence of cross-sectional dependence
1	94.032	0.0000	yes

Note: The value for the Pesaran test of cross-sectional Independence presented in this table are based on a simple OLS (pooled) regression model for Equation (1).

Source: own calculation.

Table A.5. Hausman test

Variable	Random effects (B)	Fixed effects (b)	(b – B)	Standard error
Trading Volume Ratio	–0.0279	–0.0299	–0.0021	0.0010
Trading Volume Ratio × Cash dividend$_{t-1}$	0.2831	0.3021	0.0189	0.0056
Firm size	0.1419	0.2414	0.0994	.
Profitability	0.0071	0.0069	–0.0002	0.0026
Leverage	0.0002	0.0007	0.0005	.
Age	0.0095	0.7448	0.7352	.
GDP growth	–3.7378	–4.5893	–0.8516	0.4763
Inflation rate	–5.5911	–3.9832	1.6079	.
Unemployment rate	–2.3957	–10.3801	–7.9844	.
Real interest rate	0.2833	0.2889	0.0056	0.1093
Chi2 (9)	5.69			
Prob > chi2	0.7702			

Note: The standard error in the Hausman test for the difference between fixed effects and random effects estimators is calculated as $\sqrt{\mathrm{diag}(V_b - V_B)}$, where V_b is the variance-covariance matrix of the fixed effects estimator and V_B is that of the random effects estimator.

Source: own calculation.

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Algorithmic trading, liquidity and volatility: Evidence from Poland

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Abstract

The aim of this paper is to examine the causality between pairs of measures that describe the intensity of algorithmic trading, market liquidity and volatility for selected blue-chip companies from the Warsaw Stock Exchange, which were permanently included in the WIG20 index from January 1, 2020, to August 31, 2023. In the study, both daily and high-frequency intraday data are used. The research is based on fundamental concepts of information theory, namely entropy and transfer entropy. Additionally, Rényi entropy is used to examine the causal relationships between extreme values of the variables. Our results, based on Shannon's transfer entropy, suggest that algorithmic trading affects liquidity and volatility. The main finding is that if the frequency increases, the number of companies for which information transfer is significant also grows. However, this relationship is not observed for extreme values, for which Rényi entropy is applied.

Keywords

- algorithmic trading intensity
- liquidity
- volatility
- transfer entropy

JEL codes: G10, G19

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Introduction

New directions in financial research stem from the rapid growth of algorithmic trading (AT). This phenomenon has been observed since the 1990s (Mestel et al., 2018). Trading in financial markets has experienced an extremely strong shift towards AT.

The first studies to be conducted in this field, e.g., by Hendershott et al. (2011), indicated that AT improves liquidity and enhances the informativeness of quotes. However, they did not use the exact AT data, but the rate of electronic message traffic as a proxy for the amount of AT taking place. Although Desagre et al. (2022) observed a general improvement in liquidity on the Euronext stock exchange between 2002 and 2006, they noted that those stocks that are most heavily traded algorithmically show the weakest growth in liquidity and lose the liquidity advantage observed before the advent of AT.

In stable conditions, AT increases liquidity by reducing the bid-ask spread and increasing the depth of the order book. However, it may happen that algorithms cancel orders before they are executed, thus creating apparent liquidity (which happens in conditions of high volatility). Regulators in some countries including the United States, the European Union, China, and Japan have been concerned about the possible negative effects of AT on market quality. Therefore, they have taken measures to limit its expansion.

The first exact high-frequency trading (HFT) data on trades and quotes from 26 high-frequency trading firms across 120 stocks was provided by NASDAQ and covered 2008 and 2009. Using the supplied dataset, Brogaard et al. (2014) examined the role of high-frequency traders (HFTs) in price discovery and price efficiency. NASDAQ's 120 stock sample (with explicit HFT flags) from 2008–2009 has become the most popular dataset for training purposes. Other authors have access to proprietary account-level data on trades produced by AT firms. Most of these first exact AT datasets covered only short sample periods, not more than several months (e.g., for Germany only 13 trading days, Hendershott & Riordan, 2013). We note that "algorithmic" and "high-frequency" are not synonyms, but they are closely related notions. All high-frequency data is used in an algorithmic way, but not all algorithmic processing involves high-frequency data. HFT is a specialised subset of algorithmic trading.

The research is based on unique intraday data for the WSE, which is the largest stock exchange in Central and Eastern Europe. To the best of the authors' knowledge, the paper is the first in the literature to empirically establish the relationships between algorithmic trading intensity (*ATI*), volatility, and liquidity for different data frequencies. Moreover, our data (trades and quotes) contains an algorithmic trade indicator, which allows us to estimate *ATI* directly (without using proxies for AT).

The main purpose of this study is to identify the direction of dependence between *ATI*, market liquidity and volatility for daily and intraday data at different time frequencies for selected blue-chips from the Warsaw Stock Exchange (WSE). Identifying causal relationships can help answer questions such as whether *ATI* has an impact on market quality, which is primarily determined by liquidity and risk (measured, e.g., by volatility). These questions provide a basis for constructing dynamic econometric models with regressors that identify a causal relationship.

Our empirical research utilises information-theoretic methods related to entropy and transfer entropy. The study was conducted for pairs of variables formed from *ATI*, volatility, and liquidity. The concept of Rényi entropy is also used to investigate causality between extreme values of variables. Our results, based on Shannon's transfer entropy, suggest that *ATI* affects both liquidity and volatility.

The rest of the paper is organised as follows: Section 1 is devoted to the literature review and hypotheses development. This is followed by Section 2, which includes an outline of research methodology. In Section 3, the dataset and variables are described. In Section 4, the empirical results are presented and discussed. The final section concludes the paper.

1. Literature review and hypotheses development

The literature indicates that AT typically improves liquidity and reduces the volatility of stock returns. However, there are studies, particularly those concerning emerging stock markets, whose results differ from those obtained for developed stock markets. In emerging or developing markets, some researchers have observed a decrease in liquidity and an increase in volatility as AT's share in stock trading grows. Due to the limited length of this study, however, only a few examples of studies representing these groups of results are provided.

Transfer entropy has been previously employed in the economic context by numerous scholars (Abdi & Ranaldo, 2017; Będowska-Sójka & Kliber, 2021; Brauneis & Mestel, 2018; Diaz & Escribano, 2020; Dionisio et al., 2004; Garma, Klass, 1980; He & Shang, 2017; Leone & Kwabi, 2019; Lesmond, 2005), and especially more recent contributions (Ao & Li, 2024; Banerjee & Nawn, 2024; Lacava et al., 2023; Mestel et al., 2024).

Our paper focuses on the relationships between *ATI*, volatility, and liquidity in the emerging stock market. Empirical research on these relationships requires using different measures of these three variables, and the key works on this topic devote considerable attention to these measures and their properties.

Popular measures of volatility are Garman–Klass (1980) volatility, realised volatility and bi-power variation (Aggarwal &Thomas, 2014), while widely used measures of liquidity include proportional effective spread or quoted spread (Mestel et al., 2018) and realised Amihud illiquidity (Lacava et al., 2023). The authors of the latter paper investigated the theoretical and empirical properties of a refinement of the classic daily Amihud measure. They suggested two measures of realised illiquidity: realised Amihud and high-low Amihud.

Liquidity is commonly measured based on different daily proxies versus benchmarks related to high-frequency data (Abdi & Ranaldo, 2017; Lesmond, 2005). In the literature, a rich body of approaches to approximate liquidity and volatility were proposed by Diaz and Escribano (2020). Moreover, Dionisio et al. (2004) used the concept of mutual information to measure both linear and nonlinear interdependence in financial time series.

The seminal paper by Aggarwal and Thomas (2014) provides evidence of the causal impact of AT on the stability of prices and liquidity. According to these authors, policy makers and regulators often exhibit concerns that the higher level of liquidity is transient because AT exits the market rapidly when unexpected news appears. Their main criticism is that AT causes a higher probability of extreme drops and reversals over a very short period of time during the trading day. The results showed that AT lowered the intraday liquidity risk. It is also demonstrated that higher AT leads to a lower incidence of extreme price movements during the trading day. This paper's contribution lies in its moving towards a causal analysis of the impact of AT upon market quality. The analysis used a high-quality dataset with a long time span and numerous securities. In addition to the well-studied measures of liquidity and volatility, the paper also provides evidence about intra-day ash crashes and intra-day liquidity risk.

In some papers, the correlation between several liquidity proxies and benchmarks is investigated. This stream of research was undertaken and significantly extended by Będowska-Sójka and Kliber (2021). The main aim of their paper was to compare the mutual information shared by various liquidity and volatility estimators within each group separately for a sample period from January 2006 to December 2016. The proxies were computed using either daily data from www.stooq.pl or transaction data from the WSE directly. In this way, daily measures of liquidity and volatility were obtained. The authors found that in terms of their information content, volatility measures are much more coherent, while liquidity ones are more dispersed. The Garman–Klass volatility estimator seemed to be the broadest measure of volatility, while Amihud illiquidity and volatility over volume shared the highest amount of mutual information among liquidity proxies.

Jain et al.'s (2021) findings reflect a new perspective on the impact of algorithmic traders on liquidity provision. These authors examined the importance of AT across a sample of stocks listed on the NYSE between 2001 and 2005, demonstrating that the role of algorithmic traders as liquidity providers declines during periods of high information asymmetry.

Ekinci and Ersan (2022) conducted a study on a sample of 30 blue-chip companies listed on Borsa Istanbul between December 2015 and March 2017. They found a negative impact of high-frequency trading on market quality (high liquidity, narrow bid-ask spread, low transaction costs, efficient price discovery i.e. the speed and accuracy with which the market incorporates new information into asset prices), despite its minor role on Borsa Istanbul. The authors emphasised that the provision of liquidity by entities outside FX markets is significantly reduced as the extent of high-frequency trading increases.

Ramos and Perlin (2020) provided the first evidence of AT reducing liquidity in the Brazilian equities market. Their results were contrary to most studies that found a positive relationship between AT and liquidity. However, this was based not on actual AT data, but on two types of AT proxies. In contrast, Dubey et al. (2021) found conclusively, using Indian data, that a rise in AT led to significant improvements in liquidity and reduced market volatility, especially for large-cap stocks.

In their study, Courdent and McClelland (2022) used a set of AT indicators for South Africa: average trade size, odd-lot volume ratio, and trade-to-order volume ratio. Panel regressions were used to determine the relationship between these indicators and two measures of market quality, namely market liquidity and short-term volatility. The study found a strong positive relationship between market liquidity and average trade size, but the opposite relationship for the other two AT indicators. The study points to a strong positive relationship with short-term volatility. In general, AT has a positive impact on market quality, despite the risk of volatility in some markets. Mestel et al. (2018) found that an increase in the market share of AT causes a reduction in quoted and effective spreads while quoted depth and price impacts are unaffected. These findings are consistent with algorithmic traders, on average acting as market makers.

According to Mestel et al. (2024), the relationship between algorithm intensity and volatility can be both positive and negative. Under normal conditions, high-frequency trading reduces volatility. However, during periods of market shocks, algorithms may react synchronously (e.g., by withdrawing from the market), causing price jumps. High frequency, in turn, can lead to more frequent "mini-flash crashes". In their empirical paper, the authors studied 144 mini flash crashes on the Austrian stock market between 2011 and 2015. Using panel logit models, they tried to relate mini flash crashes to

AT for the Vienna Stock Exchange. The authors addressed endogeneity by using a control function approach, but found no evidence that AT significantly affects mini flash crashes.

Banerjee and Nawn (2024) provide the first direct evidence of the behaviour of proprietary algorithmic traders during the COVID-19 pandemic. In turn, research by Ao and Li (2024) shows that the performance of trading algorithms that refer to directional changes in stock markets may not be as expected.

Arumugam et al. (2023) analysed the impact of trading by algorithmic (AT) and non-algorithmic (NAT) traders on volatility. They also investigated the possible inverse relationship, i.e. the impact of volatility shocks on AT and NAT. ATs are classified as high-frequency traders (HFTs) and buy-side algorithmic traders (BATs). Using spike-resistant volatility estimates, the authors concluded that abnormal directional and non-directional trading by BATs and HFTs increases volatility, while trading by NATs slightly decreases it. One hour after a volatility shock, all traders increase their non-directional trading. The authors reported that BATs engage in more directional trading during volatility shock, while HFTs retreat from such activity.

The ambiguous results presented in the literature demonstrate the difficulty in analysing the relationships between measures of algorithmic / high-frequency trading and liquidity or volatility, indicating that these results may depend not only on the stock markets or data periods but also the explanatory variables used in the analysis.

Based on the literature review, especially approaches and results presented in the papers by Mestel et al. (2018), Aggarwal and Thomas (2014), Będowska-Sójka and Kliber (2021) and Dionisio et al. (2004), we formulate:

H1: There exists a significant pairwise causal dependence between *ATI*, market liquidity and volatility.

The second hypothesis refers to some extent to the Epps effect (Gurgul & Machno, 2017). The Epps effect describes how the correlation between the returns of two different stocks decreases with the length of the interval for which the price changes are measured. By analogy, we suppose that with higher data frequency, the number of companies observed with causal relationships between the economic variables under consideration may increase, i.e. that holds true the conjecture

H2: With the increase in data frequency, the causality patterns are depicted in more companies.

In the case of stock markets, researchers are particularly interested in the relationships, or lack thereof, for extreme values of financial variables, i.e. in the tails of probability distributions (crisis phase or boom phase on the stock markets). These relationships are often more pronounced in this area than in the rest of the distribution domain (Gurgul & Syrek, 2023). This strengthened dependence in the tails of the distributions explains why the occurrence of a crisis in one country implies its rapid spread to other countries (contagion effect). Rényi entropy is used to examine causality in the case of extreme values. We presume that for extreme values, similar relationships exist as those discussed in hypotheses 1 and 2. Therefore, we formulate the following hypothesis:

H3: Extreme values of the variables under consideration exhibit causal relationships more often than for the entire distribution.

2. Research methodology

The linear Granger causality test does not appear to be suitable for testing the proposed research hypotheses, the reason being that the dependencies of economic processes are usually non-linear. Different versions of the linear Granger causality test expand the application possibilities. The Hiemstra–Jones test with Panchenko correction can be used in the case of non-linear causality. The Toda–Yamamoto test allows the examination of causality between variables with different degrees of integration. The description of these tests, conditions of their applicability with reference to source papers can be found in Gurgul et al. (2012) and Gurgul and Lach (2012), e.g. However, time series data, especially those with high-frequency data, still do not simultaneously meet all the conditions for the applicability of individual tests. Given the limitations of the article's length, the authors decided to use only entropy-based methods.

The connections between causality and transfer entropy are discussed in papers, e.g., by Hlaváčková-Schindler et al. (2007) and Syczewska and Struzik (2015).

To test the research hypotheses, we use the transfer entropy methodology based on the properties of stationary Markov processes and information theory, which measures the directional transfer of information between two variables. It shows the extent to which past values of one variable influence the future values of the other variable. Transfer entropy expresses how much information one variable transfers to another. This statistic is used in various fields, primarily in biology, engineering, and finance (to analyse the influence and predict market behaviour).

Most economists use the Diebold–Yilmaz (dynamic connectedness, based on VAR model) methodology in their analysis (Diebold & Yilmaz, 2012, 2014). A prerequisite is the correct estimation of the VAR model (selection of lags, stationarity of variables), which can prove difficult in practice. With a larger number of variables, the VAR becomes very complex, and parameter estimation can be unstable ("curse of dimensionality"). The Diebold–Yilmaz procedure helps us to understand how different markets are connected and how changes in one market can affect others, which is useful in risk management and investment strategies. This procedure evaluates connections, not directional interactions. However, since we do not focus on simultaneous relationships between different markets but on causal relationships with specific directions within the same market, the more appropriate methodology in our study is transfer entropy.

Suppose that X and Y are stationary Markov processes of order k and l, respectively:

$$P_X(X_{t+1} \mid X_t, \ldots, X_{t-k+1}) = P_X(X_{t+1} \mid X_t, \ldots, X_{t-k}) \tag{1}$$

and

$$P_Y(Y_{t+1} \mid Y_t, \ldots, Y_{t-l+1}) = P_Y(Y_{t+1} \mid Y_t, \ldots, Y_{t-l}) \tag{2}$$

The average number of bits needed to encode the observation of X in time $t+1$, once the previous k values are known (Behrendt et al., 2019; Będowska-Sójka & Kliber, 2021) and the average number of bits needed to encode the observation of Y once the previous l values are known, are given by:

$$h_X(k) = -\sum_{x_{t+1}, x_t^{(k)}} P\left(X_{t+1}, X_t^{(k)}\right) \log_2 P\left(X_{t+1} \mid X_t^{(k)}\right) \tag{3}$$

$$h_Y(l) = -\sum_{y_{t+1}, y_t^{(l)}} P\left(Y_{t+1}, Y_t^{(l)}\right) \log_2 P\left(Y_{t+1} \mid Y_t^{(l)}\right) \tag{4}$$

where $X_t^{(k)} = (X_t, \ldots, X_{t-k+1})$ and $Y_t^{(l)} = (Y_t, \ldots, Y_{t-l+1})$.

In the bivariate case, information flows from process to process is measured by quantifying the deviation from the generalised Markov property:

$$P\left(X_{t+1} \mid X_t^{(k)}\right) = P\left(X_{t+1} \mid X_t^{(k)}, Y_t^{(l)}\right) \tag{5}$$

relying on the Kullback–Leibler distance. The formula for Shannon transfer entropy is given by:

$$T_{Y \to X} = \sum_{x_{t+1}, x_t^{(k)}, y_t^{(l)}} P\left(X_{t+1}, X_t^{(k)}, Y_t^{(l)}\right) \log_2 \frac{P\left(X_{t+1} \mid X_t^{(k)}, Y_t^{(l)}\right)}{P\left(X_{t+1} \mid X_t^{(k)}\right)} \tag{6}$$

and measures the information flow from Y to X (whereas similarly defined $T_{X \to Y}$ measures the information flow from Y to X). The difference between these measures gives information about the dominant direction (net information flow). Transfer entropy can also be based on Rényi entropy to model the dependencies between events with low probability, i.e. in the tails of distributions. For the discrete random variable $X = \{x_i\}_{i=1}^{n}$ Rényi entropy is defined as:

$$H_X^q(X) = \frac{1}{1-q} \log_2 \sum_{i=1}^{n} [p_i]^q \tag{7}$$

where q is a positive weighting parameter and $p_i = P(X = x_i)$.

Rényi entropy converges to Shannon entropy as $q \to 1$. If $0 < q < 1$, then events that have a low probability of occurring receive more weight, while for $q > 1$, the weights induce a preference for outcomes X with a higher initial probability.

3. Data and variable description

The data, which was made available by the Warsaw Stock Exchange, was subject to a non-disclose agreement prohibiting its being sharing with third parties. It only indicates whether a given transaction was the result of AT. It does not include information on the known types of AT strategies, such as arbitrage, mean reversion, market timing, trend following, momentum, high frequency trading, index found rebalancing and breakout trading.

Moreover, in the case of (a few) other stock exchanges that register exact AT data, e.g., the Vienna Stock Exchange, AT types are not provided. Due to data availability, the dataset contains the trades and quotes of 12 companies (PZU, KGH, CDR, SPL, JSW, OPL, PKO, PKN, PEO, CPS, PGE, DNP) which were permanently included in WIG20 index in the period from 1 January 2020 to 31 August 2023. We use data that are time-stamped to the microsecond and contain an algorithmic trade indicator which specifies whether or not transactions were executed as a result of AT (defined in Article 4(1)(39) of Directive 2014/65/EU)). We consider trades and quotes only for a continuous trading phase and compute commonly used measures of liquidity and volatility along with the *ATI*. Most existing papers on Polish and foreign stock markets are based on these proxies.

Following Aggarwal and Thomas (2014), we used *ATI* over a fixed interval (day, 1-, 5- and 10-minutes frequencies) for each company, which is expressed as:

$$ATI_t = \frac{TTV_t^{AT}}{TTV_t} \tag{8}$$

where TTV_t^{AT} is the sum of the values of algorithmic trades and TTV_t is the sum of the values of all trades.

For all frequencies under consideration, we calculated Garman–Klass volatility (Garman & Klass, 1980) as:

$$GK_t = \sqrt{0.5\left[\log\left(\frac{H_t}{L_t}\right)\right]^2 - \left(2\log(2)-1\right)\cdot\left[\log\left(\frac{C_t}{O_t}\right)\right]^2} \quad (9)$$

where O_t, H_t, C_t, L_t are the open, highest, close and lowest prices in a given period (this measure is computed for either a one-day period or intradaily).

We also incorporated two popular measures of daily volatility: square root of realised variance and bi-power variation given by the following formulas:

$$RV_t = \sqrt{\sum_{i=1}^{m} r_{t,i}^2} \quad (10)$$

$$BPV_t = \sqrt{\sum_{i=2}^{m} \left|r_{t,i-1}\right|\cdot\left|r_{t,i}\right|} \quad (11)$$

where $r_{t,i}$ are five-minute returns within the day. Parameter is equal to 94, since continuous trading on the WSE is between 9:00 a.m. and 4:50 p.m.

For each frequency we computed two popular measures of liquidity, namely, proportional effective spread and proportional quoted spread of the form:

$$EFF_t = \frac{2\cdot D_t\cdot(P_t - MID_t)}{MID_t} \quad (12)$$

$$QUO_t = \frac{ASK_t - BID_t}{MID_t} \quad (13)$$

where D_t is equal to 1 (–1) if trade at time t was buy (sell) $MID_t = \frac{BID_t + ASK_t}{2}$ is midpoint at time t and P_t is trade price. For each frequency, we computed the size-weighted effective spread and time-weighted quoted spread (Mestel et al., 2018).

In addition, for daily frequency, we computed the recently introduced measure of illiquidity (Lacava et al., 2023), that is, realised Amihud illiquidity:

$$RAI_t = \frac{RPV_t}{V_t} \quad (14)$$

where $RPV_t = \sum_{i=1}^{m}\left|r_i\right|$ is realised absolute variation based on log-returns and $V_t = \sum_{i=1}^{m} v_i$ is trading volume generated in the same period. We used five-minute returns, which gave $m = 94$.

4. Empirical results

4.1. Descriptive statistics

For all intraday time series of liquidity, volatility and *ATI* we reject the null of lack of autocorrelation (Ljung–Box test) and observe a departure from normality (Jarque–Bera test). For two companies' (CDR and PEO) daily series of *ATI*, we do not reject the null in the Ljung–Box test at the 10% significance level.

We analysed all series and observed that the autocorrelation function decays more slowly than exponential decay. Moreover, for intraday data (especially for liquidity and volatility measures) we observe intraday seasonality. In Figure 1, we present as an example the autocorrelation function of proportional quoted spread, proportional effective spread, volatility and *ATI* for the company PGE for five-minute intervals. Intraday seasonality can be easily noticed (investors are characterised by different activity at different moments of the session, especially greater at the beginning and end of the trading day).

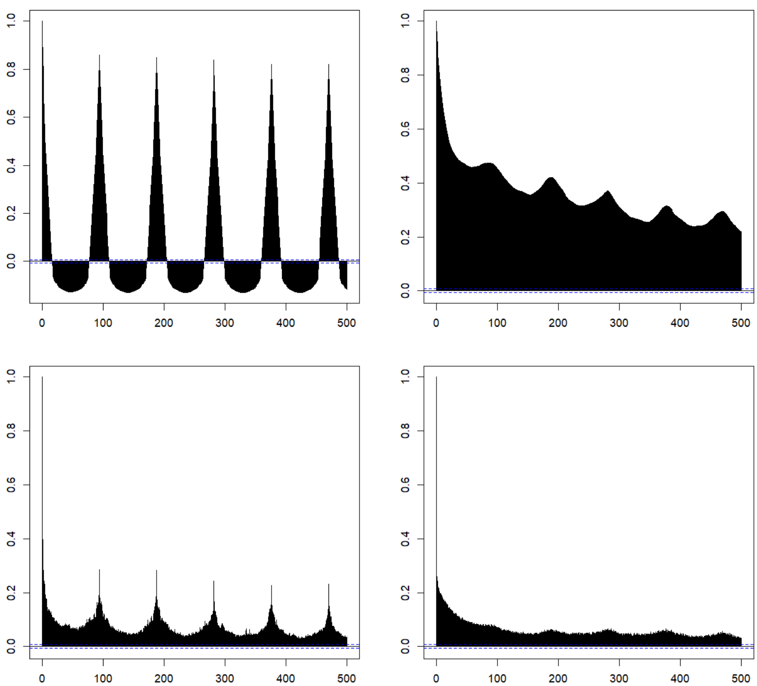

Figure 1. Autocorrelation function of proportional quoted spread (top left), proportional effective spread (top right), volatility (bottom left) and ATI (bottom right) for the company PGE (5-minute intervals)

Source: own calculations.

Table 1. Summary statistics for daily data

Variable	Mean	Standard deviation	Min	Max
ATI	0.49793	0.11687	0.15444	0.84805
EFF	0.00402	0.02227	0.00028	0.27606
QUO	0.00187	0.00082	0.00074	0.00737
RAI	0.00053	0.00060	0.00007	0.00876
GK	0.00051	0.00081	0.00003	0.01395
RV	0.00051	0.00081	0.00003	0.01395
BPV	0.00053	0.00060	0.00007	0.00876

Note: The table contains descriptive statistics calculated on the basis of trades and quotes data of 12 companies that in the period from 1 January 2020 to 31 August 2023 were permanently included in the WIG20 index.

Source: own calculations.

Table 2. Summary statistics for intraday data

Variable	Mean	Standard deviation	Min	Max
1-minute frequency				
ATI	0.452358000	0.367843900	0.000033229	1.000000000
EFF	0.003359015	0.021707740	0.000000001	0.369432100
QUO	0.001522686	0.001440292	0.000200780	0.031584196
GK	0.000000653	0.000004921	0.000000005	0.001370589
5-minute frequency				
ATI	0.539948072	0.228760851	0.000165937	1.000000000
EFF	0.003893397	0.022234950	0.000000001	0.342044900
QUO	0.001866829	0.001156481	0.000209468	0.018047837
GK	0.000004710	0.000017253	0.000000005	0.002113177
10-minute frequency				
ATI	0.531513442	0.194909261	0.000813558	1.000000000
EFF	0.003899087	0.022264020	0.000006655	0.335197400
QUO	0.001871816	0.001067395	0.000229696	0.011459379
GK	0.000010400	0.000035071	0.000000006	0.003143300

Note: The table contains descriptive statistics calculated on the basis of trades and quotes data of 12 companies that in the period from 1 January 2020 to 31 August 2023 were permanently included in the WIG20 index.

Source: own calculations.

The dataset covers the period from 1 January 2020 to 31 August 2023 (910 trading days). The sample contains trades and quotes data of 12 companies that in the mentioned period were permanently included in WIG20 index. In Table 1, we present descriptive statistics of variables for daily data (median values of means, standard deviations, minimums and maximums), whereas in Table 2 we provide the statistics for high frequency data. Daily RV is calculated by summing the squared intraday 5-minute returns over a single trading day. Also, BPV and RAI are daily data constructed from intraday data in the same way.

We remove intraday seasonality in a way described by Bińkowski and Lehalle (2018). Denoting the $x(d, \tau)$ value of variable x on day d at bin τ, we compute:

$$y(d,\tau) = \log x(d,\tau) - \log \overline{x}(d,\tau) \tag{15}$$

where $\log \overline{x}(d,\tau)$ denotes the mean value of $\log x(d,\tau)$ over all days.

Next, we set $m = n^{0.7}$ (where is the length of the series) and apply an exact local Whittle estimator of long memory (Shimotsu & Phillips, 2005). The results are mixed, but in general most of the series are stationary, with a long memory parameter less than 0.5.[3] We differentiate all series using the estimated parameters of the Whittle estimator.

4.2. Transfer entropy

In this section, we establish whether there is any information transfer between *ATI*, liquidity and volatility. We use both Shannon transfer entropy and the Rényi approach. The number of lags used in both approaches is equal to 1 for both variables in a pair. We choose such values to compare our results with Będowska-Sójka & Kliber (2021). The RTransferEntropy package was used for the calculations (Behrendt et al., 2019).

It is recommended to assign q a low value in order to give more weight to extreme events, as in the case of financial time series, the most important information comes in the tails. Following (Będowska-Sójka & Kliber, 2021), we set $q = 0.1$ (default value in the function from the RTransferEntropy package) and this highlights the information in the tails. In Table 3, we present the number of companies of significant information transfer from (to) liquidity and volatility and *ATI* at a significance level of 0.05. The results show that for daily data there is no information transfer from *ATI* to either liquidity or volatility in the case of Shannon entropy. The same holds true for the opposite direction for all variables. When considering the tails of the Rényi approach, the results are quite similar. Table 3 shows the number of companies with significant information transfer from *ATI* to liquidity and in the opposite direction, and from *ATI* to volatility and in the opposite direction.

[3] More detailed results of estimation are available upon request.

Table 3. Number of companies of significant information transfer between *ATI* and liquidity, and between *ATI* and volatility for daily data

***ATI* → liquidity**			
Variable	*QUO*	*EFF*	*RAI*
Shannon	3	1	1
Rényi	2	0	1
liquidity → *ATI*			
Variable	*QUO*	*EFF*	*RAI*
Shannon	1	0	0
Rényi	3	1	0
***ATI* → volatility**			
Variable	*QUO*	*EFF*	*RAI*
Shannon	0	0	1
Rényi	0	1	1
volatility → *ATI*			
Variable	*QUO*	*EFF*	*RAI*
Shannon	0	2	0
Rényi	0	2	1

Note: Significance level is equal to 0.05 using either the Shannon or Rényi approach. The number of lags used in both approaches is equal to 1 for both variables in a pair. In the case of Rényi transfer entropy, we set $q = 0.1$. *RV*, *BPV* and *RAI* are based on 5-minute log returns.

Source: own calculations.

We apply the same procedure to investigate the information transfer between volatility and liquidity. The results are presented in Table 4. In most companies, information transfer exists between volatility and liquidity, especially when liquidity is proxied by *EFF*. When considering Rényi transfer entropy, we do not find information transfer in both directions (with an exception for the direction from volatility to liquidity for the measure of illiquidity). Note that *RV* offers a good approximation of unobservable volatility, but it is sensitive to price spikes, while *BPV* is robust to them. The differences in the results for both estimators are small (maximum two companies). This may indicate a lack of essential jumps.

We compute estimators of and for intraday data (1-, 5-, 10-minute frequencies) to check whether the choice of frequency has an impact on the results obtained. Table 5 contains the number of companies of significant information transfer for high frequency data.

From Shannon transfer entropy, we observe that as the frequency increases, the number of companies of significant information transfer also increases

Table 4. Number of companies with significant information transfer from volatility to liquidity, and from liquidity to volatility for daily data

Volatility → liquidity						
Entropy	Shannon			Rényi		
Variable	*QUO*	*EFF*	*RAI*	*QUO*	*EFF*	*RAI*
GK	4	6	2	0	0	6
RV	6	9	1	0	0	2
BPV	8	9	0	0	0	4
Liquidity → volatility						
Entropy	Shannon			Rényi		
Variable	*QUO*	*EFF*	*RAI*	*QUO*	*EFF*	*RAI*
GK	4	11	1	0	0	0
RV	4	11	0	0	0	1
BPV	6	10	1	0	0	2

Note: Significance level is equal to 0.05 using either the Shannon or Rényi approach. The number of lags used in both approaches is equal to 1 for both variables in a pair. In the case of Rényi transfer entropy, we set $q = 0.1$. *RV*, *BPV* and *RAI* are based on 5-minute log returns.

Source: own calculations.

Table 5. Number of companies of significant information transfer for intraday data

Pair	*ATI- GK*		*ATI- QUO*		*ATI- EFF*		*GK- QUO*		*GK-EFF*	
Direction	→	←	→	←	→	←	→	←	→	←
1-minute frequency										
Shannon	12	12	12	12	12	12	12	12	12	12
Rényi	11	3	4	0	7	2	3	12	12	12
5-minute frequency										
Shannon	12	12	12	12	12	12	12	12	12	12
Rényi	1	3	0	1	0	1	2	2	3	1
10-minute frequency										
Shannon	11	11	8	8	12	12	12	12	9	7
Rényi	0	0	0	0	0	0	2	1	1	2

Note: Significance level is equal to 0.05 using either the Shannon or Rényi approach. The number of lags used in both approaches is equal to 1 for both variables in a pair. In the case of Rényi transfer entropy, we set $q = 0.1$.

Source: own calculations.

for all variables under consideration. For 1-minute and 5-minute frequency, we observe directional dependence for all companies. When considering the tails (Rényi approach), there is a huge difference for 1- and 5-minute (10-minute) frequencies, and we do not observe as many companies of significant information transfer, although their number increases as frequency increases (Rényi approach compared to Shannon approach). In Table 6, we present net information flow, that is, the difference $T_{X \to Y} - T_{Y \to X}$.

Table 6. Net information flow for 1- and 5-minute data (Shannon entropy)

Variable X	Variable Y	1-minute frequency	5-minute frequency
ATI	*GK*	0.0006	0.0014
ATI	*QUO*	0.0003	0.0005
ATI	*EFF*	–0.0062	–0.0007
GK	*QUO*	–0.0008	–0.0008
GK	*EFF*	–0.0011	–0.0008

Note: This table shows the net information flow for 1- and 5-minute frequencies for which significant information transfer was detected for all companies (Shannon entropy).

Source: own calculations.

From Table 6, we conclude that more information flows from *ATI* to volatility than vice versa. The situation is similar when we consider quoted spread instead of volatility. However, for the effective spread (when we consider variables X = *ATI* and Y = *EFF*), the sign of net information flow is negative—the impact of *EFF* on *ATI* is greater than *ATI* on *EFF*. The situation is similar when we consider volatility as X and liquidity as Y. The conclusions are the same regardless of whether we consider 1- or 5-minute data.

Conclusions

We analysed causal dependencies between financial time series of *ATI*, liquidity and volatility. The paper employs a dataset for 12 companies listed on WSE that identifies transactions made through algorithmic trading in the period from 1 January 2020 to 31 August 2023. One limitation of the study is the sample size, limited to the most liquid companies listed on the Warsaw Stock Exchange. For daily data, we observe a lack of information transfer from *ATI* to liquidity and volatility regardless of whether we use Shannon or Rényi entropy. Regarding pair volatility-liquidity, we find information transfer from liquidity to volatility for almost all the companies considered, but only for proportional effective spread. This does not hold for tail dependencies when applying Rényi transfer entropy. Our result is in line with Będowska-Sójka and Kliber (2021), who used similar volatility estimators (realised variance and bi-power variation) and method (transfer entropy), but different liquidity measures. The results are similar regardless of whether we use a robust measure for price jumps (bi-power variation) or not (realised variance).

For intraday data, the results are quite different. Considering the entire distributions, we observe that as the data frequency increases, the number of companies with significant information transfer also increases (for all pairs of variables). For 1- and 5-minute frequencies, significant information transfer is observed for all companies under consideration.

Causal relationships of *ATI*, liquidity and volatility for high frequency data are identified. HFT has an impact on volatility and liquidity, and there is also a feedback loop. Given that the one lag is used in the transfer information approach, it shows the fast reaction of one variable to another. Trading bots that enter into transactions react to market signals in a short time, thus causing changes in market volatility and liquidity. The empirical results offer support for H1 and H2, according to which a significant pairwise causal dependence exists between *ATI*, market liquidity and volatility, and an increase in data frequency leads to more numerous causality patterns. However, these results are not comparable with previous studies because we use high-frequency data and investigate causality using changes in entropy.

Based on the results of net information flow, we conclude that more information flows from *ATI* to volatility rather than vice versa. The situation is similar when we consider quoted spread instead of volatility. However, the impact of effective spread on *ATI* is greater than vice versa. When considering volatility and liquidity, we observe a greater flow of information from liquidity to volatility. The conclusions are the same regardless of whether we consider 1- or 5-minute data.

When considering the tails (describing events that have low probability), there is a clear difference in the number of companies of significant information transfer for 1-minute and 5-minute (10-minute) frequencies, and we observe causal dependence only between volatility and the effective spread in both directions for highest frequency. The results based on Rényi entropy do not support H3, which posited more frequent causal relationships among variables in the tails of the distribution.

These interesting findings require further research into their causes and economic interpretation. By varying the number of lags in the computation of transfer entropy, we could determine the duration of information's influence on the market. Other types of entropy could also be used, e.g., Tsallis entropy, which is a versatile tool for measuring the degree of uncertainty in complex or non-linear systems.

From both the theoretical and practical point of view, it would be interesting to take into account the intra-day seasonal pattern when examining the relationships between variables. We intend to expand our sample to include mid-sized companies. The goal would be to examine how company size affects performance. The plan also includes investigating the impact of changes in the value of the Rényi entropy weighting parameter on the results. An interesting topic for future research concerns the shapes of the probability distributions of intraday values of the studied variables, their similarity and changes over time.

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