

Ownership Structure and Firm Performance. A Panel Regression Study from Poland

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Abstract

This paper explores the relationship between ownership structure and firm performance, using the framework of agency theory. Given a sample of 126 Polish non-financial firms listed on the Warsaw Stock Exchange between 2016-2021, we applied fixed and random effects panel regressions with robust standard errors, tested for specification, endogeneity, cross-sectional heteroskedasticity, and serial autocorrelation. Our models identify several significant associations of adopted ownership structures with firm performance measured by accounting and market-based measures. We find that ownership concentration by the largest shareholder is negatively related to enterprise market-measured performance. Additionally, our results indicate the significance of shareholder heterogeneity.

Keywords: ownership structure, ownership concentration, dual-class shares, shareholder heterogeneity, panel regression

JEL Classification: C23, C58, G30, G32

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1 Introduction

The link between corporate ownership and firm performance has been at the centre of financial economics literature for years. Initially, studies on the implications of corporate ownership structure on firm performance have been dominated by the concept of shareholder-manager conflict. As early as in the 18th Century, Smith (1776) noted that managers of joint-stock companies cannot be expected to exercise the same level of vigilance in creating firm value as sole owners. Two centuries later, this divergence of interest was put forward into a conceptual framework of agency theory (Jensen and Meckling, 1976). Despite the fact that the effects of ownership structure on firm performance have been a subject of extensive literature for decades, there is no academic consensus on the direction of these effects.

In the mid-80s, a proliferation of dual-class shares (DCS) listings associated with the allowance of DCS at the American Stock Exchange (AMEX) and the introduction of the NASDAQ stock exchange brought greater awareness to another spectrum of corporate ownership implications. Despite being a fundamental topic in the corporate governance literature, the effects of divergence between cash flow and control rights on firm value remain unclear. Some scholars note that these alternative ownership structures amplify an agency conflict since insiders controlling disproportionately more voting rights than ownership rights do not bear the financial consequences of their decisions to a proportionate extent (Masulis et al. 2009, Gompers et al. 2010). On the contrary, other researchers argued that a dual-class shares structure allows decision-makers to execute their long-term strategy, reducing the adverse impact of short-term market pressures (Bebchuk and Kastiel 2017, Shleifer and Vishny 1997, Jordan et al. 2016).

Nonetheless, most studies on the impact of corporate ownership and control on companies performance are based on the US equity market, where common ownership is most evident and external corporate governance mechanisms are strong (Hilli et al. 2013). Over the last three decades, a wave of political events combined with the impact of the globalisation process resulted in the creation of several new equity markets, in which ownership structures differ substantially from those observed among their Anglo-Saxon counterparts. Maury and Pajuste (2002) note that the pivotal agency conflict in the context of emerging markets is the one between controlling shareholders and minority owners. The review of existing studies on corporate ownership in emerging markets shows that ownership structures are vastly dominated by large blockholders, which combined with weak investor protection mechanisms and less efficient legal systems (Pistor 2013) provides an excellent ground for investigating further implications of the agency conflicts. In addition, more recent research on the effects that ownership structure has on firm performance revealed the significance of the blockholder identity, as different shareholders may engage in corporate governance through various channels (Aluchna and Kamiński 2017, Edmans 2014). Starting in Poland, a revolution of 1989 resulted in several economic reforms being

implemented in Central and Eastern European countries. Political freedom was followed by economic liberalisation. The transition from a centrally planned to a market economy was associated with the introduction of privatisation schemes, substantial regulatory changes, and the foundation of several new business entities. This led to the development of new corporate structures, the investigation of which might make a significant contribution to the literature on the implications of agency conflicts. In 1991, after 52 years, the Warsaw Stock Exchange reopened. As of the last trading session of 2021, it hosts 430 companies, which account for approximately 50% of annualised GDP (Warsaw Stock Exchange 2021). In 2004, Poland joins the European Union, marking the end of the transition period and opening up new possibilities for Polish enterprises (Kolodziejczyk 2016). Finally, in 2018, FTSE Russel, a global index provider, updated Poland's status from an emerging market to a developed market (Warsaw Stock Exchange 2021). This alongside still relatively weak external corporate governance mechanisms and a large concentration of ownership by heterogeneous shareholders provide a unique landscape to investigate the implications of adopted ownership structures in the context of the economy transitioning to a status of a developed market.

In this study, we investigate the links between ownership structures and firm performance in the context of publicly listed Polish companies. We utilised a sample of 126 Polish non-financial firms listed on the Warsaw Stock Exchange between 2016-2021, which combines data from Thomson Reuters Eikon and manually extracted from annual activity reports. We use two alternative measures of firm performance. The first measure is a return on assets, a common ratio employed in several studies. The second measure of firm performance is the approximation of Tobin's Q, which allows capturing market perception of firm performance. To model those metrics we applied a modelling pipeline inspired by Aluchna and Kamiński (2017) which consists of both fixed and random effects models, making a choice based on the results of Hausman specification tests. The plethora of hypotheses investigated in this paper enables a comprehensive exploration of various facets of ownership structure and their implications for firm performance. Despite the complexity of the analysis, the structured approach facilitates direct comparisons with prior research, particularly the study conducted by Aluchna and Kamiński (2017). In addition to testing the persistence of effects identified by Aluchna and Kamiński (2017), this study introduces market-based performance metrics, as expropriation of minority owners may occur irrespective of firms' accounting performance. Furthermore, recognising the substantial effort involved in manually collecting data, this study is, to the best knowledge of the authors, the first to investigate the relationship between firm performance and dual-class share structure among Polish public firms. We tested for endogeneity and variance-covariance matrix issues which led us to use robustified standard error adjusted for both cross-sectional-heteroskedasticity and serial correlation. It all enabled reliable inference about the hypotheses initially made. The main thesis of this article is that in the context of post-transitional economies,

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ownership structure is linked with the performance of public companies and that factors such as ownership concentration and controlling shareholder identity should be taken into consideration in investment decision-making process.

This paper is structured as follows. The paper will begin by stating our research hypotheses and providing a literature review of existing studies on links between ownership structures and firm performance. The second section explains the sample selection process and presents our data set. Subsequently, we describe the employed methodology and present the empirical results of our panel data regression models. Next, we provide a discussion of our findings. The final section summarises the results of our study and highlights key contributions to the literature, practical implications, and possible research gaps.

2 Literature and hypotheses

Ownership concentration, which refers to the level of control by a single or a few large blockholders, has received significant attention in the academic literature as a factor influencing firm performance. Nevertheless, the review of existing literature gives no universal conclusion on the effects of ownership concentration. Many scholars derived their assumptions from the theory of ownership structure of the firm, which presumes the existence of an optimal ownership structure, that minimises the adverse impact of agency costs (Jensen and Meckling 1976). In accordance with that theory, since agency costs arise from conflicts of interest between owners (principals) and managers (agents), the only zero agency-cost company will be one fully owned by management. Consequently, the agency costs will grow alongside ownership diffusion and will be inversely related to management ownership (Ang et al. 2000). Shleifer and Vishny (1994) note, that in firms with highly diffused ownership, shareholders possessing a relatively small portion of cash flow rights might lack the incentive to bear monitoring costs. Thus, the presence of a blockholder such as family or institutional investors might be beneficial for alleviating agency costs. On the contrary, in the analysis of 200 large US corporations Demsetz and Lehn (1985) report a negative relationship between ownership concentration and firm value and argue that large shareholders may use their power to extract private benefits rather than maximise the firm's value (Zwiebel 1996, Lambrecht and Myers 2008). The existing literature also provides some evidence of the positive impact that entrepreneurial risk-seeking behaviours have on the future growth of companies (Baumol et al. 2007, John et al. 2008). Hence, the substantial concentration of ownership in the hands of one non-diversified shareholder might reduce corporate risk-taking abilities and therefore, disincentives undertaking riskier investments, which are one of the pivotal drivers of growth (Faccio et al. 2011). Whilst the image of widely held firms controlled by opportunistically acting managers, coined by Berle and Means (1932), might be associated with Anglo-Saxon public companies, the ownership concentration of public companies incorporated in other wealthy economies alongside emerging markets is significantly higher. Moreover,

the controlling shareholders often possess control rights that substantially exceed their cash-flow rights (La Porta et al. 1999). Consequently, due to differences in both internal and external corporate governance mechanisms, agency conflicts systematically vary across different firms and countries (Morellec et al. 2018).

In countries where the legal environment is weak and ownership is highly concentrated, minority investors may be at greater risk of abuse (Wang and Shailer 2015). Since weaker external governance mechanisms exacerbate expropriation risks (La Porta et al. 1999), the emerging markets environment highlights the significance of the second type of agency costs. The existing studies on the effects of ownership concentration in emerging equity markets, divided between the monitoring and expropriation hypotheses, provide mixed results. To address the diversion of empirical findings, using 419 correlations from 42 studies on ownership concentration in 18 emerging markets, Wang and Shailer (2015) introduced a meta-analytical technique. After adjusting for population differences and researchers' modelling choices, they find a negative relation between ownership concentration and firms' financial performance. In a study of 513 non-financial companies listed on the Warsaw Stock Exchange between 2005 and 2014, Aluchna et al. (2019) report that the ownership concentration by the two largest shareholders is negatively related to the dividend payout ratio, providing further support for the expropriation hypothesis.

Using a sample of 12,652 European and US firms, Morellec et al. (2018) investigate agency costs across 14 OECD countries. In order to quantify second-type agency costs, authors constructed novel firm-level indexes for agency conflicts and found, that costs of principal-principal conflicts are substantial across firms and countries. Furthermore, private control benefits, measured as a share of free cash flows, are reported to be highest in Poland (5,2%), as compared to a sample median of 2,6%. Finally, in a panel study of 495 Polish non-financial firms listed on the Warsaw Stock Exchange, Aluchna and Kamiński (2017) report a negative and significant relationship between the largest shareholder's ownership concentration and the firm's financial performance. Since the link between ownership concentration and firm performance is found to be strongly dependent on country-specific factors, it is expected that in a market with relatively weak external corporate governance mechanisms (Pistor 2013), the extraction of private benefits of control will prevail in the monitoring-expropriation trade-off. Hence, the following hypotheses are formulated:

H1. Ownership concentration by the largest shareholder is adversely related to firm performance

Agency problems have been at the centre of the corporate governance literature for decades. However, the majority of the existing studies concentrated on firms, in which voting and cash-flow rights are generally parallel (Masulis et al. 2009). The introduction of NASDAQ, implementation of the 9C-4 rule by the Securities Exchange Commission, relaxing of voting policies by NYSE, and consequential, albeit gradual, proliferation of DCS's IPOs heated the debate over the one-share, one-vote rule.

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Studies on the link between the divergence of cash-flow rights and voting rights and firm performance provide ambivalent conclusions. Jordan et al. (2016) find that dual-class shares allow managers to pursue a long-term growth strategy and alleviate the impact of short-term market pressures. Furthermore, the results indicate higher sales growth and R&D intensity among dual-class firms. This reasoning is further supported by Dugar and Nathan (1995), Hong and Kubik (2003), and He and Tian (2013), who contend that financial analysts concentrate on short-term earnings, which discourages managers from making strategic investments with a long-term horizon. Johnson et al. (2015) argue that dual-class shares reduce the likelihood of a takeover, which decreases the risk of changing the firm's operating strategy and imposing costs on its business partners. Hence, the adoption of DCS strengthens its business relationships and is positively related to their longevity. Bebchuk and Kastiel (2017) conduct a comprehensive analysis of the benefits and drawbacks of dual-share structures and note that advantages such as founders' superior leadership skills gradually recede and agency costs increase as time passes from the initial public offering. Chemmanur and Jiao (2012) analyse a firm's choice between dual-class and single-class share structures at its IPO and find that implementing a dual-class structure might substantially increase a firm's long-term value in high-ability managers' hands. However, it is also reported that in the case of low-ability incumbents, the presence of DCS aggravates principal-principal conflict and adversely affects the firm's value. Using a sample of US dual-class companies, Masulis et al. (2009) find that as the divergence between cash-flow rights and voting rights widens, corporate cash holdings are worth less to outside shareholders, management compensation is higher, value-destroying acquisitions are reported more often and corporate investments provide lower returns. This is consistent with the results of several studies supporting the hypothesis that dual-class shares structure adversely affects the firm performance (Gompers et al. 2010, Baulkaran 2014, Wang and Xie 2009, Bebchuk et al. 2000). Amoako-Adu et al. (2014) conduct an analysis of cash distribution in US dual-share companies and report that dual-class firms repurchase fewer shares and have lower dividend payout ratios, providing further support for the expropriation reasoning. These conclusions are also consistent with the findings of Gugler and Yurtoglu (2003), who examine 736 dividend change announcements in Germany and find that deviations from the one-share-one-vote rule result in lower pay-out ratios and highly adverse wealth effects.

La Porta et al. (1999) note that firms listed outside of Anglo-Saxon countries tend to have controlling shareholders, who often possess control rights, substantially exceeding their cash-flow rights. Moreover, the findings of Gopalan and Jayaraman (2012) indicate that the effects of divergence between cash-flow rights and control rights on the extraction of private benefits are most evident in countries with poor investor protection mechanisms. The expropriation hypothesis is further supported by the results of a study on publicly traded corporations in Asian emerging economies by Claessens et al. (2002), who report that firm value falls when the control rights exceed

cash-flow rights. Since the issuance of multiple voting shares is not permitted in the case of firms listed on the Warsaw Stock Exchange, the existing divergence between voting rights and cash-flow rights results from the series issued before the initial public offering. In a panel study of 105 WSE-listed companies using control-enhancing shares, Jewartowski and Kaldowski (2015) report that dual-class firms are associated with higher levels of debt. Preference for debt, as non-diluting security, indicates that financial decisions in dual-class firms may be driven by control motivations. Combined, these factors increase the potential for extreme agency problems. Hence, we formulate the following hypothesis.

H2. The positive difference between voting rights and cash-flow rights is negatively related to firm performance

Shareholder identity. Whilst the implications of the separation of ownership and control have been present in corporate governance literature for some time, the heterogeneity of blockholders received substantial attention more recently. Clifford and Lindsey (2016) argue that different shareholders may engage in corporate governance through various channels. Consequently, assuming the homogeneity of blockholders could lead to overlooking important effects at a more granular level (Edmans 2014). This is consistent with the conclusions of a number of scholars, who note that the effects of ownership concentration are dependent on controlling blockholder identity (Thomsen and Pedersen 2000, J Hadlock and Schwartz-Ziv 2019). Since various shareholders have different motivations and goals, the identity of the controlling owner might determine the direction of skewness in the trade-off between providing monitoring and extracting private benefits of control. Moreover, large non-controlling shareholders have the incentive, power, and resources to bear monitoring costs due to the substantial size of their shareholdings (Maury and Pajuste 2005). Therefore, the presence of large non-controlling blockholders might provide monitoring of the controlling shareholder and reduce agency costs derived from the expropriation of minority shareholders (Laeven and Levine 2008). Alternatively, large non-controlling shareholders may collude with the controlling entity and effectively amplify the agency costs by participating in the expropriation of other minority owners (Cai et al. 2016).

State ownership. Since state participation in the shareholding of public companies is hardly ever visible in terms of public enterprises incorporated in Anglo-Saxon countries, the Western literature on the subject is not particularly extensive. In the past, economists have generally seen state-owned enterprises as a cure to market inefficiencies, particularly in terms of monopolistic services, which may create substantial divergence of interest between private and social goals (Shleifer and Vishny 1994). In general, the review of existing financial economics literature indicates, that state ownership is negatively associated with firm performance (Estrin et al. 2009, Wang and Shailer 2018, Djankov and Murrell 2002). The divergence of goals and motivations between the state and other shareholders goes beyond first and second-

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type agency costs. The state in the role of controlling blockholder might pursue political objectives rather than maximise firm value (Liu and Zhang 2018, Szarzec and Nowara 2017). The subsequent viable factor impacting the performance of public companies partially owned by the state can be explained by the theory of “tragedy of commons”. This theory can apply to SOEs in the sense, that both managers and employees can perceive the company as public property, which in turn might result in the creation of additional incentives to extract private benefits (Estrin and Perotin 1991). In a study of nearly twelve thousand Polish joint-stock companies (Szarzec et al. 2022) report, that changes in management and supervisory boards of state-controlled public enterprises are more frequent than those observed in privately held companies, reaching a peak three months after the formation of a new government. Apart from the adverse influence of high management turnover on firm performance (Kim et al. 2021), the dependency of management on the outcome of elections may amplify managerial entrenchment and disincentivise agents, whose tenure is not determined solely by their professional achievements. On the contrary, several scholars observed some benefits associated with state ownership such as access to government contracts (Goldman et al. 2009), lower cost of debt capital (Le and Tannous 2016), and lower risk of default (Del Bo et al. 2017).

Although there is substantial literature on the negative implications of state ownership, studies on the effects of state ownership on firm performance in Central and Eastern Europe, in general, do not provide strong evidence on the effects of state ownership on firm performance. Using 4425 estimates gathered from 204 existing studies Iwasaki et al. (2022) conduct a comparative meta-analysis of the ownership structure of firms publicly listed in Russia, China, and Eastern-European EU member states and note that state ownership adversely affects firm performance. Moreover, the hypothesis stating the adverse implications of state ownership is tested and accepted for all three distinct regional groups. Aluchna et al. (2019) investigate the impact of ownership concentrations alongside blockholder identity and note that dividend payout is much lower in terms of state-owned enterprises, providing further support for the expropriation hypothesis. In contrast, Hanousek et al. (2007) examine the relationship between the performance and ownership structure of Czech public companies and find a positive correlation between government ownership and firm value. The results of univariate analysis and discriminant analysis performed on a group of Polish non-financial both private and state-controlled public companies by Kabaciński et al. (2020) indicate that SOEs realise a higher return on assets, albeit underperform in terms of liquidity and inventory management. In a study of 500 largest enterprises operating in 13 post-socialist Central and Eastern European countries, Szarzec and Nowara (2017) find that SOEs’ performance is comparable to their private peers. Finally, in a panel study conducted by Aluchna and Kamiński (2017), the authors do not observe any significant links between government ownership and the performance of firms listed on the Warsaw Stock Exchange. Considering the evidence provided by previous studies and following the reasoning that implications

of state ownership on performance are highly dependent on the quality of the institutional environment (Borghesi et al. 2016, Castelnovo et al. 2019), we formulate the following hypothesis.

H3. State ownership is negatively related to firm performance

Corporate Investors. In general, results of studies on the links between industry investors' ownership and firm performance indicate a positive association. Morck et al. (2005) observe that controlling corporations might benefit their portfolio companies through technology transfers, synergistic effects, and infrastructure sharing. Since corporate blockholders possess industry expertise as compared to other shareholders, they may also provide more effective monitoring and positively influence the composition of management (Allen and Phillips 2000). In contrast, Atanasov et al. (2010) examine potential agency costs at publicly listed subsidiaries in the United States and provide evidence of the expropriation of minority shareholders by controlling corporations. The most evident effects are observed when the stake owned by the dominant shareholder is relatively small. In a panel study of non-financial WSE-listed companies, Aluchna et al. (2019) investigate the relationship between dividend payouts and ownership structure and report that stakes held by industry investors constrain the expropriation of minority owners. These results are also consistent with an earlier study conducted by Aluchna and Kamiński (2017), who report a positive relationship between ownership by industry investors and ROA. Considering the efficient monitoring perspective, internalization of synergy effects, and the long-term investment horizon that industry investors take, we state the following hypothesis.

H4. Ownership by industry investors is positively associated with firm performance

Managerial Ownership. The agency theory assumes that principal-agent conflict arises because managers do not face the financial consequences of their decisions to the same extent the owners do (Jensen and Meckling 1976). Under that reasoning, the zero-agency cost structure is one, when the manager is the sole owner of the business (Ang et al. 2000). However, the ownership structures of companies around the world differ substantially (La Porta et al. 1999). Maury and Pajuste (2005) argue that in the context of emerging markets, the pivotal agency conflict is the one between the controlling blockholder and minority shareholders. Even though the effects of conflict between managers (owners) and non-controlling shareholders have been at the center of corporate governance literature for some time, studies investigating the implications of managerial ownership and firm performance remain divided between entrenchment and unification of interests hypotheses (Cheng et al. 2012). The studies based on Anglo-Saxon markets, when minority investor protection mechanisms are stronger, provide generally favorable evidence of managerial ownership. Lilienfeld-Toal and Ruenzi (2014) argue that managerial ownership provides a strong incentive for

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managers to increase firm value and report that firms with high managerial ownership have a higher return on assets by 5 percent. Nonetheless, the studies on managerial ownership in other markets, where minority investor protection mechanisms are less effective and ownership concentration is higher, provide mixed results. In a meta-analysis performed individually for Eastern European EU countries, China and Russia, Iwasaki et al. (2022) find a positive relationship between managerial ownership and firm performance in all three regions. Managerial holdings often constitute a pivotal part of their personal wealth, which in turn may impede corporate risk-taking abilities (Faccio et al. 2011), negatively impacting growth perspectives. For Poland, Aluchna and Kamiński (2017) hypothesise that managerial ownership adversely influences firm performance. However, the results do not indicate any significant links. In view of the results of provided studies, we expect the convergence of interests reasoning to prevail in the context of Polish public firms. Therefore, we propose the following hypothesis:

H5. Managerial ownership is positively associated with firm performance

Institutional investors. The academic debate on the implications of institutional investors' presence in ownership structures has been dominated by two conflicting reasonings: monitoring and short-termism (Callen and Fang 2013). Previous research provides substantial evidence of the positive impact that the presence of institutional investors has on firm performance (Thomsen and Pedersen 2000, Ferreira and Matos 2008, Dimson et al. 2015). Scott (2014) argues that institutional investors' presence is positively related to R&D expenditures, which increase the long-term performance of portfolio companies. Results of studies also indicate that institutional ownership diminishes agency costs (Mizuno 2014), reduces CEO turnover (Helwege et al. 2012), and more often links CEO compensation to long-term firm performance (David et al. 1998, Zhang et al. 2021). In contrast, Bebchuk et al. (2017) contend that institutions are controlled by investment managers, who may act opportunistically and in turn, create an additional layer of agency costs. Following the short-termism reasoning, Davis (2008) notes that institutional owners are often more likely to exit than to exercise their control rights. Institutional investors with short-term investment horizons may also apply pressure on portfolio company management to reduce strategic investments with an aim to meet short-term earnings goals (Bushee 1998). The short-termism of institutional investors as transient owners seems to be more evident in terms of liquid companies in which the liquidity encourages institutions to sell their stake rather than intervene (Black et al. 2014). Studies on links between institutional ownership and firm performance in emerging markets deliver mixed results. Using 3297 observations covering 516 WSE-listed non-financial firms, Aluchna et al. (2019) investigate collusion among blockholders to reduce dividend payouts. Findings reveal that the presence of institutional blockholder provides monitoring the potential for reducing agency costs and potentially reduces the risk of expropriation. Based on a sample of 180 Bangladeshi companies listed from

2008 to 2018, Abedin et al. (2022) investigate the link between institutional ownership and firm performance measured by ROA and Tobin's Q and report that both domestic and foreign institutional ownership positively affects firm performance. These results are partially consistent with Lin and Fu (2017), who employ a simultaneous equations model with a generalised method of moments to explore the relationship between the presence of institutional blockholders and firm performance of Chinese publicly listed firms. The results suggest that large foreign institutional investors' presence has greater positive effects on firm performance measured by Tobin's Q, as compared to small domestic entities. Finally, in order to address the relationship between ownership structure and the financial performance of Polish firms, Aluchna and Kamiński (2017) analyse the data from the unique sample of 495 Polish non-financial companies listed on the Warsaw Stock Exchange from 2005 to 2014. Even though the authors hypothesise that stakes held by financial institutions positively influence portfolio companies' performance, they do not find any significant links between institutional ownership and firm performance measured by both ROA and ROE. In light of discussed results and weak external corporate governance mechanisms in Poland, we follow the monitoring reasoning and expect that:

H6. Institutional ownership is positively related to firm performance

Individual investors. The literature on the effects of large individual investors on firm performance concentrates on family firms. In Poland, due to the relatively young age of public companies, most firms with significant ownership of individual blockholders are "first generation", which means that the founder is often the largest individual shareholder (Kowalewski et al. 2009). Therefore, the review of studies of non-managerial family involvement should give an insightful perspective on the impact of large individual owners on firm performance. Following the reasoning that large shareholders might provide effective monitoring and directly influence managers (La Porta et al. 1999), the presence of large blockholder such as family can be beneficial for alleviating agency costs (Shleifer and Vishny 1994). A number of studies on the link between non-managerial family ownership and firm performance have shown positive effects. In a study examining the effects of family non-managerial ownership, Poutziouris et al. (2015) report that this relationship is positive, albeit not linear. The performance is observed to increase until family shareholdings exceed 31 percent. Apart from providing effective monitoring, non-managerial family involvement may also limit managerial opportunism and myopia (Anderson and Reeb 2003, Block et al. 2011). On the other hand, several scholars noted that non-managerial family involvement leads to agency problems that erode firm performance due to oligarchic control (Morck et al. 2005), nepotism (Schulze et al. 2003), and expropriation of private benefits (D'Angelo et al. 2016). Similarly to managerial ownership, large individual shareholders are mostly not diversified, what might decrease their risk appetite and in turn, negatively affect the firm growth perspectives (Faccio et al. 2011).

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Although non-managerial family ownership has been a subject of extensive literature, the majority of studies concentrate on developed markets characterised by high investor protection. Thus, conclusions from Anglo-Saxon countries may not be transferable to markets with less effective external corporate governance mechanisms (Wang and Shailer 2018). In a panel study of 217 Polish public companies listed on the WSE between 1997 to 2005, Kowalewski et al. (2009) investigate the influence of family ownership on firm performance. The results provide evidence of a U-shaped relationship between family involvement in ownership and firm performance measured by ROA and ROE. Whilst authors report that moderate involvement in ownership positively relates to firm performance, the second conclusion is that family firms with family CEOs outperform their peers with non-family CEO. Following the alignment of interest and long-term investment horizon we expect that:

H7: Ownership by individual investors is associated with higher firm performance

3 Data

This study uses 2016-2021 annual data from companies listed on the Warsaw Stock Exchange. The extensive use of pyramidal ownership structures in Poland requires us to collect ownership data manually. To identify the ultimate owners of the investigated firms and to obtain detailed data on cash flow and voting rights possessed by the three largest shareholders, we made an effort to go through 756 annual consolidated financial statements. The financial data for this panel study are obtained from Eikon Thomson Reuters.

Table 1: Descriptive statistics

variable	mean	q1	median	q3
roa	0.058	0.024	0.050	0.084
q_tobin_approx	1.367	0.507	0.797	1.205
assets (mln PLN)	4519.120	438.353	992.795	2889.238
leverage	2.275	0.648	1.043	1.788
beta	0.739	0.414	0.684	1.033
concentration1	0.451	0.270	0.440	0.620
state	0.070	0.000	0.000	0.000
individual	0.240	0.000	0.000	0.499
corpo	0.071	0.000	0.000	0.000
managment	0.328	0.000	0.000	0.708
institutions	0.223	0.000	0.167	0.307
votes_over_shares	0.015	0.000	0.000	0.000

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The sampling process begins with obtaining financial data on 782 companies listed on the Warsaw Stock Exchange between 2016 and 2021. Subsequently, the following selection criteria are applied:

1. firms listed on the NewConnect (alternative trading system) are excluded due to due to the utilisation of disparate accounting standards, limited liquidity, and distinct financial reporting regulations (389 companies),
2. financial companies are excluded due to the different formats of financial statements (Vieito et al. 2011) and other performance metrics used in terms of banks, insurers, and investment firms (53),
3. companies with a market capitalisation below 100 million PLN in 2016 are excluded as the lack of liquidity might be a source of substantial measurement errors (169),
4. foreign firms, restructuring companies, delisted companies, companies with missing data, and firms having free float below 10% of outstanding shares are excluded (45).

Aluchna and Kamiński (2017) have shown that there is no survivorship bias among when regressing performance indicators over ownership structure. That enables filtering out delisted companies, and, as a results, obtaining a balanced panel.

The final sample consists of 126 firms observed through 6 consecutive year, what makes 756 observations in total. All financial data comes from audited consolidated financial statement that adhere to International Financial Reporting Standards (IFRS). Although the selection criteria are strict, they should benefit the variables' quality and in consequence, decrease the risk of measurement errors.

Table 2: Descriptive statistics of independent variables by industry

industry	n companies	ROA			Tobin's Q approximation		
		mean	median	sd	mean	median	sd
Communication Services	6	0.051	0.058	0.063	1.170	1.030	0.659
Consumer Discretionary	24	0.082	0.068	0.104	2.148	1.165	2.634
Consumer Staples	7	0.060	0.059	0.053	1.085	0.745	0.771
Energy	4	0.061	0.060	0.062	0.594	0.578	0.266
Health Care	10	-0.040	0.030	0.260	2.582	1.352	2.916
Industrials	28	0.056	0.050	0.070	0.937	0.711	0.742
Information Technology	9	0.150	0.053	0.271	2.994	0.762	6.148
Materials	16	0.063	0.049	0.054	0.753	0.699	0.379
Real Estate	14	0.055	0.056	0.034	0.645	0.628	0.223
Utilities	8	0.007	0.021	0.048	0.447	0.437	0.185

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The selection of variables employed in this paper builds on frameworks established in several studies on the implications of ownership structure on firm performance. In order to provide more comparable evidence, firm performance is measured by two distinct dependent variables. The first performance measure is the return on assets (ROA), a common accounting performance measure used by several scholars. The selection of this variable is also aligned with facilitating direct comparative analysis with the study conducted by Aluchna and Kamiński (2017). This deliberate approach affords the opportunity to discern the persistence of the relationships scrutinised within their scholarly inquiry. Since minority owners' expropriation may take place regardless of the firm's financial performance, we introduce an additional dependent variable. To mitigate data availability issues, the second explained variable (Q) is an approximation of Tobin's Q (Chung and Pruitt 1994), which is reported to explain at least 96.6% of the variance of the original indicator, proposed by Lindenberg and Ross (1981), see Equation (1). The choice of this alternative dependent variable is dictated by the relatively small share of bonds in total value of Polish public companies' total liabilities (Białek-Jaworska and Krawczyk 2019) and limited availability of market data on existing and historic issues. The difference between Tobin's Q and the approximation of Tobin's Q lies in that the approximation employed in this study assumes that the replacement value of firms' assets and the market value of the firm's debt are equal to their respective book values.

$$\text{Tobin's Q approximation} \approx \frac{\text{market capitalisation} + \text{debt}}{\text{total assets}} \quad (1)$$

Table 3: The type of the first shareholder by industry

industry	C	I	IP	M	S	total
Communication Services	12	6	6	12	0	36
Consumer Discretionary	6	17	29	92	0	144
Consumer Staples	24	0	7	11	0	42
Energy	0	0	0	6	18	24
Health Care	0	6	22	32	0	60
Industrials	19	16	80	41	12	168
Information Technology	15	9	0	30	0	54
Materials	0	6	36	30	24	96
Real Estate	13	22	8	35	6	84
Utilities	1	0	12	0	35	48

Note: Every single year considered separately.

The ownership concentration of the largest shareholder (*concentration1*) is measured as a percent of the total outstanding shares owned by the largest blockholder. Following the methodology used by Claessens et al. (2002) and Villalonga and Amit (2006), the divergence between cash flows and voting rights (*votes_over_shares*)

is calculated as the difference between cash flow and voting rights of the largest shareholder of the firm. To capture the effects of shareholders' heterogeneity, we introduce explanatory variables denoting state ownership (*state*), corporate ownership (*corpo*), institutional ownership (*institutions*), managerial ownership (*management*), and individual ownership (*individual*). All variables denote the percent of total outstanding shares among the three largest shareholders whose stake exceeds the five percent threshold. Since investors exceeding five percent ownership are obliged to report that, the 5% threshold criterion allows us to avoid measurement errors. As compared to studies using binary variables, the percentage measurement of ownership stake should increase the quality of obtained results as the relative size of shareholding plays a pivotal role in incentivising monitoring (Shleifer and Vishny 1994).

The utilised dataset also includes variables related to the general characteristics of the company. The size of the company is measured by the logarithm of total assets at the end of the financial year in millions of PLN (*assets*). The financial leverage (*leverage*) is calculated as Equation (3). Finally, the firm-specific risk is measured using β coefficient (*beta*). The β coefficients were calculated separately for each year and stock using Equation (2). The minimisation problem was solved using R-builtin ordinary least squares (OLS) estimator

$$\hat{\beta}_{OLS} = \arg \min_{\alpha, \beta} \left\{ \sum_{i=1}^n (y_i - \alpha - \beta x_i)^2 \right\} \quad (2)$$

where:

n – number of business days within a given year

$y_i |_{i=1, \dots, n}$ – a specific stock daily close prices

$x_i |_{i=1, \dots, n}$ – a WIG index daily close valuations

$$\text{leverage} = \frac{\text{Total Liabilities}}{\text{Total Shareholders Equity}}. \quad (3)$$

The variables *mean_roa* and *mean_q_tobin_approx* contain arithmetic means of the dependent variables for each industry and year. To categorise firms into the different industries, global industry classification standards (GICS) were applied.

The data was carefully checked for outliers with regard to dependent and control variables. We noticed some abnormal behaviours, namely:

- i) *equity* < 0
- ii) *roa* < -1
- iii) *leverage* < 0 or > 5, up to 300

for 7 different firms: BML, CCC, ENT, ZWC, MAB, MDG and OTS. To account for that, we applied four steps:

1. the values of the variable *equity*, below its 1st percentile were replaced by it,

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Table 4: Pearson's correlation between explanatory variables

	log_assets	leverage1	beta	concentration1	state	individual	corpo	management	institutions	votes_over_shares
log_assets										
leverage1										
beta	0.32 (0.000)									
concentration1	0.16 (0.000)									
state	0.44 (0.000)	0.19 (0.000)	0.26 (0.000)							
individual				-0.24 (0.000)						
corpo				-0.12 (0.001)	-0.12 (0.001)	-0.22 (0.000)				
management	-0.36 (0.000)			-0.30 (0.000)	-0.30 (0.000)	-0.41 (0.000)	-0.30 (0.000)			
institutions		0.13 (0.000)		-0.14 (0.000)	-0.14 (0.000)	-0.25 (0.000)	-0.13 (0.000)	-0.27 (0.000)		
votes_over_shares				-0.15 (0.000)	-0.15 (0.000)	-0.23 (0.000)	-0.10 (0.004)	0.23 (0.000)		

Note: Only non-diagonal correlations above 0.1 absolute value reported.

2. variable *leverage* was recalculated according to Equation (3),
3. variable *leverage* was winsorized using 1% tails (pre-calculated on original *equity*),
4. values of *roa* below -1 were replaced by -1 .

As the variable *equity* was not directly used in our regression models and the majority of issues related to *leverage* were the results of negative *equity*, we decided not only to winsorise *leverage*, but first to adjust for its components, using steps 1-2.

4 Methodology and empirical results

To verify the proposed hypotheses, we introduce regression models separately for both considered performance metrics, respectively, ROA and Tobin's Q approximation; see Equation (4) and Equation (5). Both specifications share all explanatory variables. We originally intended to also include average values of dependent variable by industry and year to control for time-varying industry effects. This would be numerically feasible as those averages would be calculated for each year separately, thus varying over time and not collinear with fixed effects estimated via within transformation. However, as we show in Appendix, the empirical sample on which we work does not contain evidence proving, in a statistically rigorous way, that those industry-wise averages differ over time. The effects differ only between industries, which is clearly noticeable when looking at Table 2. To account for those time-invariant effects, however, including individual intercepts (either fixed or random) is sufficient, as there exists a deterministic implication between the firm identity and its' industry. That means that both factors will be automatically controlled for if either fixed or random effects least-squares estimators are applied, see Aluchna and Kamiński (2017). As long as the industry-wise effects are not the topic of the research itself, that is a sufficient solution.

One more alternative way to include the second grouping factor is to use ML-estimated linear mixed models instead of least-squares panel regression typical econometrics. These allow for an arbitrary number of distinct random effects along different grouping factors and random coefficients. For the purpose of comparison, we estimated such linear mixed model with two separate random effects for a company and the industry. However, the residuals were heavy-tailed, following a scaled t-student distribution, which violates the most important assumption behind linear mixed models, making reliable inference impossible. Additionally, linear mixed models in a general form do not provide exact p -values for parameter estimates, as their standard errors do not follow t-student distribution under null hypothesis, see Bradic et al. (2020), making the use of the numerical approximations required. Finally, as already shown, individual-level random effects will include the industry ones anyway if the latter were

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not separately included in the specification. Because of all that, we decided to further continue our analysis using least-squares panel regression models.

Table 5: Estimation results

	(1)	(2)
dep. var.	roa	log(q_tobin_approx)
model	fixed	random
(Intercept)		1.2642 ** (0.5002)
beta	-0.1065 (0.0842)	-0.5456 ** (0.2626)
leverage	-0.0006 (0.0014)	-0.0059 * (0.0034)
leverage:beta	0.0012 (0.0034)	0.0156 * (0.0081)
log(assets)	0.0639 ** (0.0283)	-0.1999 *** (0.0444)
log(assets):beta	0.0075 (0.0110)	0.0490 (0.0348)
concentration1	-0.1357 (0.1162)	-0.5347 * (0.2729)
state	0.2196 (0.1681)	0.2624 (0.5318)
corpo	0.2235 * (0.1233)	0.6330 * (0.3686)
institutions	0.1814 (0.1242)	0.1149 (0.3994)
managment	0.3835 ** (0.1700)	0.4999 (0.4198)
individual	0.3255 ** (0.1473)	0.4509 (0.3974)
votes_over_shares	-0.0602 (0.1532)	0.3299 (0.7148)
R^2	0.1219	0.1289
Adj. R^2	-0.0727	0.1148

Note: * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$.

The estimators for both regressions were chosen between fixed and random effects using the Hausman test. We did not follow the approach of i.e. Anderson and Reeb (2003), who focused only on fixed effects estimator without further justification. We argue that as the random effects estimator is more statistically efficient under the random effects assumption holding, see Clarke et al. (2010), considering it will improve the reliability of inference. Also Clarke et al. (2010) concluded that the random effects model should be especially considered when the sample selection procedure is well understood, which is the case in terms of this research. We did

not test for poolability as the descriptive statistics has already shown significant differences between the industries which make the model with all parameters holding globally extremely unrealistic.

Moreover, following the procedure applied by Aluchna and Kamiński (2017), we tested (using Hausman test again) for endogeneity, comparing the models estimated with and without instrumental variables (IV). We considered all variables possibly endogenous and used their first lags as an independent instruments.

As presented in Table 7 the empirical results of the specification tests led us to estimate fixed effects for ROA and random effects for Tobin's Q approximation. Based on those results we also conclude there is no severe endogeneity, as in both cases we do not have evidence to reject null hypothesis, of both models being correctly specified, by a large margin.

We found the empirical distribution of q_tobin_approx to be positively skewed and truncated at zero but without reaching it, which suggests applying log-transformation. We also found no theoretical justification to insist on linear character of dependence between q_tobin_approx and explanatory variables. The residual term estimates from a Equation (5) with a log transformation are symmetrically distributed, being very close to Gaussian. To confirm it, we applied the Kolmogorov-Smirnov test, with null hypothesis stating that the data follow the normal density with parameters obtained via MLE adjusted for outliers. We obtained a p -value equal to 0.09436, slightly above the common critical level 0.05, which led us to keep the q_tobin_approx log-transformed, as the residuals do not suggest a misspecification problem.

As shown in Table 4, the correlation between the explanatory variables is low, at max 0.44. Thus, we do not risk multicollinearity between explanatory variables inflating the standard errors of the estimated models. To confirm this we also tested for aliases using *plm* library builtin functionality, which further stated that no columns are linearly dependent. We also considered interaction terms between control variables and we found *leverage:beta* effect to be significant in both regression models. Moreover, we decided to keep the *log(assets):beta* interaction, even if the respective coefficient was not significant. This makes sense as we have already shown that those variables are relatively strongly correlated, and accounting for situations in which they behave differently can stabilise the model. We did not consider interaction terms including hypotheses-related variables, as within the scope of this work we focus on unconditional relations between ownership structure and firms performance.

Finally, both estimated models were tested for serial correlation (Breusch-Godfrey test) and cross-sectional heteroskedasticity (Pesaran's CD test). The results, see Table 6, lead to the clear conclusion that the estimated models suffer from both mentioned issues. To account for this, we decided to utilise a robust standard errors estimator. A common approach, followed by Anderson and Reeb (2003), is to apply the so-called Huber-White estimator. However, such an estimator accounts primarily for heteroskedasticity, leaving serial correlation unaddressed. Thus, we considered a better alternative, namely White-Arellano estimator, an extension of White estimator

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adjusted specifically for panel data. As explained by Millo (2017), White-Arellano estimator clustered by entities is a right choice, because it is suited for short panels like ours and accounts for cross-sectional heteroskedasticity and serial correlation in any form. To improve comparability with *Stata*, which is a very common choice for panel data econometrics, we parameterised the White-Arellano estimator so that it uses the same small sample correction weighting scheme, as implemented in the *R* package *plm*, see Croissant and Millo (2018).

We are aware that using using robust SE estimators leads to the loss of power, but because of our sample being relatively large we do not consider this an issue. Moreover, the robust SE estimator should be applied only if the parameter estimates are consistent (especially asymptotically unbiased). But again in our setting we found no evidence of a miss-specification of a mean model. Thus we assume that our models are correctly specified and only suffer from heteroskedasticity and serial correlation, which both was proved empirically. That makes usage of the robust SE estimators justified.

The final results involving robust standard errors are presented in Table 5. The R^2 were calculated using *R plm* default “cor” method, which means they represents the “coefficients of correlation between the fitted values and the response”, see Croissant and Millo (2018). We obtained R^2 estimates around 12-13% being on pair with the results of other (financial) econometric regression studies, see deHaan (2020); Faccio and Masulis (2005); Barbopoulos et al. (2018). Low adjusted R^2 is not a concern, as this measure was designed to work only for jointly-estimated regression models. For two-stage within estimator, it treats the within-group means as typical parameters, thus for the wide panels it will be inevitably very low and not suitable for comparison with other estimators. Due to the mechanism of two-stage estimation, the large number of estimated fixed effects may, but do not have to increase the adjusted R^2 (deHaan 2020), while it will always severely drive down the classical, adjusted R^2 for the wide panels. Aluchna and Kamiński (2017) did not report the R^2 among the other results of their study thus we cannot provide a comparison here.

We observe that for ROA all the control variables are significant. Additionally, only two variables are insignificant with regard to the first model. We observe that, with regard to the ownership structure, ROA is primarily positively associated with large proportions of management (CEO), individual (family) and corporate (industrial) owners. For the Tobin’s Q approximation, however, we observe only the corporate ownership to be positively and significantly associated with the dependent variable. Moreover, the Tobin’s Q approximation is strongly, negatively associated with *concentration1*. For every 10 extra percentage points of shares owned by the largest blockholder, the Tobin’s Q approximation is approximately $0.5347 \times 100\% = 53.47\%$ lower.

$$\begin{aligned}
 roa = f(\log(assets), leverage, beta, concentration1, state, corpo, \\
 institutions, managment, individual, votes_over_shares)
 \end{aligned}
 \tag{4}$$

$$\log(q_tobin_approx) = g(\log(assets), leverage, beta, concentration1, state, corpo, institutions, management, individual, votes_over_shares) \quad (5)$$

Table 6: Residual term test results

test	Test p -value given dependent variable	
	ROA	Tobin's Q approximation
Breusch-Godfrey ^a	< 0.0001	< 0.0001
Pesaran's CD ^b	< 0.0001	< 0.0001

Note: ^a H1: serial correlation in idiosyncratic errors.

^b H1: cross-sectional dependence.

Table 7: Hausman test p -values by dependent variable and hypothesis

null hypothesis	ROA	Tobin's Q approximation
random effects	0.000	0.242
exogeneity	0.422	1.000

5 Discussion

Results of the panel study indicate significant and negative effects of ownership concentration by the largest shareholder on approximation of Tobin's Q, being a market-measured firm performance. Moreover, looking from the point of view of confidence intervals, the coefficient of ownership concentration is small enough relative to its standard deviation such that the whole 95.45% confidence interval fits below zero, according to *68-95-99.7 rule*. It provides additional confirmation for the first hypothesis, emphasising the magnitude of the second type of agency costs. Even though the results do not show the significance of the link between ownership concentration and reported return on assets, the findings are partially consistent with those of Aluchna and Kamiński (2017) and Wang and Shailer (2015). The market-based performance metric also supports the findings of Aluchna et al. (2019), who note that ownership concentration leads to a reduction in dividend payouts, highlighting the implications of private benefits of control extraction. The strong and significant negative relationship between ownership concentration and firm performance stands in striking contrast to the conclusions of scholars following the monitoring reasoning (Hanousek et al. 2007, Moscu et al. 2015). Since the market value of firms' assets decreases with ownership concentration, it can be argued that in the light of weak

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external corporate governance mechanisms, the largest shareholders might expropriate minority owners regardless of firms' accounting performance.

The results of the panel data regression models do not indicate a significant link between excess of control rights over cash-flow rights and firm performance measured by both ROA and approximation of Tobin's Q. The insignificance of this variable could also be attributed to the fact that controlling shareholders might implement different control-enhancing mechanisms by a way of including certain privileges directly in companies' charters. Those privileges include provisions such as the exclusive right to appoint a fixed number of directors or the limit of votes at the general meeting that can be exercised by shareholders (other than controlling shareholder). Whilst these provisions cannot be quantified and hence, they do not fit the methodology applied in this panel study, we believe that the inclusion of different control-enhancing mechanisms could provide additional insights.

The results of the panel study indicate the significance of shareholder heterogeneity in regard to firm performance measured by both approximation of Tobin's Q and accounting performance metric. The model does not provide confirmation for H3, as the relationship between state ownership and firm performance is insignificant. This is consistent with the results of a panel study conducted by Aluchna and Kamiński (2017). We believe that insignificance could be associated with a low frequency of transactions conducted by the government. Since in most observations, state ownership remains at a constant level throughout the studied period, a fixed-effects model might fail to capture the possible link between state ownership and accounting-based performance measure.

The results of the regression model showed a positive and significant relationship between both performance measures and ownership by corporate investors. This finding is consistent with observations made by numerous scholars investigating the industry investors' role in emerging equity markets (Aluchna and Kamiński 2017, George and Kabir 2012). It is also in line with the reasoning that in the context of emerging markets, firms benefit from synergistic effects and transfer of know-how associated with ownership by industry investors (Morck et al. 2005). In addition, in our sample industry investors are generally dominant shareholders. Hence, our findings contradict the findings of scholars, who provide evidence on the expropriation of minority shareholders in publicly listed subsidiaries (Atanasov et al. 2010, Chernenko et al. 2012).

In terms of managerial ownership, our model shows a positive and highly significant effect, that it has on firm performance, providing some confirmation for H5. It is in line with the results of the meta-analysis by Iwasaki et al. (2022), who investigates the implications of managerial ownership in Eastern European EU member states. Our finding is also consistent with the agency theory by Jensen and Meckling (1976), supporting the alignment of interests reasoning.

The link between large individual investors and firm performance is found to be strongly positive and most significant among all shareholder types. This finding

is consistent with Gugler et al. (2014), who investigates the role that large individual investors play in CEE equity markets. The authors conclude, that when large investors employ professional managers and provide active monitoring, firms achieve better performance, as compared to those controlled by state or managers. Therefore, large individual investors can provide efficient monitoring when external corporate governance mechanisms are not strong. Since many of the largest individual shareholders are identified to be founders of investigated companies, we interpret this effect also as evidence supporting the stewardship theory (Miller and Le Breton-Miller 2006).

The effect of ownership by institutional investors is observed to be positive, albeit weakly significant. We believe that, as compared to other shareholders, the weaker influence of institutional investors on firm performance can be associated with the size of stakes they hold (Table 2). Following the argument made by Shleifer and Vishny (1994), investors with relatively small stakes may lack the incentive to bear monitoring costs. Nevertheless, the results are consistent with the conclusions of several scholars observing the positive association between ownership by financial investors and firm performance in the context of post-transitional economies (Lin and Fu 2017).

In addition to our main research objectives, our model delivers some insights into other factors influencing firm performance. The firm risk control variable denoting volatility of the firm in relation to the overall market is highly significant and negative both for the market-based and accounting-measured performance metrics. The relationship between firm size and return on assets is found to be significant and positive. This is consistent with Aluchna and Kamiński (2017) and reasoning that larger enterprises achieve higher profitability as a result of economies of scale (Dogan 2013, Lee 2009). On the other hand, firm size is observed to have an adverse and significant effect on market-measured firm performance. Gala and Julio (2016) note that smaller firms have significantly higher corporate investment rates. Therefore, it can be argued, that their Tobin's Q approximation accounts for brighter growth prospects. Finally, the leverage is negatively and significantly linked to the reported return on assets. Even though various debt variables have been employed by scholars to control for leverage, this is generally consistent with existing literature (Aluchna and Kamiński 2017).

6 Conclusions

This paper explores the links between ownership structure and firm performance in the context of the market transitioning to a status of a developed market. Using a unique sample of 126 Polish non-financial firms listed on the Warsaw Stock Exchange between 2016-2021, we introduce regression models, controlling for firm risk, size, debt, and industry. The financial data is collected using Eikon Thomson Reuters. To identify the ultimate owner and gather data on special-voting rights, we manually review 756 annual activity reports. In order to make our study more comparable, we employ two alternative explanatory variables, an approximation of Tobin's Q, and

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a reported annual ROA. Our modeling pipeline consists of both fixed and random effect models. We present the ones chosen on the basis of Hausman's tests results. Our results indicate that ownership concentration by the largest shareholder is adversely related to market-measured firm performance. This is consistent with the findings of a number of scholars investigating this relationship in the context of emerging markets. Our findings provide support for the reasoning that in light of weak external corporate governance mechanisms, the pivotal agency conflict is the one between controlling shareholders and minority owners (Maury and Pajuste 2005). Our results on the implications of divergence between cash-flow and control rights on firm performance measured as ROA and approximation of Tobin's Q do not indicate any significant relationship. Referring to shareholder identity, we found that ownership by industry investors, financial institutions, managers, and individual shareholders is positively related to firm performance measured by ROA. The effects are strongest and most significant in terms of management, industry, and individual investors. In addition, the ownership by corporate investors is identified to be positively linked to the approximation of Tobin's Q. No significant links are observed in terms of ownership by the state, possibly due to a low frequency of transactions completed by the state during the investigated period.

Our study makes a number of contributions. First, the results emphasise the significance of second type agency costs and provide support for reasoning that in a post-transition environment, large blockholder do not provide efficient monitoring and that ownership concentration creates a substantial risk of expropriation. Secondly, our findings provide evidence on the significance of shareholder heterogeneity. Finally, to our best knowledge, this is the first study investigating the link between dual-class share structure and firm performance in the context of Polish enterprises. Referring to the practical implications of our study, our findings generally contradict those of scholars exploring parallel correlations in countries with strong external corporate governance mechanisms. This highlights the role that regulators play in the development of efficient equity markets. In addition, our study provides useful insights for investors in terms of highlighting the channels through which ownership and control may affect portfolio firm performance. There are a few possible research gaps that future studies could explore. First, considering the dynamic regulatory environment, it would be beneficial to replicate our analysis in a few years to verify if the identified effects persisted. Secondly, this panel study concentrates exclusively on Polish enterprises listed on the Warsaw Stock Exchange. Expanding the sample size by including firms from other post-transition Central-European economies could provide evidence of the parallelity of this link. Finally, the inclusion of non-quantitative special control rights included directly in companies' charters could provide more insights into the effects they have on firm performance.

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A Appendix

In this Appendix, we intend to provide detailed inference about the time-invariance of industry-wise averages of both ROA and log-transformed Tobin's Q approximation. It was extracted from the main text as an appendix because the whole procedure of proving the time-invariance of averages is fairly distant from the main topic of this work, while it also requires the use of less common statistical techniques which can be unfamiliar to researchers primarily focused on finance and/or panel data econometrics. Figure A1 shows that the most industries persist in their averages with respect to both dependent variables. However, we see two noticeable violations (industries deviating from the global mean), at least at a scale the data is visualised, namely *Energy* and *Health Care*. With regards to *Energy* industry, such a volatility is reasonable as this is the least represented one, consisting only of 4 firms, see Table 3. The same statement cannot be made in respect to *Health Care*, being represented by as many as 10 companies. Both those violations, noticed through graphical analysis made us suspect that the industry-wise averages of dependent variables can be actually varying over time strong enough to justify their inclusion in panel regression model.

To further examine this issue and obtain decisive conclusions, we decided to evaluate the hypothesis that within-industry averages of dependent variables are time-invariant in a statistically rigorous way. The standard approach in this case is two-way parametric, fixed-effects ANOVA, testing significance of two categorical grouping factors and their interaction (which is clearly of main interest here). Assuming both grouping factors to be indexed by i and j respectively, it can be formulated as Equation (6).

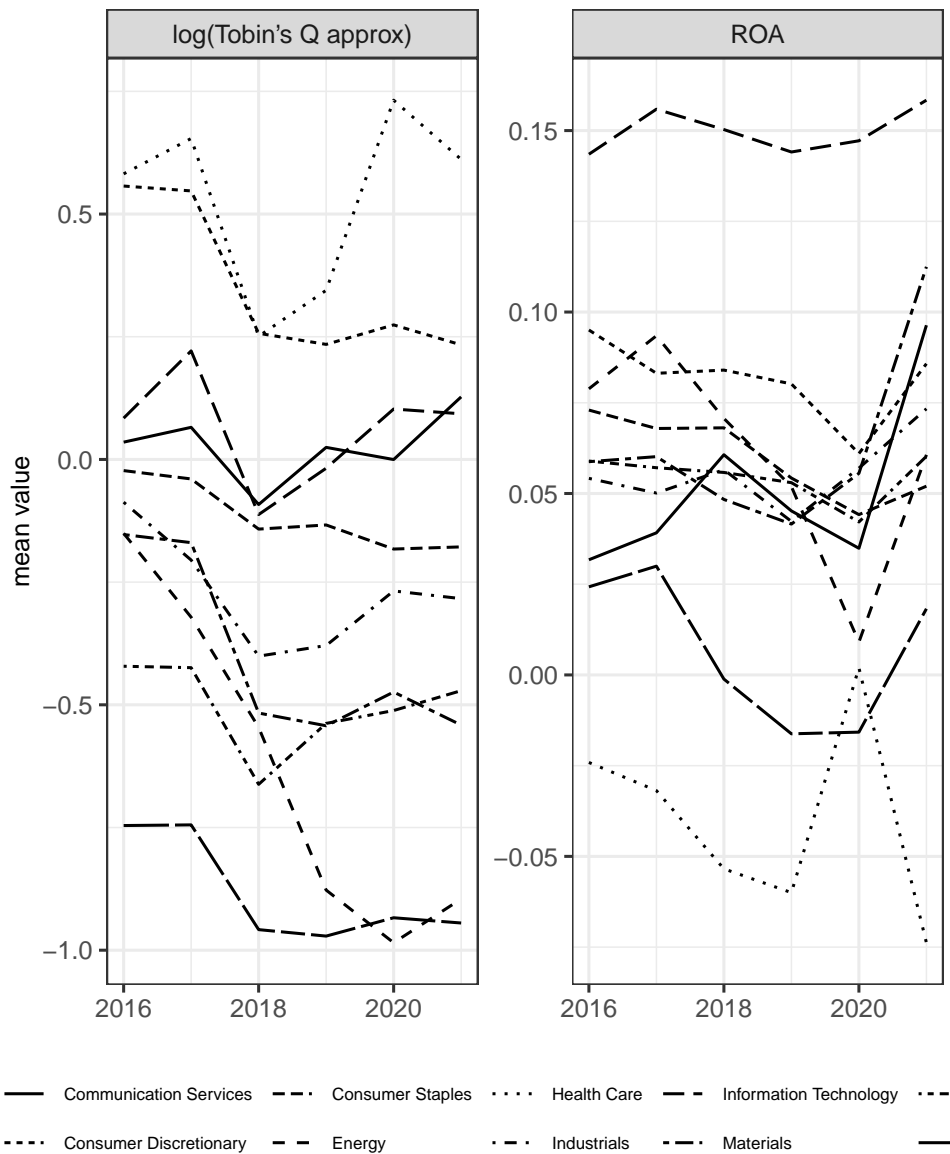
$$Y_{ij} \stackrel{i.i.d.}{\sim} \mathcal{N}(\mu_{ij}, \sigma^2) \quad \text{where} \quad \mu_{ij} = \mu + \alpha_i + \beta_j + \gamma_{ij} \quad (6)$$

However, the validity of inference using such formulated ANOVA is conditional on the fulfilment of three assumptions (Sokal and Rohlf 2012), arising directly from Equation (6):

1. residuals are normally distributed: $Y_{ij} - \mu_{ij} = \varepsilon_{ij} \sim \mathcal{N}(0, \sigma^2)$,

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Figure A1: Dependent variables means by year and industry



2. within-group variances are homogeneous: $\forall_{ij} \sigma_{ij} = \sigma = const$,
3. the observations are independently sampled.

There are no reasons to conclude that assumption 3 is violated. However, the two other assumptions need to be properly tested based on an empirical sample.

To evaluate whether the residuals are normally distributed, we applied the Shapiro-Wilk test (Royston 1982), proven to be the most efficient (with respect to statistical power) among all popular univariate normality tests, see Mohd Razali and Yap (2011). Unfortunately, the results require the rejection of the null hypothesis of normality in both cases, see Table A8. Additional graphical analysis (omitted here) shows that for ROA, the distribution is symmetric but heavily tailed. Simultaneously, the empirical distribution of the log-transformed Tobin's Q approximation shows a noticeable right skewness. Lack of normality becomes an issue because the elementary test for variance homogeneity, Bartlett's one, requires (assumes) it, see Sokal and Rohlf (2012). We decided to apply Fligner-Kileen nonparametric test instead, by definition robust to violations from normality (Conover et al. 1981). As shown in Table A8, the test results are not fully conclusive, supporting (borderline) the null hypothesis of homogeneity at critical level 0.05 for log-transformed Tobin's Q approximation and its' rejection for ROA. We thus conclude that the problem exists, but is neglectable.

Table A8: Combined test results related to the hypothesis that industry averages are time-dependent

	Dependent variable	
	ROA	log(Tobin's Q approximation)
Shapiro-Wilk Test ^{ac}	< 0.0001	< 0.0001
Fligner-Kileen Test ^{ad}	0.0305	0.0668
Two-way ANOVA ^{ae}	≈ 1	≈ 1
Scheirer-Ray-Hare Test ^{ae}	≈ 1	≈ 1
Two-way Bayesian ANOVA ^{be}	15876	15753

Note: ^a p -value reported,

^b BF_{01} (Bayes factor, H_0 against H_1) reported,

^c H_1 : residuals not normally distributed,

^d H_1 : > 1 group variances differ (with respect to *industry:year*),

^e H_1 : > 1 group means differ (with respect to *industry:year*).

Sokal and Rohlf (2012) argue that violations of both assumptions, unless very severe, do not immediately invalidate the ANOVA F-statistic. Therefore, we estimated the model, given our data sample, obtaining decisive evidence (p -value ≈ 1) to reject the null hypothesis that the grouping factor *industry:year* is insignificant, see Table A8. However, to further ensure that our conclusions are correct, we applied two alternative approaches that account for the violations of the ANOVA assumptions we

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have detected. The first of those is Scheirer-Ray-Hare test, a nonparametric, rank-based method being a generalisation of Kruskal-Wallis test and thus a substitute for two-way ANOVA (Sokal and Rohlf 2012). The second approach, being primarily robust to variance heterogeneity, is a Bayesian ANOVA. We used it to test whether it is more probable that the data was generated by a regression without interaction term (so only including indicators variables of both the company and the year) instead of the one with it. That is equivalent to a standard frequentest procedure of Two-Way ANOVA. In the case of the Bayesian approach, the inference is based on a score named “Bayes factor”, reported in Table A8, which is a ratio of probabilities that the data is generated by one model versus another. Jeffreys (2003) suggests, as a rule of thumb, that $BF_{01} > 100$ means that there exists decisive evidence for the correctness the first model (in our case H0, no interaction term) instead of the second one (H1). Gathering all the results presented in Table A8, we observe that all three tests leave us with the same conclusive results, respectively with p -value ≈ 1 and $BF_{01} \approx 16000$. To summarise the whole investigation we conducted above, we conclude, without a doubt, that the industry-wise effects are time invariant. Therefore, we will not include those dependent variable averages as additional regressors in our models.