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THE IMPACT OF INSTITUTIONAL INVESTORS ON RISK AND STOCK RETURN AUTOCORRELATION IN THE CONTEXT OF THE POLISH PENSION REFORM

The main aim of this paper is to examine the relationship between the increasing share of institutional investors resulting from the pension reform in Poland and stock return autocorrelation as well as risk level on the Warsaw Stock Exchange. The problem under consideration is investigated by applying the M–GARCH model for the individual stocks included in the investment portfolios of the pension funds operating in Poland.

Keywords: *stock return autocorrelation, risk, institutional investors, pension reform*

1. Introduction

Since Chile in the early 1980s, and then Peru in the early 1990s, implemented systematic pension system reform, in many countries the reform of national pension system became a topic that has been repeatedly under discussion. Multinational organizations and economists around the world have been making huge contribution to this debate. The main objections raised against existing pension systems in different countries are: low benefits, widespread contribution evasions, huge number of retirees, benefits paid almost without regard to work history, and inefficient administration. Since the problem of an unpredictable and unreliable pension system remains still unsolved in many countries, this discussion will probably continue also in the future. Pioneering countries which reformed their pension system in time to avoid

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a financial crash, provided valuable hints for other countries, especially those from Eastern and Central Europe, whose pension system operated in relative financial equilibrium only with the help of budget subsidies. It is worth highlighting that Argentina and Colombia are the first countries that have carried out fundamental reform of the pension system through the democratic process. Chile and Peru implemented their reforms under autocratic regimes.

Countries, like Poland, which faced the necessity of reforming their inefficient and costly pension system, followed the World Bank's advice supported by the example of such countries as Chile, Peru and Argentina and implemented one of the permuted forms of the basic framework which is outlined below. The first pillar of the pension system continues the idea of government social insurance guarantee. This component is compulsory, financed on the "pay-as-you-go" basis, which means that the contributions of the current generation of workers are directly transferred to the current generation of retirees. The pension level from this pillar is determined by working record or, more often, a person's contributions to it. The second pillar is a mandatory forced savings program. The concept of this program is quite simple. A fixed percentage of a worker's wages is accumulated on an individual account over the whole working period. Being formed in this way capital is privately managed and invested in an individual account to gain in value and ultimately provide the retirement payments stream. Unlike the first two pillars, the third one is voluntary. It is a savings plan whose target is to supplement retirement income. To make this scheme more attractive for employees, it should be accompanied by a relief from income tax, but as practice shows, this kind of motivation is applied rather occasionally.

Nowadays, the list of countries which transformed their pension system into one of the permutations of the above mentioned basic framework includes Uruguay, which implemented a less ambitious reform in 1995, Bolivia, El Salvador and Mexico as well as several Eastern and Central European countries, namely Hungary, Poland, Slovenia, Croatia, Russia and Kazakstan in Asia.

Given this background, in the next part we review in some detail the Polish pension system reform which took place at the end of the 1990s. The reader who is interested in other instances of pension system reform according to the Latin model is referred to the rich literature that deals with this subject (see, e.g., [7], [14], [20], [28], [32], [35], [40], [41], [44]).

1.1. The pension system in Poland after 1999

The old Polish pension system suffered from an extremely low ratio of working population to pensioners, and unclear pension law causing a great deal of misuse of funds. The Social Insurance Institution (ZUS) was able to operate in relative financial equilibrium only with the help of budget subsidies increasing each year. The level of contributions to the system became so high that any further increase could not be considered by any sagacious politician. To avoid the total collapse of the Polish pension system the work on its reform started at the very beginning of the transition process in the 1990s. The effort undertaken to improve the Polish pension system and to render it free of political manipulation, and at the same time more in line with demographic trends resulted in its fundamental reorganization by the end of the 1990s.

First, the compulsory, financed on the "pay-as-you-go" basis, pillar was launched on 1st January, 1999. The major role in this component is played by the Social Insurance Institution (ZUS). The rule that the pension level from this pillar is determined by a person's contributions to the system is one of the main features which distinguishes the new system from that existing before 1999.

In spring 1999, the second pillar was started. This mandatory, but privately managed part of the reformed system consists of open-ended pension funds (OFE). The pension funds have legal status and their period of operation is unlimited. The scope of their activity is accumulating and investing funds for disbursement to fund members upon reaching retirement age. Each open pension fund is managed and represented in relation to third parties by general pension societies (PTE), operating as joint stock companies. Each PTE establishes a fund and manages it on a non-gratuitous basis. PTE's are authorised to manage no more than one OFE. The Social Insurance Institution transfers 7.3% of remuneration to the OFE selected by the insured person. Contributions paid to OFE are converted into clearance units and booked into individual accounts of the insured. The increase in the value of accumulated funds (contributions) occurs as a result of increasing the number of clearance units (through the payment of subsequent contributions) as well as the growth in the value of a single clearance unit resulting from the investment policy conducted by the selected OFE. The value of a clearance unit, calculated as the relation of an OFE's net assets and the number of issued clearance units, is the basis for calculating the return rate of each OFE and the combined weighted average return of all OFEs. An OFE invests its assets seeking to maximize the security and profitability on funds invested. Investment limits and the mechanism of a minimum required rate of return assure the security of funds entrusted to OFEs. Investment limits determine the investment policy of open pension funds as well as the types of investments (i.e., equities, bonds).

The third (optional) pillar enables employers to register special savings plans which can be offered to their employees within one of the four forms belonging to the so-called Employee Pension Programmes (EPP).

A considerable body of literature has emerged which review these issues, e.g., Chlon et al. [13], Hausner [25].

In the next section, we provide a brief literature survey and specify main conjectures which are subsequently tested in the empirical part of our study.

1.2. Literature survey and main conjectures

Apart from a few distinguishing features, funds operating within the second pillar of the Polish pension system can be seen as a group of institutional investors such as banks, open-ended funds and others. There is a common belief that these institutions invest their members' money based on a long-term calculation. However, the latest evidence shows that this belief may appear to be particularly illusory (see [22]). Nevertheless, their impact on capital market still remains unquestionable and many empirical studies provide evidence supporting such a statement. Schmidt-Hebbel [34], for example, states that Chile's pension reform improved labour-market performance and raised savings, investments, and factor productivity. Investigating direct and indirect links between the fully-funded privately managed pension system and the capital market in Chile, Argentina and Peru, Leofort and Walker [28] found that fund activities lead to reductions in firms' cost of capital, lower price volatility and higher trading volumes. This matter was continued in Winter [44], Börsch-Supan et al. [8], Börsch-Supan and Winter [9], Bosworth and Burtless [10].

It should be noticed that these types of relations were also partially explored in Poland. Gurgul and Majdosz [23] provided evidence of the influence of reform on mean stock returns for companies listed on the Warsaw Stock Exchange (WSE). Some attempt was made to capture changes in the first-order autocorrelation of stock returns after the reform of the pension system in 1999 (see [19]). However, the methodology used by the authors based on a simple regression with a dummy variable and heteroscedastic error term seems not to be relevant to the capturing of such an effect.

In this chapter, we go into the matter of risk and stock return autocorrelation changes in the context of an increasing share of institutional investors. Unlike the above-cited work of Gebka et al. [19], in this study we generate conditional first-order autocorrelation estimates using a multivariate generalized autoregressive conditional heteroscedasticity (M–GARCH) model. The risk of individual stocks is proxied by time dependent conditional beta estimates fitted to series of given firm returns and market portfolio returns.

Now, we will specify the main hypothesis underlying our study. Formulated by Fama in 1970 and extended in the following years the efficient market hypothesis (EMH) states that stock prices reflect immediately all new information available (see [17], [18]). A simple consequence of the weak-form of the EMH is that stock return autocorrelation estimates should all be close to zero and no predictable patterns in stock returns should be observed. However, there exist many theories, and empirical evidence supporting them, that predict a positive impact on autocorrelation in stock returns attributed to institutional investor activities. First, Barclay and Warner [3] built a model where informed institutional investors distribute their trades over time to lower their price impact inducing an increase in stock return autocorrelation. Wang [42] put forward a model based on information asymmetry between traders. The main conclusion is that an increased percentage of informed traders causes an increase in the autocorrelation of stock returns. Institutional investors such as pension funds seem to be well informed. Owing to this institutional ownership, it increases return autocorrelation. Supporting evidence can be found in Arbel and Strebel [1], Sias and Starks [38], Sias et al. [37]. Another explanation for positive autocorrelation dependence on the amount of institutional ownership appeals to positive feedback trading and herding by institutional investors. However, evidence of positive feedback trading and herding is rather modest in the case of pension funds (see [26]). Overall, we expect autocorrelation in stock returns to be essentially higher over the subperiod after 1999 compared to the subperiod prior to the reform.

Secondly, if our hypothesis is true that with the presence of pension funds on the capital market an increase in autocorrelation of stock returns will be observed, the future behaviour of stock prices would become more foreseeable. This means that the level of risk would also be lower. If, with access to more information and professional analysis, pension funds are assumed to be well informed, stock prices are likely to fluctuate more in line with fundamental values (see [28]). Consequently, a decline in risk should occur. Thus, our second prediction is that pension fund activities over the subperiod after 1999 would induce a decrease in risk level on the WSE.

The remainder of this chapter is organized as follows. In Section 2, we outline the basic methodology used to generate conditional first-order autocorrelation estimates and time varying conditional beta estimates. Section 3 provides a brief description of the criteria for firm selection and preliminary findings. Our results for conditional autocorrelation in stock returns and conditional beta are summarized in Section 4. In the same section, we focus our attention on the issue of whether the transfer of insurance contributions to pension funds is a determinant of conditional autocorrelation in stock returns. The last section concludes the paper.

2. Methodology

W. Sharpe's market model is commonly used to estimate a systematic risk of asset *i*. Taking into account the widely-documented fact that daily stock returns appear to be serially correlated, and that one autoregressive term is usually sufficient to shape it, this model is given as:

$$r_{i,t} = \alpha_i + \rho_{i,1} r_{i,t-1} + \beta_i r_{M,t} + \varepsilon_{i,t}, \qquad t = 1, ..., T, \tag{1}$$

where $r_{i,t}$ stands for the return of asset *i* for day *t* and $r_{M,t}$ denotes the return of market portfolio for day *t*. This value is commonly proxied by the return of a market-capitalization weighted stock index. Assuming homoscedasticity of the error term and a lack of autocorrelation of lagged errors, an unconditional ρ_i and β_i can be obtained from the equations:

$$\hat{\rho} = \frac{\operatorname{cov}(r_{i,t}, r_{i,t-1})}{\operatorname{var}(r_{i,t})}, \tag{2}$$

$$\hat{\beta}_i = \frac{\operatorname{cov}(r_{i,t}, r_{M,t})}{\operatorname{var}(r_{M,t})}. \tag{3}$$

Alternatively, one can employ the M–GARCH model to generate the conditional variance-covariance matrix of given time-series and by replacing the unconditional covariance and variance in (2) and (3) with their conditional equivalents get the conditional first-order autocorrelation estimates $(\hat{\rho}_{i,t})$ and time varying beta estimates $(\hat{\beta}_{i,t})$.

In the late 1980s, Bollerslev et al. [6] proposed a model in which each element of covariance matrix ($\mathbf{H}_t$) is a linear function of the lagged squared errors and cross products of errors, and lagged values of all elements of $\mathbf{H}_t$. The VEC(1,1) can be expressed as:

$$\mathbf{h}_t = \mathbf{c} + \mathbf{A}\eta_{t-1} + \mathbf{G}\mathbf{h}_{t-1}, \tag{4}$$

where $\mathbf{h}_t$ = vech($\mathbf{H}_t$), $\eta_t = \operatorname{vech}(\varepsilon_t \varepsilon_t')$ and vech(.) denotes the operator that stacks the lower triangular portion of an $N \times N$ matrix as an $N(N + 1)/2$ vector. **A** and **G** are square parameter matrices of order $(N + 1)N/2$, and **c** is an $(N + 1)N/2$ parameter vector where N denotes the number of time series included in the M–GARCH framework.

Fitting the general VEC model (4) to a set of three time-series, it is necessary to estimate 78 parameters. The huge number of parameters to be estimated has led to the general VEC model being in practice mainly used in the bivariate case. Bollerslev et al. [6] suggest imposing assumptions on parameters to overcome this problem. In the diagonal VEC (DVEC) model the matrices **A** and **B** are assumed to be diagonal. This means that each element $h_{ij,t}$ depends only on its own lag and on the lagged value of $\varepsilon_{i,t}\varepsilon_{j,t}$. Under these assumptions, one has to estimate 18 parameters instead of 78 ($N = 3$).

Bollerslev [5] put forward the constant conditional correlation M–GARCH specification (CCC). If the correlations between conditional variances are assumed to be constant, the conditional covariances are proportional to the product of the corresponding standard deviations. The CCC model is defined as:

$$\mathbf{H}_t = \mathbf{D}_t \mathbf{R} \mathbf{D}_t = [\rho_{ij}\sqrt{h_{ii,t}h_{jj,t}}], \tag{5}$$

where $\mathbf{D}_t = \operatorname{diag}(\sqrt{h_{11,t}}, \sqrt{h_{22,t}}, \ldots, \sqrt{h_{NN,t}})$, $h_{ii,t}$ can be defined as any univariate GARCH model, and $\mathbf{R} = [\rho_{ij}]$ is a symmetric positive definite matrix with $\rho_{ii} = 1$ for any i.

In many empirical applications, the assumption about a conditional correlation constancy seems to be invalid (see [6]). Unlike the CCC model, the dynamic conditional correlation (DCC) model makes the conditional correlation matrix time dependent. Engle (see [15] and [16]) introduced the DCC model defined as:

$$\mathbf{H}_t = \mathbf{D}_t\mathbf{R}_t\mathbf{D}_t\,, \tag{6}$$

where

$$\mathbf{R}_t = \mathbf{Q}_t^{*-1}\mathbf{Q}_t\mathbf{Q}_t^{*-1}\,, \tag{7}$$

$$\mathbf{Q}_t = \left(1 - \sum_{k=1}^{K}\alpha_k - \sum_{l=1}^{L}\beta_l\right)\overline{\mathbf{Q}} + \sum_{k=1}^{K}\alpha_k(\varepsilon_{t-k}\varepsilon'_{t-k}) + \sum_{l=1}^{L}\beta_l\mathbf{Q}_{t-l}\,, \tag{8}$$

$$\mathbf{Q}_t^* = \mathrm{diag}(\sqrt{q_{11,t}}, \sqrt{q_{22,t}}, \ldots, \sqrt{q_{NN,t}})\,. \tag{9}$$

In (6)–(9), $\mathbf{D}_t$ is, as in (5), the $N \times N$ diagonal matrix of time varying standard deviation with $\sqrt{h_{ii,t}}$ on the main diagonal, $h_{ii,t}$ can be defined as any univariate GARCH model, $\mathbf{Q}_t$ is the $N \times N$ positive definite matrix, $\overline{\mathbf{Q}}$ is the $N \times N$ unconditional variance matrix of the standardized residuals, and $\mathbf{Q}_t^*$ is the $N \times N$ diagonal matrix composed of the square root of the diagonal elements of $\mathbf{Q}_t$.

Of course, the brief review of the M–GARCH theoretical frameworks given above does not aspire to be a comprehensive presentation of the many variations of the M–GARCH model which have been developed since the late 1980s up to now. The aim is rather to give some basis for a discussion about the relevant parameterization selection of the M–GARCH model under study. An excellent survey of M–GARCH models can be found in Bauwens et al. [4].

In many papers, including contributions published most recently, the CCC model is applied to generate conditional first-order stock return autocorrelation estimates and time varying beta (see, e.g., [11], [12], [29], [31]), where $h_{ii,t}$ is chosen to be the GARCH(1,1) process. Taking into account the experience of the cited researchers, this class of M–GARCH models seems to be thoroughly sufficient for the empirical proof of our conjectures. However, not all the time-series included in our sample fulfil the assumption about conditional correlation constancy.

We use the test for constant correlation proposed by Engle and Sheppard [16], in which the null and alternative hypotheses can be expressed as:

$$H_0 : \mathbf{R}_t = \overline{\mathbf{R}}, \quad \forall t \in T,$$
$$H_a : \mathrm{vech}(\mathbf{R}_t) = \mathrm{vech}(\overline{\mathbf{R}}) + \beta_1\mathrm{vech}(\mathbf{R}_{t-1}) + \beta_2\mathrm{vech}(\mathbf{R}_{t-1}) + \ldots + \beta_p\mathrm{vech}(\mathbf{R}_{t-p})\,.$$

By this test we found that in many cases the null of constant correlation should be rejected in favour of a dynamic structure. On the basis of the test results, the DCC model is used in all those cases where it appears to be suitable, instead of the CCC model.

Univariate GARCH processes were selected according to the minimum AIC principle. We chose the simplest dynamic correlation structure in (8) for $K = L = 1$. The whole DCC framework was estimated by means of the two-stage estimation procedure developed by Engle and Sheppard [16].

3. Data and preliminary results

To be included in our sample, a given company has to meet the following selection requirements. Firstly, it should be listed on the WSE at least from May 1997. In addition, the company should be reliable and large enough to be quoted on the primary market of the WSE. Secondly, the stock price series of a given firm taken from the PARKIET database should not exhibit a sequence of missing data over at least 500 trading days prior to 19 May, 1999, and also in the following 500 days. In a few cases a single missing piece of data was filled in (as a mean of its direct neighbours). Thirdly, the share of pension funds in the capitalization of a firm which is a candidate for the sample should be significant.

On the basis of these rules, twelve firms representing eight different sectors, namely Construction (Budimex, Echo, Elektrobudowa, Mostostal Zabrze hereafter BDX, ECH, ELB, MSZ, respectively), Chemicals (Jelfa hereafter JLF), Informatics (Computerland hereafter CPL), Light Industry (Novita hereafter NVT), Building Materials (Krosno and Lentex hereafter KRS and LTX), Media (Poligrafia hereafter PLG), Metals (Kęty hereafter KTY), and Motorization (Dębica hereafter DBC) were selected. It is worth noting that the Polish capital market's post-war history is rather modest. Since its launch in the first half of 1991, the WSE has developed quickly and at the end of 1996 there were 83 companies listed on this stock exchange. However, many firms that are now strongly represented in the investment portfolio of pension funds were added to the WSE afterwards and this fact is mirrored in our sample.

After controlling stock price time series for mistakes, dividends, stock splits and other events, we start off with the estimation of unconditional beta and unconditional first-order autocorrelation in stock returns. In the entire study, continuous stock returns are used. Approximating market portfolio returns by returns of the market-capitalization weighted stock index called WIG, we obtained estimates of ρ_i and β_i in (1) by means of the OLS method. The Durbin-Watson statistics are close to 2. Therefore, no significant first-order serial correlation occurs. Next, we generated the conditional variance of corresponding return series as well as conditional covariance between them based on (5) and (6)–(9) and computed conditional equivalents of ρ_i and β_i on the basis of (2) and (3) in which unconditional variances and covariances are substituted for conditional counterparts. In both cases, i.e., under unconditional and conditional estimation, the full sample covering the period from 15th May, 1997 to 21st May, 2001 (500 trading days prior to and after the 19 May, 1999) is used. Our results are summarized in Table 1.

All unconditional ρ_i (except for NVT) are negative and statistically significant. Comparing the unconditional ρ_i with corresponding mean conditional first-order autocorrelation estimates, one can find that the differences between them are rather fundamental. In more than half of the cases, the average conditional ρ_i is positive while the unconditional counterpart is negative. A discrepancy between conditional and unconditional first-order autocorrelation estimates is not surprising and can also be found, for example, in McKenzie and Faff [29]. In this case, however, a specific factor that contributes to this disproportion is likely to exist, making it so striking. If our conjecture that an increase in the share of institutional ownership (caused by pension reform) induces an increase in the autocorrelation of stock returns holds true, then conclusions drawn on the basis of unconditional ρ_i are essentially invalid. The reason is that model (1) is incapable of uncovering such an effect. The strong tendency revealed in the data is probably mirrored adequately within the M–GARCH framework. The final judgements (main results) are presented in Section 4 where the models are fitted to the data over two subperiods.

In the second column, the pension fund share in the capitalization of the company in 2001 is presented (Source: KNUiFE). Unconditional first-order autocorrelation in stock return and unconditional beta were estimated on the basis of (1) by using the OLS method. Standard errors are shown in parenthesis. The superscripts "a" and "b" denote values which are significantly different from zero and one, respectively.

The part of the results concerning risk level (reported in Table 1, columns 6–8) also prompts us to make some interesting remarks. Firstly, all unconditional β_i significantly differ from zero and differ from one, except for CPL where beta is significant, and it does not differ from one. In principle, this time without exception, the same conclusions are true for the mean conditional β_i. Secondly, none of the betas (unconditional and mean conditional) is greater than one. This throws light on the issue as to which one of the firms listed on the WSE is preferred by the pension funds. Note that our sample comprises twelve companies out of the twenty-four in which the pension fund share in capitalization was above 14% at the end of the 2001. All of the sampled firms are marked by beta below unity. Firm-specific effects in the stock returns of these companies (absorbed, for example, into the constant terms α_i) are unimportant and betas below 1 mean that their stocks underperform the market index. Thirdly, only in the case of NVT, is the mean conditional β_i slightly greater than the corresponding unconditional β_i. Apart from this isolated case, these estimates show a reverse relationship. Simultaneously, the conditional beta estimates range from 0.006 (KSR) to 2.480 (BDX) and this is fully reasonable, taking into account results presented by, e.g., Brooks [12]. In the next section, we turn to the comparison of beta and first-order autocorrelation estimates in two subperiods.

Tabela 1

4. Main results

4.1. Subperiod analysis

We divided our sample into two subperiods. The first extends from May 15th, 1997 to May 18th, 1999, the second – from May 19th, 1999 to May 21st, 2001. May 19th, 1999 is chosen as a boundary date, because the first transfer of insurance contributions from the Social Insurance Institution to the OFE took place on this day. Note that in the initial stage ranging over the period from April 1st, 1999 when the pension fund sector in Poland invited the first payment of insurance contributions, the OFE activities were fully devoted to the acquisition of members. Therefore, we use as a boundary date May 19th, 1999, instead of April 1st, 1999. Gebka et al. [19] proceeded in a similar way.

Both subperiods comprise 500 daily observations of each firm included in our sample. We re-estimated our basic models in the subperiods separately, and computed the estimates of beta and first-order autocorrelation in stock returns again. Therefore, if any tendency exists in the data, it should be reflected in the values of these parameters. Our results are reported in Table 2, separately for each subperiod.

Comparing the unconditional estimates of first-order autocorrelation in stock returns prior to pension system reform and after it, one can find ten cases where this parameter increases over time. In the case of JLF, MSZ, and NVT, it even becomes positive though statistically insignificant. However, one should be particularly careful not to make a mistake in inferences, because in model (2.1) the parameter ρ_i is assumed to be constant over time. Alternatively, going on conditional estimates of stock return autocorrelation (see columns 2 and 7 of Table 2), we find that the mean conditional ρ_i increases for BDX, CPL, DBC, ELB, JLF, KTY, MSZ, and NVT. In all of these cases, the mean conditional ρ_i is positive in the subperiod after pension system reform. Except for CPL, JLF, and NVT, it is simultaneously positive in the subperiod after reform and negative in the subperiod prior to reform.

On the other hand, the mean conditional β_i decreases over time in the case of all firms but CPL. In addition, the highest conditional betas in the subperiod after the reform are mostly lower when compared with the subperiod prior to reform. A substantial degree of range around the mean value is presented in both the subperiods.

It is worth noting that the comparison of the unconditional betas within subperiods fundamentally leads to the same conclusion. To answer the main questions of our chapter more formally, we use a technique based on the empirical cumulative distribution function introduced by Gonzales-Rivera (21) and extended by Brooks et al. [12].

Tabela 2

Let

$$\delta_{xy}^{\rho} = \sup_{\rho \in \mathbb{R}} (F_x(\rho) - G_y(\rho)), \tag{10}$$

and

$$\delta_{xy}^{\beta} = \inf_{\beta \in \mathbb{R}} (F_x(\beta) - G_y(\beta)), \tag{11}$$

where $F_x(\rho)$ and $G_y(\rho)$ denote the cumulative distribution function of ρ (belonging to the set of real number) for assets x and y, respectively, and $F_x(\beta)$ and $G_y(\beta)$ stand for the cumulative distribution function of β for assets x and y. From (10) and (11) follow stochastic dominance criteria:

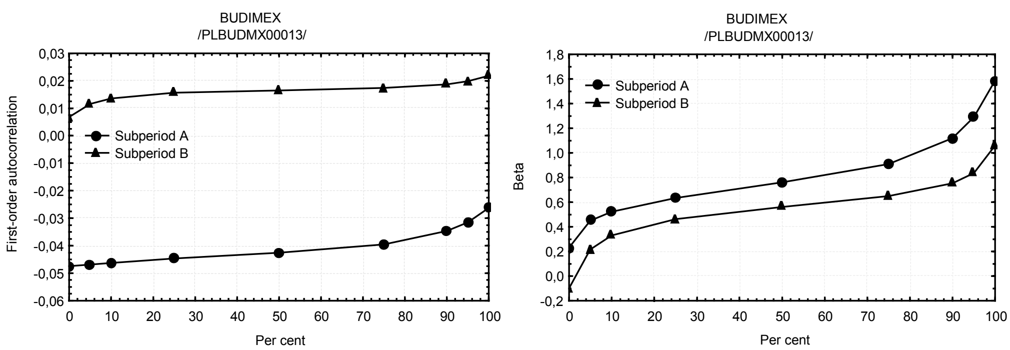
Definition 1. Autocorrelation in stock returns of x is higher than those of y if $\delta_{xy}^{\rho} = 0$.

Definition 2. Risk level of x is lower than those of y if $\delta_{xy}^{\beta} = 0$.

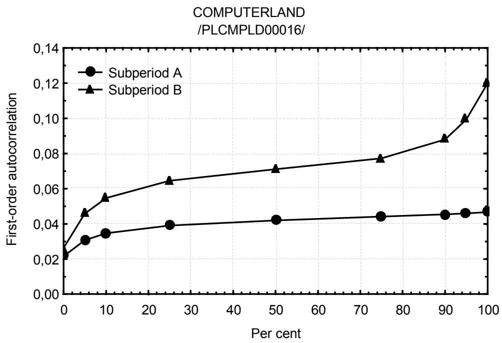
It is very important to make clear that the higher autocorrelation in stock returns of x identified on the basis of Definition 1 does not imply that its stock returns are automatically stronger serially correlated compared to y. This simple reasoning would be true only if the stock returns of y exhibit positive serial correlation.

The empirical cumulative distributions of conditional first-order autocorrelation in stock returns as well as conditional beta for all firms included in our sample are plotted in Figure 1.

a) b)



c)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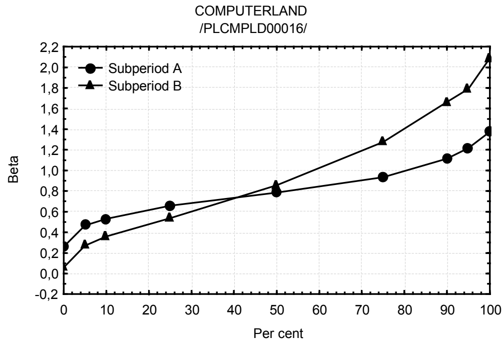
Subperiod A
Subperiod B
First-order autocorrelation
Per cent

d)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Subperiod A
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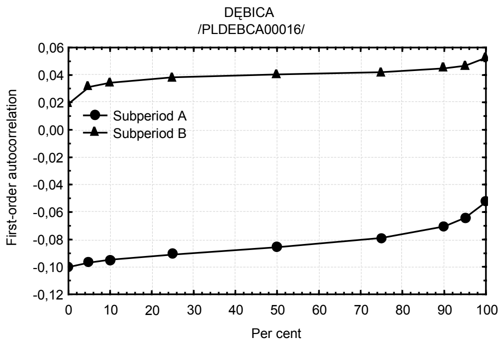
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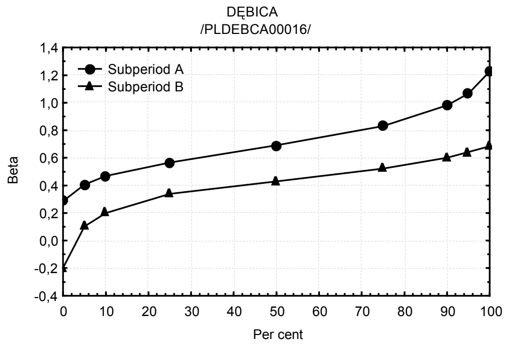
Subperiod A
Subperiod B
First-order autocorrelation
Per cent

f)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Subperiod A
Subperiod B
Beta
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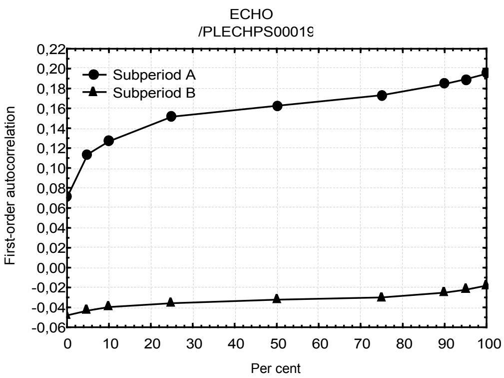
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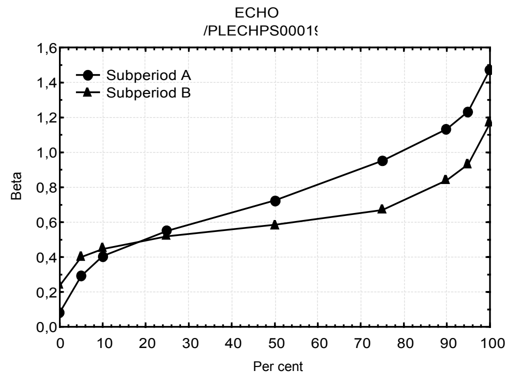
Subperiod A
Subperiod B
First-order autocorrelation
Per cent

h)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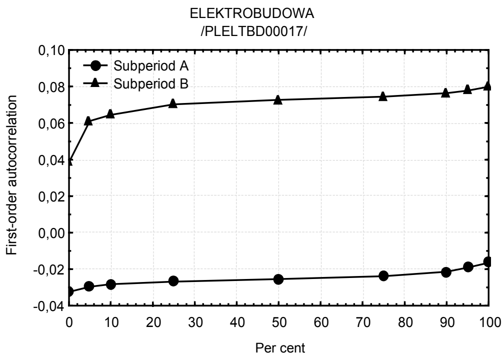
Subperiod A
Subperiod B
Beta
Per cent

i)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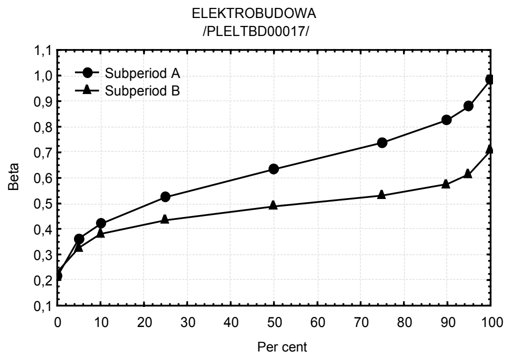
Subperiod A
Subperiod B
First-order autocorrelation
Per cent

j)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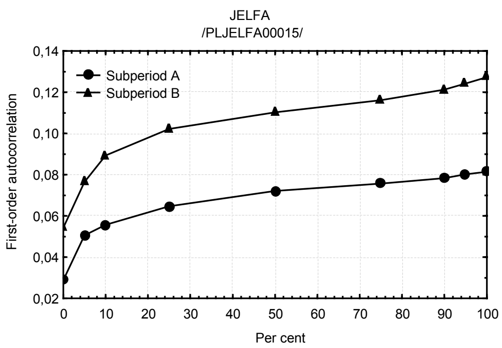
Subperiod A
Subperiod B
Beta
Per cent

k)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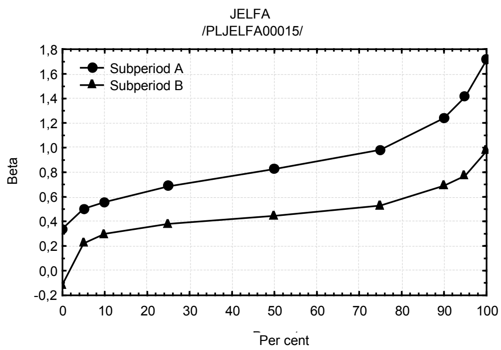
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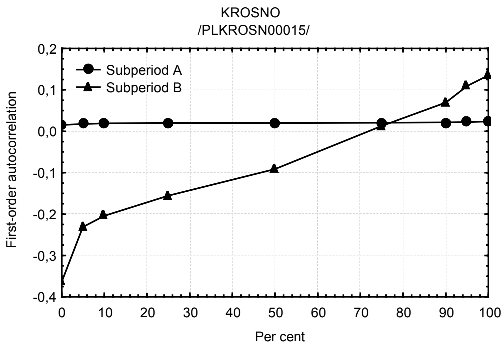
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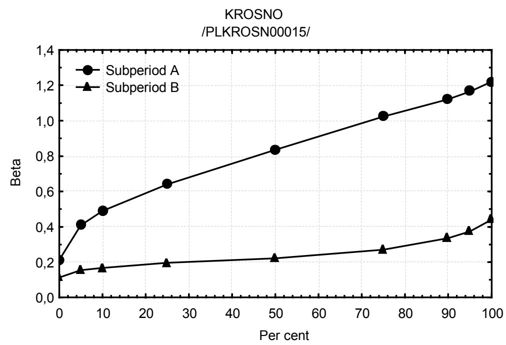
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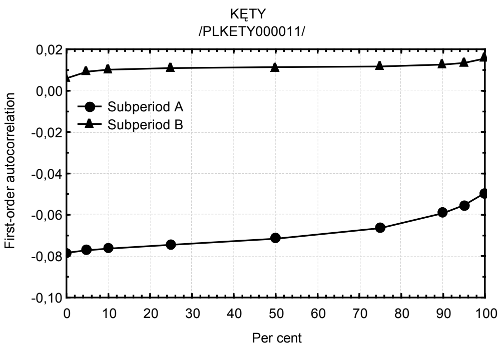
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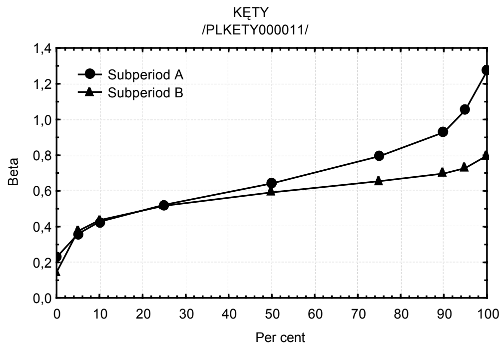
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Per cent

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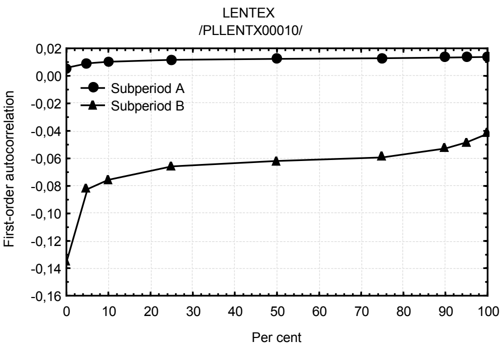
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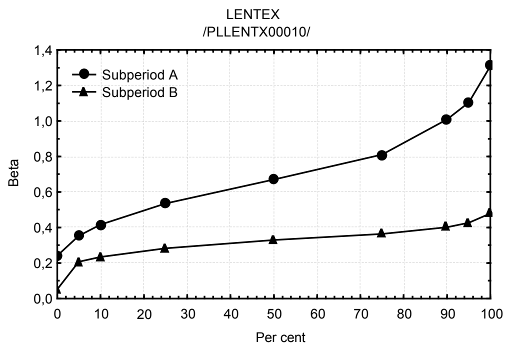
Subperiod B
First-order autocorrelation
Per cent

r)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Subperiod B
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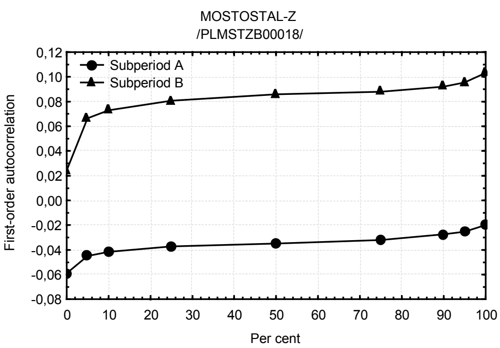
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Subperiod A
Subperiod B
First-order autocorrelation
Per cent

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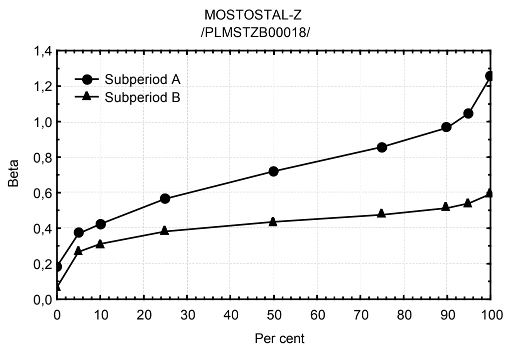
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Subperiod B
Beta
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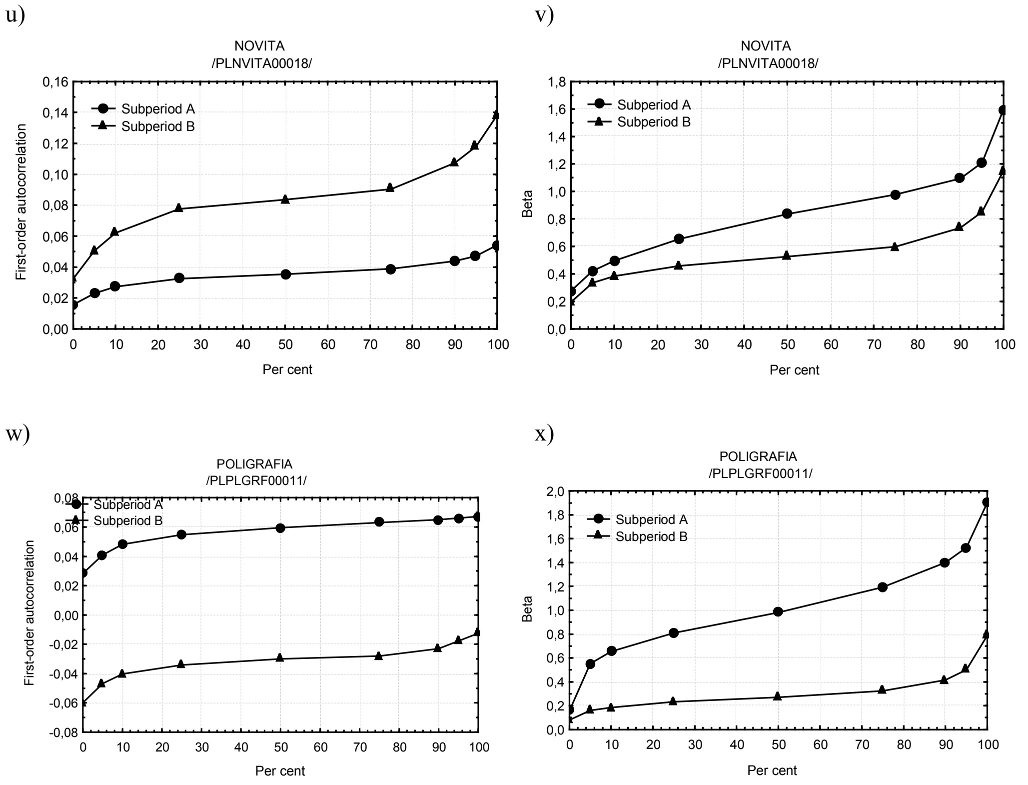

Fig. 1. The empirical cumulative distribution functions of ρ and β in two subperiods (a)–(x)

Note that the cumulative distribution functions are presented in two subperiods separately. In the case of eight firms, the cumulative distribution of ρ representing the period after the reform lies above that standing for the period prior to the reform $(\delta_{xy}^{\rho} = 0)$. This means that the autocorrelation in stock returns of these firms essentially increases over time. In the case of CPL, ELB, JLF, MSZ and NVT, this also implies that their stock returns are stronger serially correlated in the period after the reform. The increases in autocorrelation of the stock returns of the remaining three firms, namely BDX, DBC, and KTY are sufficient to change the direction of serial correlation from negative to positive over the time under study. These increases are not strong enough to make the return series more (significantly) serially autocorrelated in the subperiod after the reform, though in the case of DBC, $\inf_{\rho \in \mathbb{R}} F(\rho)$ equals 0.02.

We also find three cases where the impact of the institutional investor trade on the autocorrelation in stock returns has a reverse direction, not expected by us (in a term of (10) $\inf_{\rho\in\mathbb{R}}(F_x(\rho)-G_y(\rho))=0$). This implies that the cumulative distribution function of ρ for the period after the reform lies below that representing the period prior to the reform (see ECH, LTX, and PLG). Such a phenomenon in the behaviour of autocorrelation in stock returns in the context of pension system reform was also observed by Gebka et al. [19] but no explanation for it has been given.

At the end of 2001, seventeen pension funds were operating within the Polish pension funds sector. Taking into account the value of net assets, four of them, namely AIG, Commercial Union, ING Nationale–Nederlanden, and PZU can be unquestionably considered as sector leaders. The proportion of the leaders' net assets to the net assets of the whole sector is close to 74 per cent. On the other hand, the proportion of the net assets of each of the remaining thirteen funds in the sector does not exceed 3 per cent. The leaders are probably less prone to positive feedback trading and herding. On the contrary, after choosing their investment portfolio based on fundamental reasons, the leaders aim at preservation of the portfolio by buying when prices fall and selling when prices rise. This may induce a negative autocorrelation in stock returns. On the other hand, the remaining funds operating in the sector are probably most prone to positive feedback trading and herding. Thus, the actual level of autocorrelation that can be observed in stock returns of a given firm is the sum of both these effects (negative autocorrelation induced by the leaders, and positive autocorrelation as a result of the activities of the remaining funds operating within the sector).

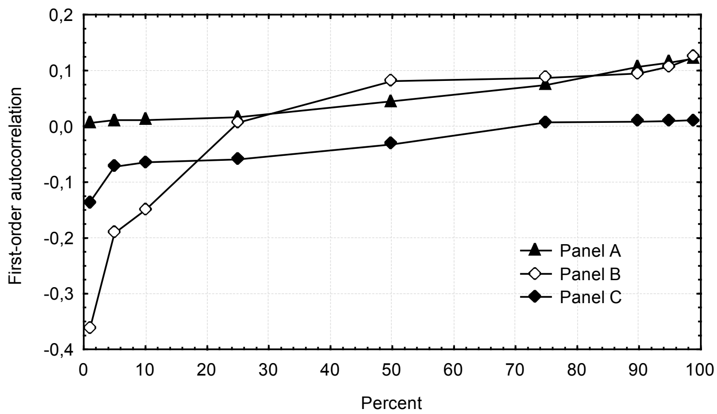

Fig. 2. The empirical cumulative distribution of ρ in the three clusters

Under the condition that the presented reasoning is true, we should observe negative autocorrelation in the case of those firms where, on the one hand, there are a few pension funds that own their shares, and where, on the other hand, the leaders' ownership is high. To test this, we divided our sample into three clusters. Panel A includes firms whose shares are traded by at least six pension funds (approximately 40% of pension funds operating within the sector). Panel B consists of firms whose shares are traded by less than six pension funds. Finally, Panel C comprises those firms included in Panel B for which the leaders' ownership amounts to at least 60 per cent. Figure 2 presents the empirical cumulative distributions for conditional first-order autocorrelation in stock returns for the three panels, separately.

Figure 2 seems to support our prediction. The cumulative distribution of ρ for Panel C is below that for Panel A. The conditional autocorrelation in stock returns of firms included in Panel C is negative, or close to zero at the higher quartiles. This may explain why the conditional autocorrelation in the stock returns of ECH, LTX, and PLG becomes negative over the period after the reform. Note that each of the abovementioned firms has one of the highest leaders' ownership in the sector (at the end of 2001 it equalled 99%, 92%, 70%, respectively). In addition, the dispersion of their shares among the pension funds operating in the sector is low for each firm.

Now, we summarize our results in the part concerning the comparison of risk level over two subperiods. One can find that the empirical cumulative distribution of β for the period after reform is below the corresponding one representing the period prior to reform in the case of all firms but two, namely CPL and ECH. On the basis of this, we can conclude that with the presence of pension funds on the domestic capital market the risk of firms that are included in the investment portfolio of funds is essentially lower.

4.2. Do insurance contribution transfers affect autocorrelation in stock returns?

In this part, we focus on stock return autocorrelation determinants. Our interest is limited to the question concerning the explanatory power of insurance contribution transfers for conditional autocorrelation in stock returns. Over the period from March 2002 to May 2004, we identified seventeen days when announcements about the transfer of insurance contributions from ZUS to OFE took place. Why might such a transfer affect the conditional autocorrelation in stock returns? Firstly, it is a widely-documented fact that some informative events like earnings or dividend announcements cause an increase in stock return volatility in surrounding days (see, e.g., [24]). The transfer of insurance contributions seems to belong to the same group of events. On the other hand, a serious body of literature is devoted to the relationship between volatility and autocorrelation and evidence to support an inverse relation between them is commonly presented (see, e.g., [27], [29], [36], [43]).

Secondly, the time-varying risk premium is another reason for the observed autocorrelation in stock returns. Säfvenblad [33] argues that the autocorrelation of stock returns is reduced if an agent's estimates of the market risk aversion become more accurate. Note that absolute volatility is likely not to have a direct effect on the welfare of pension funds managers and, consequently, on the risk tolerance (see [28]). Therefore, this aspect of risk premium variation is probably less changeable over time. Hence, market risk aversion should be more predictable after contribution transfers.

Accordingly, we expect that a transfer of insurance contributions results in the decline in the autocorrelation of stock returns. To examine this, we use a variant of event study methodology. First of all, our sample was modified. Because of limiting our interest to the period from March 2002 to May 2004, there are many more firms fulfilling the selection criteria. On the other hand, two firms were excluded from the sample, because of defects in the data. The modified sample includes twenty six firms.

As in Section 4.1, we computed estimates of the first-order autocorrelation in stock returns over the period from February 28th, 2002 to May 27th, 2004 for all sampled firms. Finally, to test the impact of contribution transfers on return autocorrelation, we employed the modified model proposed by McKenzie and Faff [29] defined as:

$$
\begin{aligned}
\rho_{1,t} &= \phi_{10} + \phi_{11} V_{1,t-1} + \phi_{12} D^{P}_{1,t-1} + \phi_{13} D^{N}_{1,t-1} + \phi_{14} D^{E}_{t} + \sum_{k=\mathrm{Mon}}^{\mathrm{Thu}} \phi_k D_k + \varepsilon_{1,t} \\
\rho_{2,t} &= \phi_{20} + \phi_{21} V_{2,t-1} + \phi_{22} D^{P}_{2,t-1} + \phi_{23} D^{N}_{2,t-1} + \phi_{24} D^{E}_{t} + \sum_{k=\mathrm{Mon}}^{\mathrm{Thu}} \phi_k D_k + \varepsilon_{2,t} \\
&\vdots \\
\rho_{n,t} &= \phi_{n0} + \phi_{n1} V_{n,t-1} + \phi_{n2} D^{P}_{n,t-1} + \phi_{n3} D^{N}_{n,t-1} + \phi_{n4} D^{E}_{t} + \sum_{k=\mathrm{Mon}}^{\mathrm{Thu}} \phi_k D_k + \varepsilon_{n,t}
\end{aligned}
\tag{12}
$$

where V_i denotes the daily trading volume for firm i measured by the number of shares traded over a day, $D_i^P\,(D_i^N)$ stands for a dummy variable that equals 1 when an above (below) average positive (negative) return occurs, D_k is a day-of-week dummy variable, and D^E is a dummy variable that equals 1 for ten trading days after the event day (an announcement about the transfer of insurance contributions) and zeros otherwise.

Note that D^E measures an average effect of transfers of insurance contributions that took place over the period from March 2002 to May 2004 on the conditional autocorrelation in stock returns of a given firm. To calculate the standard errors of the parameter estimates in (12), we use the heteroscedasticity and autocorrelation consistent (HAC) covariance matrix estimator (see [30]). Our results are summarized in Table 3.

Tabela 3

Tabela 3

Surprisingly, the relationships between autocorrelation in stock returns and trading volume are either insignificant or their direction is contradictory to some theories and empirical results (see, e.g., [29]). Note, however, that in predicting a reverse relation between autocorrelation and volume, it is assumed that non-informal trading plays a negligible role (see [33]). This assumption cannot be justified by the equities included in our sample. At least part of the transactions made by pension funds on the capital market has a non-informative nature. Hence, in the surroundings of trading days with high volume, we should not expect reductions in autocorrelation. The autocorrelation response to above-average stock price changes is significant and negative in the majority of cases. This is consistent with McKenzie and Faff [29]. The day-of-week effect is rather modestly present in the data.

As expected, in the days surrounding the day of insurance contribution transfers, decreases in the autocorrelation of stock returns can be observed. We find twenty one cases where the corresponding parameter is negative, but only for four firms, namely DBC, KTY, LTX, and RELPOL, this parameter is statistically significant.

To test jointly the hypothesis of the zero impact of pension fund presence on WSE on the autocorrelation of equities against the alternative one that at least for one company such an effect is statistically significant, the robust (HAC) Wald test was used. The test statistic is significant at 5% (see Table 3).

Note that there is no inconsistency between these findings and the evidence presented in Section 4.1. Here, we are interested in the short-term impact of transfer announcements on autocorrelation in stock returns. In Section 4.1, on the other hand, we focus our attention on the long-run relation between pension funds shares on the WSE and autocorrelation in stock returns. For example, a transfer of insurance contributions causes an increase in the conditional variance and a consequent reduction in the autocorrelation of stock returns. But with the presence of pension funds on the capital market, stock return volatility declines (see [28]). Consequently, autocorrelation in stock returns increases. Thus, our results are consistent.

5. Conclusions

The main goal of this paper was to examine whether the autocorrelation in stock returns and the risk level on the domestic capital market were affected by the increasing share of institutional investors resulting from pension system reform according to the Laotian model. A further interest is to explore the response of stock return autocorrelation to announcements about the transfer of insurance contributions.

Applying stochastic dominance criteria, we found increases in the autocorrelation of stock returns after pension reform. In many cases, a consequent gain in the serial correlation of stock returns can be observed. We also found three cases where the impact of the higher institutional investor trade on autocorrelation has a reverse (from that expected) direction. In our opinion, however, such a phenomenon can be, at least partly, attributed to the pension funds sector leaders' behaviour, which aims at the preservation of their investment portfolio by buying when prices fall and selling when prices rise. Our preliminary results seem to support this.

This study also contributes to a better understanding of the determinants of conditional autocorrelation in stock returns in the period after the reform. Our findings suggest that the reverse relation between autocorrelation and volume predicted by some theories and widely-documented in empirical contributions does not hold on a small capital market like the WSE with the presence of pension funds. On the other hand, additional determinants such as announcements about the transfer of insurance contributions to pension funds seems to play an important role in the explanation of conditional stock return autocorrelations.

As expected, we found evidence supporting risk level reduction in the response to pension fund activities on the domestic capital market. This finding supports the hypothesis about the stabilizing impact of pension reform on the domestic capital market and corresponds with the empirical evidence presented in the literature.

Comparison of unconditional estimates with their conditional counterparts shows that the commonly employed models with a dummy variable to capture the change in stock return autocorrelations seem not to be proper tools to accommodate such an effect. The M–GARCH framework, on the other hand, allows us to reflect changes in stock returns autocorrelations by changes of the model parameters.

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Wpływ inwestorów instytucjonalnych na ryzyko i autokorelację stóp zwrotu akcji w kontekście reformy systemu emerytalnego w Polsce

W artykule podjęto próbę odpowiedzi na pytanie, czy rosnący udział inwestorów instytucjonalnych, wskutek reformy systemu emerytalnego wzorowanej na modelu latynoamerykańskim, prowadzi do zasadniczych zmian w poziomie autokorelacji i ryzyka stóp zwrotu akcji, notowanym na rynku wewnętrznym. Nasze zainteresowanie w wyżej wymienionym przedmiocie zostało ograniczone do przypadku reformy systemu emerytalnego, którą przeprowadzono w Polsce w końcu lat dziewięćdziesiątych. Wyniki empiryczne uzyskano w oparciu o warunkowe współczynniki autokorelacji i warunkowe współczynniki beta, wygenerowane przy użyciu modelu M–GARCH dla pojedynczych akcji, notowanych na Giełdzie Papierów Wartościowych w Warszawie. W toku badań stwierdzono, że autokorelacja stóp zwrotu akcji ma tendencję do wzrostu, a ryzyko do spadku, w okresie po reformie systemu emerytalnego. Obecność funduszy emerytalnych na krajowym rynku akcji umożliwia zatem inwestorom budowanie lepszych prognoz przyszłego kursu akcji oraz kontrolę poziomu ryzyka. Przy użyciu odmiany analizy zdarzeń wykazano również, że poziom autokorelacji stóp zwrotu akcji staje się nieistotny w sąsiedztwie ogłoszeń o transferze składek ubezpieczeniowych na konta otwartych funduszy emerytalnych.

Słowa kluczowe: *autokorelacja stóp zwrotu akcji, ryzyko, inwestorzy instytucjonalni, reforma emerytalna*