Institutional Investors and Stock Market Efficiency: The Case of the January Anomaly[*]

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Abstract: In this paper, we investigate the effect of institutional investors on the January stock market anomaly. The Polish and Hungarian pension system reforms and the associated increase in investment activities of pension funds are used as a unique institutional characteristic to provide evidence on the impact of individual versus institutional investors on the January effect. We find robust empirical results that the increase in institutional ownership has reduced the magnitude of an anomalous January effect induced by individual investors' trading behavior.

JEL Classifications: G14, G23

Keywords: Institutional Traders, Individual Investors, January Effect, Polish and Hungarian Pension Fund Investors

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I. Introduction

Since the late 1970s, researchers have discovered several seasonal patterns in stock returns that constitute a challenge to the efficient markets hypothesis. Regularities in stock returns or stock market anomalies comprise, among many others, the January effect, the Monday seasonal, and the size effect. In this paper, we focus on the following aspect of stock market anomalies: if stock returns exhibit exploitable regularities, then smart traders are expected to take advantage of these patterns, thereby earning abnormal profits. Consequently, on stock markets with a sufficiently large number of smart traders, anomalies are supposed to disappear as the trading of this investor group arbitrages away seasonal patterns in stock returns.

Recent empirical findings suggest that institutional investors play the role of smart traders on stock markets and, therefore, may have an impact on stock market anomalies. Institutional investors can be characterized as informed traders who speed up the adjustment of stock prices to new information, thereby rendering the stock market more efficient. Institutions can obtain an informational advantage by exploiting economies of scale in information acquisition and processing. The marginal costs of gathering and processing information are lower for institutional than for individual traders. In addition, institutional investors may be better trained and have superior resources than individual investors. Moreover, for many years it has been common practice of companies to inform securities analysts in advance about company-specific news, and only recently regulatory measures have been launched (namely the SEC's Regulation FD) to prevent this habit. Hence, institutional investors' trading decisions may be stronger information-driven than those of individual investors.

Dennis and Weston (2001) support this view by providing evidence for U.S. stock exchanges that institutions are better informed than individual investors. Cohen, Gompers, and Vuolteenaho (2002) show that institutional investors push stock prices towards their fundamental values by exploiting individual traders' sentiment. Following Barber and Odean (2005), individual investors display attention-based buying behavior, whereas institutions do not exhibit this kind of non-fundamental trading pattern.

The impact of institutional trading on stock market anomalies has recently been covered by three papers. Kamara (1997) and Chan, Leung, and Wang (2004) highlight the role of institutional investors on the Monday seasonal. They present evidence for U.S. stock markets that an increase in institutional ownership decreases the magnitude of the Monday effect. Gompers and Metrick (2001) show that an increase in institutional trading is partly responsible for the disappearance of Banz' (1981) small stock premium.[1]

In this study, we focus on the impact of institutional trading on a third major anomaly, namely the January effect (Reinganum 1983, Gultekin and Gultekin 1983, and Ritter 1988). Two of the most prominent explanations for the January effect refer to the specific trading behaviors of individual and institutional investors. First, the tax-loss-selling hypothesis explains the January anomaly with tax-motivated trading of individual investors. As the end of the year approaches, individual investors sell stocks that declined in value in order to realize tax losses. After the turn of the year they re-invest in these securities, which pushes stock prices up (Ritter 1988). Second, the window-dressing hypothesis suggests that institutional investors' portfolio rebalancing activities are responsible for the January anomaly. Institutions are evaluated relative to their peers and, therefore, buy winners and sell losers in order to present respectable year-end portfolio holdings (Lakonishok et al. 1991). The findings in Sias and Starks (1997) are favorable for the

[1] Another strand of the finance literature views institutions as investors which induce non-fundamental dynamics in stock returns due to their specific trading behavior. The main arguments in this context are investment activities relying on herding, positive feedback trading, and window-dressing strategies (Lakonishok, Shleifer, and Vishny 1992; Grinblatt, Titman, and Wermers 1995; Nofsinger and Sias 1999; Badrinath and Wahal 2002; Griffin, Harris, and Topaloglu 2003).

tax-loss-selling hypothesis and show that individual traders are primarily responsible for the January anomaly.

This study highlights the impact of institutional traders on the January effect in Poland and Hungary. The history of both emerging stock markets provides a unique institutional environment to investigate the influence of individual and institutional investors on the January anomaly. In Poland, the pension system reform on May 19, 1999, separates the history of the stock market into a period of predominantly individual trading and a period of increased institutional trading. Similarly, in Hungary, private pension funds were founded in 1997 and started their financial activities in 1998. Before 1998, primarily small individual investors populated the Hungarian stock market.

The pension system reform in both countries changed the investor structure due to the enrichment of the old pay-as-you-go system with a privately managed pension funds pillar. Since 1999, pension funds are an important group of institutional investors on the Polish and Hungarian stock markets. In addition to the change of the investor structure, in both countries capital gains taxes do not exist, which rules out the tax-loss-selling hypothesis as a rationale for the January effect. Consequently, if a January effect can be detected in the data during the period before the entrance of pension fund investors in both stock markets, then it must be driven by an anomalous trading behavior of Polish and Hungarian individual investors. We exploit the increased institutional ownership in both emerging capital markets to provide evidence on the impact of individual and institutional investors' trading decisions on the January anomaly.

Relying on the institutional background of the Polish and the Hungarian stock markets, we contribute to the literature answering the following two questions. First, is there evidence in favor of a January effect during the period of individual trading? If this is the case, we can conclude that individual investors' non-fundamentally driven trading decisions led to the January anomaly.

Second, in which way did Polish and Hungarian pension fund investors contribute to the January anomaly after 1999 and 1998, respectively? In case pension funds exhibit window-dressing behavior, we expect a strengthening effect on the January anomaly. In contrast, if pension funds' trading decisions are more influenced by fundamental information, a dampening effect on unusually high stock returns in January can be expected.[2]

The remainder of the paper proceeds as follows. The next section outlines the institutional background for Poland and Hungary. Section III introduces the data set, Section IV describes the econometric methodology, and Section V contains the empirical findings. Section VI provides robustness checks, and Section VII summarizes and concludes.

II. Institutional Background

Poland

Re-established in 1991, the Polish stock market has grown rapidly during the last decade in terms of both the number of companies listed and market capitalization. In comparison to the two other European Union accession countries in the region, namely the Czech Republic and Hungary, the capitalization of the Polish stock market is significantly higher. It is comparable to that of the smaller mature European stock markets like Austria and was about 60 billion U.S. $ at the end of 2004 (Warsaw Stock Exchange 2005).

[2] It is obvious that the date of entrance of pension funds into the stock market plays an important role in the following investigation. Similarly, one branch of the literature studies the impact of the introduction of futures markets on stock return anomalies of the spot market underlying (Kamara 1997, Szakmary and Kiefer 2004). In our investigation, we can exclude an influence from the introduction of futures markets because these markets were established earlier (January 16, 1998, in Poland and March 31, 1995, in Hungary) than the appearing of pension fund investors on the stock markets took place. Nevertheless, we acknowledge that there may be other possible explanations for the decreasing January effect, like an ongoing efficiency with increased integration of transition economies (Rockinger and Urga 2001). We control for this influence and provide empirical evidence in Section V.

The change in the investor structure on the Polish stock market has its origin in the pension system reform. In 1999, the public system was enriched by a private component, represented by open-end pension funds. Participation in this component, often called the “second pillar”, is mandatory for employees below certain age. They are obliged to transfer 7.3% of their gross salary to the government-run social insurance institute called Zakład Ubezpieczeń Społecznych (ZUS), which in turn transfers the collected contributions to the pension funds. The first transfer of money from the ZUS to the pension funds took place on May 19, 1999. This date marks a change of the investor structure on the Polish stock market. In 1999, about 20% domestic institutional investors and 45% domestic individual investors traded at the Warsaw Stock Exchange. Over time the proportion of domestic institutional traders has increased, whereas the relative importance of individual investors has decreased. In 2004, approximately one-third of the investors were domestic individuals, and about one-third were national institutions. Constantly about one-third of the investors on the Polish stock market adhere to the group of foreign investors.

While before May 19, 1999, the majority of traders were small, private investors, after that date pension funds became an important group of institutional investors on the stock market in Poland. There were also some mutual funds active in the market, but they had relatively small amounts of capital under management. Moreover, the role of corporate investors, i.e., companies investing their capital surpluses, was very small. This unique institutional characteristic allows us to compare the period before May 19, 1999 – characterized by predominantly non-institutional trading – with the period after that date, when pension funds as institutional investors started to act on the stock market.

The number of pension funds in the 1999–2003 period varied between 15 and 21. The change in their number occurred mainly due to some acquisitions of smaller funds by larger ones.

It is important to note, however, that their structure as well as the structure of the assets under their management remained invariant. By the end of 2003, 17 pension funds operated in the Polish stock market with about 12 billion U.S. $ under management. In comparison, Polish insurance companies and mutual funds had only 3 and 1 billion U.S. $ of assets, respectively. In 2003, pension funds invested about 4 billion U.S. $ in stocks listed on the Warsaw Stock Exchange. Their stock holdings predominantly consist of large-capitalization stocks that are listed in the blue-chip index WIG20 and usually belong to the Top 5 in their industries.

Concerning capital gains taxation, in Poland the following regulations were in force: Until the end of 2003, capital gains from the sale of shares were tax-exempt for domestic individual investors. Since January 1, 2004 capital gains of this type have been taxed at a flat rate of 19%. For corporations, capital gains have consistently been treated as part of the company's profits and therefore been taxed at the regular corporate income tax rate. Polish pension fund investors are tax-exempt. Dividend withholding taxes varied over the period under review, the latest rate being 19% (effective 2004). However, the number of firms paying dividends is low.

Hungary

The Budapest Stock Exchange, re-established in 1990, experienced a significant increase in its capitalization, attaining about 6 billion U.S. $ in 1996, mainly due to the privatization of Hungary's bigger state-owned companies such as Mol, OTP, Gedeon Richter, and Matav. In the following years, the stock market went through a phase of continuous growth, reaching a capitalization of 30 billion U.S. $ at the end of 2004.

The introduction of a three-pillar pension system on January 1, 1998, had an influence on the Hungarian stock market because a growing share of households' savings was channeled to stock market investments through pension funds. Since 1998, individuals can choose between the mandatory public system – the first pillar – and the mandatory private system. Open-end private mandatory pension funds represent the second pillar of the Hungarian pension system. The first 38 mandatory private funds started their activities in 1998 with 134 million U.S. $ of assets under management and about 1.3 million members. The third pillar consists of voluntary pension funds, which can be both open-end and closed-end funds and also play an important role with a comparable amount of assets.[3]

The establishment of the private mandatory pension funds in 1998 was beneficial and stimulating for voluntary pension funds. The year 1998 can therefore be considered as the year when pension funds appeared as institutional investors on the Hungarian stock market. However, compared to the institutional framework in Poland, the exact date of entrance of pension funds into the Hungarian stock market is less clear-cut and hardly traceable. Whereas for Poland May 19, 1999, is known as the start date of pension funds' investment activities and well-documented as such, the investment activities of Hungarian pension funds seemed to develop gradually over the year 1998. Detailed information on this issue is practically not available. Consequently, we choose January 1, 1999, as the start date of increased institutional ownership on the Hungarian stock market to ensure that the entire post-event period is characterized by institutional trading activities. The pension funds' capital was growing during the following years and, by the end of 2004, amounted to 4 and 2.5 billion U.S. $ for the mandatory and voluntary pension funds, respectively.

The number of pension funds decreased over time, mainly due to acquisitions, and by the end of 2004, 18 private and 75 voluntary pension funds remained in the market. Contrary to other countries, where pension funds participate directly in the stock market, in Hungary an increasing

[3] The first voluntary pension funds started their activity already in 1994. However, the assets under their management were marginal at that time.

number of pension funds entrusted their assets to investment fund managers. Consequently, the impact of pension funds on stock market prices should be evaluated by means of portfolio managers' investment activities. At the end of 2004, 23 investment fund managers had under their management 4.9 billion U.S. $ of pension fund assets, 5.2 billion U.S. $ of investment fund assets, and 3.7 billion U.S. $ of contributions from other sources. Notwithstanding the assignment of pension funds' assets to portfolio managers, their investment activities have to adhere to the pension funds' investment regulations specified by law. In addition, the accumulated accounts can be invested in the longer term since contributions are not accessible before retirement.

In Hungary, capital gains realized by individual investors on the domestic or any other European Union stock exchange were considered as non-taxable interest-type income during our sample period.[4] Capital gains on transactions not qualified as stock exchange deals are, however, subject to tax – at a top tax rate of currently 25%. Only for a short period of time (2001–2002), stock market gains were also taxed at the then applicable uniform 20% capital gains tax rate. For corporate investors, capital gains are included in taxable income and taxed at standard rates. While inter-company dividend payments are tax-exempt, the dividend withholding tax rate for individual recipients varied, current rates being flat at either 25% or 35% (effective 2005). Pension funds are not subject to tax on the proceeds of the funds; these are only taxed once they are paid out to contributors.

[4] As of September 1, 2006 a 20% capital gains tax rate applies in this case.

III. Data

Poland

The data for Poland contain daily closing prices for all stocks listed on the Warsaw Stock Exchange in the period from October 3, 1994 to March 31, 2004.[5] These time series were directly provided by the Warsaw Stock Exchange. Altogether, the sample comprises 278 firms over the indicated sample period. The time series are stock-split adjusted and corrected for outliers to assure that our results are not driven or distorted by few extreme values. For this purpose, the 0.5% of highest and lowest returns observed in the data set are excluded from the investigation and, therefore, deleted from all sub-samples.

To investigate the impact of the pension funds' investment activities, we construct two sub-samples of actively institutionally traded stocks as follows. We calculate a measure of each stock's institutional coverage by dividing the aggregate pension fund holdings of that stock by the overall aggregate pension fund holdings in a particular year. This measure can be interpreted as the percentage share of a particular stock in the aggregate pension fund holdings. A stock is defined as actively institutionally traded in a given year if the measure of relative institutional holdings exceeds 1%.[6]

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[5] The selection of the start date is due to the following reasoning. Shortly after its re-opening, the Polish stock market experienced a stock price increase of 924% from May 6, 1993 to March 8, 1994, and a subsequent crash. Furthermore, it was not until October 3, 1994, that trading on the Warsaw Stock Exchange was extended from four days to five days a week. Starting our inquiry at the beginning of October 1994 ensures that the empirical findings are neither distorted by the bubble and crash periods nor affected by the change in trading frequency.

[6] We drop stocks with only marginal institutional coverage as for these stocks institutional trading behavior may not have a large impact on stock returns. The 1% cut-off point is arbitrarily chosen but proved to be an acceptable compromise for the purpose of our study. On the one hand, it allows us to eliminate those stocks which are not at all or only marginally covered by institutional investors and to come up with a limited number of stocks that are actively traded by institutions. On the other hand, the size of the resulting sub-samples is still sufficient for econometric testing.

We calculate this measure for all stocks and all years separately during the 1999–2003 period and end up with five yearly measures of relative pension fund holdings for each individual stock. A stock is included in the first sample of actively institutionally traded stocks if the pension fund holding measure of this stock exceeds the 1% level in at least three out of the five years. This amounts to 60% of the post-event period. In an alternative, less strict definition a stock has to exceed the 1% cut-off point in at least two of the five years, i.e., during 40% of the post-event period. These criteria result in the identification of 20 stocks for the stricter definition and 28 stocks for the less strict definition of institutionally traded shares. Columns 1 and 2 of Table 1 provide additional information about these stocks. Whereas Polish pension fund investors do not have a preference for stocks of a specific sector, they concentrate their investments on large firms' stocks.

[Insert Table 1 here]

Hungary

For Hungary, the data consist of daily closing prices for the stocks listed on the Budapest Stock Exchange in the period from January 3, 1994 to December 31, 2004. The time series were obtained from Thomson Financial Datastream. Altogether, the cross-section of the sample comprises 84 firms. The same trimming procedure was applied to the data set as described above for the Polish case. In contrast to Poland, we do not have reliable information regarding stock splits, dividends, and other impact factors on stock returns. This provides an additional reason for the exclusion 0.5% of the extreme stock return observations in both tails of the distribution.

To determine a sub-sample of institutionally traded stocks for the Hungarian stock market, we requested the portfolio holdings of all Hungarian pension funds. The pension funds' replies show that their stock market investment decisions closely mirror the composition of the main stock index BUX. In the sample of Hungarian stocks actively traded by institutional investors, we therefore focus on the stocks included in the BUX. Information on the BUX composition was provided by the Budapest Stock Exchange for the 1996–2004 period. Contrary to Poland, we do not use a 1% cut-off criterion because the BUX is dominated by very few stocks with high weights. Hence, a cut-off point as the one mentioned above would considerably reduce our sample in size. The number of stocks included in the institutional sample would be too small to conduct a cross-sectional investigation.

For a strict definition of institutionally traded stocks that is roughly in line with the selection criterion for Poland, we use all stocks that are included in the BUX for at least 60% of the time in the post-reform sample period 1998–2004. This definition results in the identification of 17 institutionally traded stocks. For a less strict definition, we require inclusion in the BUX for at least 40% of the same time period. The less strict definition increases the sample of institutionally traded stocks to 19. We use these two sub-samples of 17 and 19 stocks to investigate the effect of institutional trading on the Hungarian stock market. Columns 3 and 4 of Table 1 list the Hungarian companies selected together with their sector affiliation.

IV. Methodology

Groupwise Regressions

In the empirical investigation we distinguish between the impact of predominantly individual investors versus increased institutional ownership on stock returns in January. First, the hypothesis is investigated that individual investors exhibit anomalous trading behavior and cause abnormally high stock returns in January. Second, we analyze the hypothesis that institutions are informed traders relying on fundamental information and, consequently, the entrance of pension funds on the stock market dampens the anomalous January effect. If the contrary holds, the trading behavior of pension funds can be ascribed a positive contribution to higher stock returns in January relative to other months of the year, which would be in line with the window-dressing hypothesis.

The hypotheses are investigated within a panel framework and separately tested for different sub-samples of stocks from Poland and Hungary. The advantages of a panel data model over a purely time-series investigation of index data or individual shares are manifold (see, e.g., Baltagi 2001). Most importantly, unobserved individual heterogeneity can be controlled for that would otherwise have to go undetected and could generate biased results. Specifically, the following one-way error component regression model is run:

$$r_{i,t} = \beta_0 + \beta_1 JAN_t + \beta_2 JAN_t^{Post} + \beta_3 r_{i,t-1} + u_i + e_{i,t}, \tag{1}$$

where the subscript i denotes the cross-sectional and t the time-series dimension of the data set. The dependent variable is the daily stock return $r_{i,t}$, calculated as the logarithmic difference in prices $r_{i,t} = 100\ln(P_{i,t} / P_{i,t-1})$. $P_{i,t}$ denotes the individual stock price at the close of every trading day. JAN_t is a dummy variable which takes on the value of 1 in January throughout the whole sample period. The dummy variable JAN_t^{Post} is 1 only for those January observations that fall into the post-pension system reform period, i.e., beginning with January 2000 for Poland and January 1999 for Hungary. In addition, we allow for stock returns autocorrelation in the time-series dimension by including the lagged dependent variable $r_{i,t-1}$ as an additional explanatory variable.[7] u_i denotes an unobservable firm-specific random effect. $e_{i,t}$ is the remainder disturbance.

In the above specification, a positive and significant parameter β_1 provides evidence in favor of a January effect in stock returns. For the interpretation of the parameter β_2, three cases have to be distinguished. First, a negative and significant coefficient β_2 indicates a reduction of positive January stock returns (estimated by $\hat{\beta}_1$) due to the entrance of pension funds as institutional investors into the market. Second, if β_2 is positive and significant, then institutional investors' trading behavior is in line with the window-dressing hypothesis because a strengthening of the January anomaly can be observed. Third, if β_2 is statistically insignificant, institutions do not have an influence on the January anomaly. The sum $(\beta_1 + \beta_2)$ provides a measure of the magnitude of the January effect in the period of increased institutional trading.

Joint Estimation

In addition to testing the hypotheses separately for the four different sub-samples described above, we estimate the following joint model with interaction variables:

$$r_{i,t} = \beta_0 + \beta_1 (JAN_t \cdot INST_i) + \beta_2 (JAN_t^{Post} \cdot INST_i) + \beta_3 INST_i + \beta_4 POST_t + u_i + e_{i,t}, \quad (2)$$

where all previously introduced variables are defined as in equation (1).[8] In addition, the indicator variable $INST_i$ equals 1 for those companies included in the sample of institutionally traded stocks and is 0 otherwise. $POST_t$ is a dummy variable with value 1 for the period of

[7] Parameter estimates remain unbiased due to the large time-series dimension of the data. We can, therefore, rely on asymptotic properties and obtain consistent estimates (Baltagi 2001).

increased institutional trading and 0 otherwise. The interaction variables $(JAN_t \cdot INST_i)$ and $(JAN_t^{Post} \cdot INST_i)$ correspond to JAN_t and JAN_t^{Post} in regression equation (1) when it is estimated for the institutional sub-samples.

The model specified above is estimated for both sub-samples of institutionally traded stocks. We henceforth refer to the version estimated with the more strictly defined institutional dummy $INST_i$ as equation (2a) and to the less strictly defined variant as equation (2b). The coefficients β_1 and β_2 can be interpreted as described for equation (1). In addition, β_3 captures possible systematic differences between average stock returns of the institutional and the control sample, and β_4 displays aggregate factors that affected average stock returns over time in the same way for institutionally traded and non-traded shares.

V. Empirical Findings

Summary Statistics

All results are presented separately for the two sub-samples of stocks actively traded by institutional investors, a control sample of all stocks excluding the stocks identified as institutionally traded as well as the whole sample reflecting the entire Polish and Hungarian stock markets. Hence, we are able to analyze the impact of the Polish and Hungarian pension system reform on stock returns not only through time – before and after the pension funds' appearance as institutional traders on the stock market – but also in a cross-sectional dimension, i.e., among stocks more actively traded and those nearly non-traded by institutional investors.

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[8] The lagged dependent variable is dropped from the regressor list for the sake of brevity since its inclusion did not alter the empirical findings.

To gain some first insight into the seasonal patterns inherent in our data, daily average stock returns for January and for February to December are reported in Table 2. Daily mean stock returns in January are positive and higher than average stock returns between February and December for all samples. Furthermore, for both institutional sub-samples (Panels A1, B1, A2, B2) we observe higher average January stock returns during the 1994–1999 (1994–1998) period relative to the years 2000–2004 (1999–2004) for Poland (Hungary). This also refers to the whole samples (Panels D1, D2) which include all stocks listed on the respective stock exchange. Interestingly, for the Polish control sample (Panel C1) we observe an increase of average stock returns over time, whereas Hungarian stock returns (Panel C2) are slightly lower in the 1999–2004 period compared to the 1994–1998 sub-sample.

[Insert Table 2 here]

Regression Results for Poland

Table 3 displays the results from estimating regression (1) for Poland. When looking at the outcomes for the two sub-samples of actively institutionally traded stocks (Panels A and B), we find evidence in favor of a pronounced January effect in the period when the Polish stock market was dominated by individual investors. The estimated coefficients of the January effect are about 0.36. All coefficient estimates of the dummy variable JAN_t are statistically significant at the 1% level. The empirical findings in favor of a January effect are insofar interesting as during the period of predominately individual trading capital gains taxes did not exist in Poland. Hence, the tax-loss-selling hypothesis can be ruled out as a rationale for higher stock returns in January. We can therefore conclude that Polish stock returns dynamics exhibit an anomalous

January effect during the period prior to the entrance of institutional investors, which may be explained by individual investors' sentiment.[9]

[Insert Table 3 here]

Moreover, for both institutional samples the magnitude of the January effect decreases in the period after the pension fund investors' entrance into the stock market, measured by the coefficients of the post-reform dummy JAN_t^{Post}. The estimated parameter values are statistically significant and about –0.22. Thus, the significant negative parameter estimates of this institutional investors dummy lead us to reject the window-dressing hypothesis. The anomalous January effect in stock returns does not entirely disappear after the entrance of pension funds as institutional investors into the Polish stock market. However, its magnitude becomes substantially lower.

The results are robust towards the inclusion of the lagged dependent variable $r_{i,t-1}$. For both institutional samples, the coefficient of $r_{i,t-1}$ is positive and significant, which can be explained by the implications of strategic trading models (Kyle 1985, Barclay and Warner 1993). Rational informed investors spread their trades over time to conceal information. By breaking up a large order into several smaller trades, institutional investors reduce the overall price impact. Moreover, price impacts may be inversely related to market liquidity (Madhavan and Smidt 1993). This suggests that the benefits of trading over a longer horizon are greater in thin relative to liquid stock markets which, in turn, implies an increase in trade duration and a decrease in

[9] The existence of a January effect in stock returns without capital gains taxes is not new. Tinic, Baroni-Adesi, and West (1987) provide evidence for Canada and Jones, Pierce, and Wilson (1987) for the U.S. before capital gains were taxed in these countries.

order size. Moreover, the significance of a lagged dependent variable may indicate predictability in stock returns and a violation of the efficient markets hypothesis.

The estimated results for the control sample (Panel C) consisting of all stocks except for the 28 institutionally traded ones reinforce the above findings. The coefficients of the dummy variable JAN_t are positive and significant at the 10% and 19% levels. Hence, we find at least weak evidence indicating that a January effect exists in returns of non-institutionally traded stocks. In contrast to the results for the two institutional samples, the parameters for JAN_t^{Post} are statistically insignificant. For stocks not actively traded by Polish pension fund investors, the magnitude of the January effect does not decrease during the period after May 19, 1999.[10] The statistically insignificant parameters for JAN_t^{Post} in the control sample emphasize that the estimated decrease in the two institutional samples is caused by the institutions' trading behavior and not by other factors.[11] In addition, the estimated coefficient of the variable JAN_t^{Post} for the whole market is not significant either, which suggests that the January effect for the market as a whole continues to be driven by individual investors.

The empirical findings of model (2) are reported in the bottom part of Table 3. The pronounced January effect for actively institutionally traded Polish stocks is confirmed, as is the substantial decrease in the anomaly's magnitude after the entrance of pension funds into the stock market. In addition, stocks actively traded by institutions earn significantly higher returns relative

[10] Given the marginal level of significance of the JAN_t coefficients, we run separate regressions investigating whether a January effect exists in the post-event period. The values of the coefficients of the January dummy variables are slightly higher relative to the ones reported in Table 3 and are significant at the 1% level. Hence, a January effect exists in the period after May 19, 1999, in non-institutionally traded Polish stocks.

[11] The January effect in the pre-event period is substantially higher for institutionally traded stocks compared to the stocks in the control sample. A reason for this finding may be the extreme illiquidity of a subset of stocks in the control sample. As our study focuses on the evolution of January stock returns over time instead of the level of the January effect for particular stocks, we do not further explore this issue.

to the rest of the sample. The period of increased institutional trading is accompanied by higher average stock returns compared to the period before the pension system reform.

Regression Results for Hungary

The findings for Hungary in Table 4 are consistent with the ones for the Polish stock market and support the pension funds' impact on the January anomaly. The estimation results for the two sub-samples of actively institutionally traded stocks (Panels A and B) show a pronounced January effect in the period before the investment activities of Hungarian pension funds. The estimated parameters of the dummy variable JAN_t are about 0.44 and are statistically significant. In line with the results for Poland, the tax-loss-selling hypothesis as a rationale for higher January stock returns can be ruled out because capital gains are not taxed in Hungary. Moreover, the anomalous January effect decreases drastically after the entrance of pension funds into the stock market with statistically significant coefficients for JAN_t^{Post} of about –0.36. The findings are robust concerning the inclusion of the lagged dependent variable. The estimated parameters are positive and significant, supporting the implications of strategic trading models and market liquidity as well as considerations on the violation of the efficient markets hypothesis outlined above.

[Insert Table 4 here]

The empirical results of the control sample (Panel C) also indicate that a January effect exists in the period before Hungarian pension funds invested on the stock market. The estimated parameters of JAN_t are positive and significant at the 1% level. In line with the findings for

Poland, the magnitude of the January effect is smaller for non-institutionally traded shares relative to stocks actively institutionally traded. More importantly, the estimated coefficients of the dummy variable JAN_t^{Post} are not statistically significant. This finding supports the hypothesis that the estimated decrease in the two institutional samples is caused by institutions' trading behavior and not by other factors. The decrease in the magnitude of the January effect is also observed for the whole market.

VI. Robustness Check

All results presented so far were calculated for a sample where 0.5% of extreme stock returns in both tails of the distribution were dropped. As a check of robustness, we repeated the above analysis using the sample without excluding the outliers. The results for Poland are qualitatively identical. The same holds for Hungary except for the findings of the control sample. For this sub-sample, very few large return outliers seem to impact the findings and justify our outlier correction. For Hungary, the estimation of regression model (2) also supports the empirical results discussed above.[12]

Furthermore, we tackle the natural objection that the compelling evidence in favor of a decreasing January effect brought by institutional investors might merely be a reflection of some common influence or increasing trend towards efficiency. For this purpose, we widen the baseline regression model (1) to allow for a set of control variables:

$$r_{i,t} = \beta_0 + \beta_1 JAN_t + \beta_2 JAN_t^{Post} + \beta_3 r_{i,t-1} + \beta_4 CONTR_t + u_i + e_{i,t} \,. \tag{3}$$

[12] The findings are not reported but available on request.

All variables are defined as previously. In addition, $CONTR_t = \{r_{t-1}^{US}, TIME_t, TVOL_t\}$, where r_{t-1}^{US} is the one-period lagged return of the S&P 500 index, $r_{t-1}^{US} = 100\ln(P_{t-1}^{US} / P_{t-2}^{US})$, $TIME_t$ is a linear time trend, and $TVOL_t$ denotes the log of aggregate trading volume in the respective home market on day *t*, $TVOL_t = \ln(\sum_{i=1}^{n} TVOL_{i,t})$, where n denotes the number of individual stocks. Lagged U.S. stock returns are meant to capture international influences that might have gone undetected, whereas the latter two variables account for the development of the Polish and Hungarian stock markets over the decade investigated. Should the significant reduction in abnormal January returns be due to a common time trend or more active trading in general, we would expect such an evolution to be mirrored in significant coefficients for $TIME_t$ or $TVOL_t$.

The results of this robustness check are displayed in Table 5. In general, the empirical findings are fairly insensitive towards the inclusion of the three control variables. The estimated coefficients of the January dummy JAN_t are positive and statistically significant in all cases at the 1% level. More importantly, the coefficients of the JAN_t^{Post} dummy variable are in the majority of cases negative and statistically significant at least at the 5% level, while two parameters are still significant at the 14% level. This finding is robust towards the inclusion of the control variables either individually or jointly.[13] Our main hypothesis that the decrease in the anomaly's magnitude is driven by the institutions' trading activities is confirmed.

[Insert Table 5 here]

[13] We do, however, not bundle trading volume variables and the time trend together into one equation to avoid potential multicollinearity problems due to high positive correlation between these variables (0.91 for Poland and 0.53 for Hungary).

VII. Summary and Conclusion

The increase in the number of institutional investors trading on stock markets world-wide since the end of the 1980s has been associated with a rising interest from part of financial economists in institutions' impact on stock prices. One branch of literature investigates the effect of an increase in institutional ownership on the magnitude of stock market anomalies. This paper adds to the evidence available on the Monday effect (Kamara 1997; Chan, Leung, and Wang 2004) and the size effect (Gompers and Metrick 2001) by providing empirical results on the impact of institutional trading on the January effect.

Our results shed light on the causes for the anomaly and enhance the understanding of the relationship between asset prices and the investor structure of stock markets. The major difference between previous studies and ours is the unique institutional framework we exploit to investigate the role of institutional investors for the January anomaly. After the pension system reforms in Poland on May 19, 1999, and in Hungary in 1998, pension fund investors became traders on the stock market. In contrast, before these dates the majority of traders were small, private investors. Moreover, capital gains taxes did not exist in Poland and Hungary during the period of predominantly individual trading.

The institutional features of the Polish and the Hungarian stock markets enable us to investigate the role of individual and institutional investors on the magnitude of the January effect. Our empirical findings are twofold. First, we can empirically confirm that there is a significant January effect in Polish and Hungarian stock returns driven by the trading behavior of individuals. Due to the lack of capital gains taxes we cannot rely on the tax-loss-selling hypothesis as a rational explanation for the January effect. Instead, our findings suggest that higher stock returns in January during the period before the pension system reforms in both countries are the result of possibly sentiment-driven investment decisions by individual investors.

Second and more importantly, our empirical results show that the increase in institutional trading on the Polish and the Hungarian stock markets had a significant dampening effect on the magnitude of the January anomaly. Our evidence is comparable to the results found in Kamara (1997) and Chan, Leung, and Wang (2004) for the Monday effect as well as Gompers and Metrick (2001) for the size effect in the U.S. The window-dressing hypothesis is not supported. The empirical evidence indicates that trading by Polish and Hungarian pension funds to a certain extent arbitrages away seasonal patterns in stock returns and, therefore, increases the efficiency of both stock markets. The price effect of irrational trading patterns seems to be partly eliminated by rational investors.

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TABLE 1. Stocks Actively Traded by Institutional Investors.

Poland		Hungary	
Company	Sector	Company	Sector
	Institutionally traded stocks (strict definition)		
Agora	Media	Antenna	Broadcasting
BPH	Banking	Borsodchem	Chemicals
BRE	Banking	Danubius	Hotels
BSK	Banking	Demasz	Electricity Supply
Budimex	Construction	Egis	Pharmaceuticals
Computerland	IT	Fotex	Retail trade
Dębica	Chemicals	Magyar Telekom	Telecommunications
Echo	Construction	MOL	Oil/Natural Gas
Kęty	Metals	NABI	Engineering/Machinery
KGHM	Metals	OTP	Banking
Orbis	Hotels	Pannonplast	Plastics industry
PBK	Banking	Pick Szeged	Food products
Pekao	Banking	Rába	Machinery
PGF	Wholesale & Retails	Richter	Pharmaceuticals
PKN	Chemicals	Synergon	IT
Prokom	IT	TVK	Chemicals
Stomil	Chemicals	Zalakerámia	Construction
Świecie	Wood and paper		
TPSA	Telecommunications		
WBK	Banking		
	Additional institutionally traded stocks (less strict definition)		
BIG	Banking	Graboplast	Textile
ComArch	IT	Prímagáz	Gas services
Elektrim	Telecommunications		
Kredyt Bank	Banking		
Netia	Telecommunications		
Optimus	IT		
Softbank	IT		
Żywiec	Food		

Note: The table presents the stocks identified as actively traded by institutional investors and the corresponding sectors. The selection criteria are described in the text. When applying the stricter (less strict) definition, 20 (28) Polish and 17 (19) Hungarian companies are included in the sub-samples of institutionally traded stocks.

TABLE 2. Average Daily Stock Returns.

Poland			Hungary		
Sample Period	January	February – December	Sample Period	January	February – December
Panel A1: Institutional Sample I ($N = 20$)			Panel A2: Institutional Sample I ($N = 17$)		
1994 – 1999	0.3964	0.0624	1994 – 1998	0.4471	0.0368
2000 – 2004	0.1618	0.0186	1999 – 2004	0.0849	-0.0166
1994 – 2004	0.2452	0.0382	1994 – 2004	0.1993	0.0010
Panel B1: Institutional Sample II ($N = 28$)			Panel B2: Institutional Sample II ($N = 19$)		
1994 – 1999	0.3902	0.0642	1994 – 1998	0.4523	0.0369
2000 – 2004	0.1758	-0.0331	1999 – 2004	0.0662	-0.0176
1994 – 2004	0.2546	0.0110	1994 – 2004	0.1973	0.0018
Panel C1: Control Sample ($N = 250$)			Panel C2: Control Sample ($N = 65$)		
1994 – 1999	0.0004	-0.0582	1994 – 1998	0.1841	-0.0556
2000 – 2004	0.0190	-0.0361	1999 – 2004	0.1134	0.0568
1994 – 2004	0.0131	-0.0452	1994 – 2004	0.1410	0.0115
Panel D1: Whole Sample ($N = 278$)			Panel D2: Whole Sample ($N = 84$)		
1994 – 1999	0.0586	-0.0406	1994 – 1998	0.2611	-0.0287
2000 – 2004	0.0385	-0.0357	1999 – 2004	0.0976	0.0319
1994 – 2004	0.0450	-0.0378	1994 – 2004	0.1588	0.0084

Note: Mean stock returns are calculated as simple arithmetic averages of daily stock returns. The overall sample period is from October 3, 1994 to March 31, 2004, for Poland and from January 3, 1994 to December 31, 2004, for Hungary. The years 1999 and 1998 mark the dates of the Polish and the Hungarian pension system reforms, respectively. N denotes the number of stocks.

TABLE 3. Empirical Results for Poland.

Equation	$Const$	JAN_t	JAN_t^{Post}	$r_{i,t-1}$	
	Panel A: Institutional Sample I ($N = 20$)				
(1)	0.0399***	0.3512***	-0.2306**	0.0148***	
	(0.0151)	(0.0829)	(0.1015)	(0.0053)	
(1)	0.0382**	0.3582***	-0.2347**		
	(0.0152)	(0.0830)	(0.1017)		
	Panel B: Institutional Sample II ($N = 28$)				
(1)	0.0114	0.3730***	-0.2089**	0.0225***	
	(0.0133)	(0.0714)	(0.0882)	(0.0044)	
(1)	0.0104	0.3787***	-0.2134**		
	(0.0144)	(0.0717)	(0.0886)		
	Panel C: Control Sample ($N = 250$)				
(1)	-0.0535***	0.0559*	0.0065	-0.0336***	
	(0.0059)	(0.0341)	(0.0408)	(0.0017)	
(1)	-0.0452***	0.0456♣	0.0186		
	(0.0060)	(0.0344)	(0.0411)		
	Panel D: Whole Sample ($N = 278$)				
(1)	-0.0445***	0.1069***	-0.0317	-0.0277***	
	(0.0055)	(0.0310)	(0.0372)	(0.0016)	
(1)	-0.0378***	0.0963***	-0.0200		
	(0.0055)	(0.0313)	(0.0375)		
Equation	$Const$	$(JAN_t \cdot INST_i)$	$(JAN_t^{Post} \cdot INST_i)$	$INST_i$	$POST_t$
(2a)	-0.0700***	0.3872***	-0.2796**	0.0791***	0.0450***
	(0.0094)	(0.0991)	(0.1215)	(0.0189)	(0.0112)
(2b)	-0.0718***	0.4086***	-0.2608**	0.0534***	0.0464***
	(0.0095)	(0.0814)	(0.1008)	(0.0161)	(0.0113)

Note: The estimated models are (1) $r_{i,t} = \beta_0 + \beta_1 JAN_t + \beta_2 JAN_t^{Post} + \beta_3 r_{i,t-1} + u_i + e_{i,t}$ and (2) $r_{i,t} = \beta_0 + \beta_1 (JAN_t \cdot INST_i) + \beta_2 (JAN_t^{Post} \cdot INST_i) + \beta_3 INST_i + \beta_4 POST_t + u_i + e_{i,t}$, where stock returns are calculated as $r_{i,t} = 100\ln(P_{i,t} / P_{i,t-1})$. JAN_t (JAN_t^{Post}) denotes a dummy variable which takes on the value 1 in January throughout the whole sample period (only in the post-pension system reform period). $INST_i$ is a dummy variable indicating a stock's affiliation to the stricter (less strict) sub-sample of institutionally traded shares for equation 2a (2b). $POST_t$ is a dummy with value 1 for the period of increased institutional trading. *, **, *** denote statistical significance at the 10%, 5%, and 1% levels, respectively, and ♣ at the 19% level.

TABLE 4. Empirical Results for Hungary.

Equation	$Const$	JAN_t	JAN_t^{Post}	$r_{i,t-1}$	
	Panel A: Institutional Sample I ($N = 17$)				
(1)	0.0003 (0.0138)	0.4357*** (0.0832)	-0.3541*** (0.0991)	0.0331*** (0.0051)	
(1)	-0.0021 (0.0201)	0.4405*** (0.0837)	-0.3544*** (0.0998)		
	Panel B: Institutional Sample II ($N = 19$)				
(1)	0.0012 (0.0132)	0.4449*** (0.0770)	-0.3822*** (0.0932)	0.0309*** (0.0048)	
(1)	-0.0006 (0.0181)	0.4468*** (0.0774)	-0.3807*** (0.0939)		
	Panel C: Control Sample ($N = 65$)				
(1)	0.0094 (0.0111)	0.1701*** (0.0605)	-0.0575 (0.0760)	-0.0500*** (0.0033)	
(1)	0.0115 (0.0112)	0.1726*** (0.0608)	-0.0707 (0.0765)		
	Panel D: Whole Sample ($N = 84$)				
(1)	0.0067 (0.0087)	0.2555*** (0.0481)	-0.1598*** (0.0598)	-0.0309*** (0.0027)	
(1)	0.0062 (0.0098)	0.2518*** (0.0484)	-0.1622*** (0.0603)		
Equation	$Const$	$(JAN_t \cdot INST_i)$	$(JAN_t^{Post} \cdot INST_i)$	$INST_i$	$POST_t$
(2a)	-0.0139 (0.0155)	0.4801*** (0.0994)	-0.4135*** (0.1190)	-0.0246 (0.0223)	0.0563*** (0.0179)
(2b)	-0.0156 (0.0158)	0.4849*** (0.0909)	-0.4397*** (0.1108)	-0.0218 (0.0216)	0.0581*** (0.0179)

Note: The estimated models are (1) $r_{i,t} = \beta_0 + \beta_1 JAN_t + \beta_2 JAN_t^{Post} + \beta_3 r_{i,t-1} + u_i + e_{i,t}$ and (2) $r_{i,t} = \beta_0 + \beta_1 (JAN_t \cdot INST_i) + \beta_2 (JAN_t^{Post} \cdot INST_i) + \beta_3 INST_i + \beta_4 POST_t + u_i + e_{i,t}$, where stock returns are calculated as $r_{i,t} = 100 \ln(P_{i,t} / P_{i,t-1})$. JAN_t (JAN_t^{Post}) denotes a dummy variable which takes on the value 1 in January throughout the whole sample period (only in the post-pension system reform period). $INST_i$ is a dummy variable indicating a stock's affiliation to the stricter (less strict) sub-sample of institutionally traded shares for equation 2a (2b). $POST_t$ is a dummy with value 1 during the period of increased institutional trading. *** denotes statistical significance at the 1% level.

TABLE 5. Robustness Check.

Const	JAN_t	JAN_t^{Post}	$r_{i,t-1}$	r_{t-1}^{US}	$TIME_t$	$TVOL_t$
Panel A: Poland – Institutional Sample I ($N = 20$)						
0.0299**	0.3404***	-0.1562♣	-0.0066	0.3471***		
(0.0152)	(0.0833)	(0.1023)	(0.0053)	(0.0118)		
0.0608	0.3411***	-0.2140**	0.0148***		0.0000	
(0.0376)	(0.0845)	(0.1051)	(0.0053)		(0.0000)	
-0.4639***	0.4024***	-0.3420***	0.0143***			0.0634***
(0.1031)	(0.0835)	(0.1039)	(0.0053)			(0.0128)
0.0265	0.3420***	-0.1589♣	-0.0066	0.3471***	0.0000	
(0.0376)	(0.0850)	(0.1059)	(0.0053)	(0.0118)	(0.0000)	
-0.4752***	0.3925***	-0.2687***	-0.0072	0.3472***		0.0636***
(0.1035)	(0.0839)	(0.1048)	(0.0053)	(0.0118)		(0.0129)
Panel B: Hungary – Institutional Sample I ($N = 17$)						
-0.0084	0.4344***	-0.2880***	0.0179***	0.3405***		
(0.0139)	(0.0850)	(0.1018)	(0.0052)	(0.0115)		
0.0662**	0.3969***	-0.3024***	0.0329***		0.0000**	
(0.0334)	(0.0851)	(0.1019)	(0.0051)		(0.0000)	
-0.2838***	0.5334***	-0.4594***	0.0335***			0.0403***
(0.0760)	(0.0921)	(0.1088)	(0.0053)			(0.0106)
0.0302	0.4115***	-0.2577**	0.0179***	0.3403***	0.0000	
(0.0337)	(0.0870)	(0.1047)	(0.0052)	(0.0115)	(0.0000)	
-0.2454***	0.5021***	-0.3613***	0.0182***	0.3487***		0.0336***
(0.0769)	(0.0920)	(0.1097)	(0.0054)	(0.0118)		(0.0107)

Note: The estimated equations are variants with different regressors of the model $r_{i,t} = \beta_0 + \beta_1 JAN_t + \beta_2 JAN_t^{Post} + \beta_3 r_{i,t-1} + \beta_4 CONTR_t + u_i + e_{i,t}$, where stock returns are calculated as $r_{i,t} = 100\ln(P_{i,t} / P_{i,t-1})$. JAN_t (JAN_t^{Post}) denotes a dummy variable which takes on the value 1 in January throughout the whole sample period (in the post-pension system reform period). $CONTR_t = \{r_{t-1}^{US}, TIME_t, TVOL_t\}$ describes a set of control variables, where r_{t-1}^{US} denotes the one-period lagged return of the S&P 500 index, $TIME_t$ is a linear time trend, and $TVOL_t$ denotes the log of aggregate trading volume in the respective home market on day *t*, $TVOL_t = \ln(\sum_{i=1}^{n} TVOL_{i,t})$. *, **, *** denote statistical significance at the 10%, 5%, and 1% levels, respectively, and ♣ at the 14% level.