Pension Funds and IPO Pricing. Evidence from a Quasi-Experiment

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Abstract

We exploit a quasi-experiment arising from the government-forced changes to the assets under management and investment policy of the Polish pension funds. We test whether this new regulation and its resultant demand shock on the investors' side, leads to changes in the IPO pricing and the subsequent stock's performance. We report material and a statistically significant decrease in the IPO proceeds (IPO size) in the post-treatment period equal to over 107 million PLN (34 million USD). We find no empirical evidence that the treatment had a significant effect on the first-day IPO underpricing or on the long-term underperformance. We conclude that the demand shock resulting from the pension system reform that primarily aimed at solving fiscal problems effectively eliminated the so-called 'pension premium' of higher IPO valuations. Thus, it indirectly impaired companies' power of raising money in the public stock market. Furthermore, we report a decrease in the average first-day IPO returns among big issuers that is consistent with the book building literature.

Keywords: IPO pricing, first-day returns, long-term underperformance, quasi-experiment, event study, pension funds

JEL codes: G11, G14

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1. Introduction

Effective February 2014, the Polish government introduced new legislation that regulated the allocation of assets under management (hereafter, 'AUM') and the investment policy of the Polish pension funds. The most significant changes encompassed a lower contribution to private pension funds, the gradual transfer of funds related to persons approaching the retirement age, and a one-off transfer of all government-guaranteed securities from private pension funds to the state-governed administrator. The main reason for these changes was to aid the unsustainable pay-as-you-go state-governed part of the pension system in Poland. Moreover, the new regulation restrained pension funds from high-risk equity investments, but at the same time loosened their investment policy by eliminating the internal benchmarking mechanism as a measure of fund's performance. In consequence, we observe lower net cash flows to pension funds after the 2014 reform.

A long line of literature documents the effects of pension funds' investing process on capital markets. One stream of research finds that the investment activity of pension funds is beneficial to equity markets by promoting liquidity, boosting demand for capital-market instruments and innovation, and contributing to the development of capital markets on the whole (Davis, 1998; Walker & Lefort, 2000). Conversely, Singh (1996) and Roldos (2004) suggest that institutional investors might exert a negative influence on the capital markets in emerging economies. In particular, private pension funds' investments contribute to asset price distortions, bubbles, and concentration of risks (Roldos, 2004). In the context of the Polish Pension Reform of 1999, Zalewska (2006) shows that the benefits of pension funds' investment in the home market are only short-term and revert in the long run. Despite the intense debate on the capital market effects of pension fund investments and pension system reforms, none of the papers specifically looks at how they affect the new listings and their aftermarket performance in equity markets.

In this paper, we investigate whether there is a link between the pension fund reform of 2014 and the pricing of initial public offerings (hereafter, 'IPO'). We exploit a quasi-experiment arising from the exogenous treatment of the Polish pension funds to test whether a demand shock on the investors' side leads to changes in the IPO market. We argue that lower net cash flows to pension funds and eliminating the internal benchmark mechanism as a measure of fund's performance (the leading cause of herding among the Polish pension funds in the pre-reform time) decreased funds' aggregate demand for equities, and specifically for new listings. Our research idea is motivated by the practitioners' claim that this regulatory change could have had "a negative impact on the Warsaw Stock Exchange, where the funds have been crucial buyers of shares and have been responsible for the so-called ‘pension premium’ of higher valuations” (Cienski, 2014). Specifically, we intend to test whether a demand shock on the investors’ side, caused by the regulatory change, led to changes in the size of the IPO proceeds and the level of the IPO underpricing.

We study IPOs in the Warsaw Stock Exchange Main Market (hereafter, ‘WSE’), which has operated since 1991. Importantly, we also look at the NewConnect[1] (hereafter, ‘NC’)—an alternative stock market created in 2007 to aid smaller entities that could not meet the requirements (either capital, procedural, or disclosure) of the WSE. As of June 30, 2017, both platforms included 891 entities with a combined market capitalisation of 1,326 billion PLN, which was equivalent to 358 billion USD[2] (GPW, 2017; NewConnect, 2017). Noteworthy, years of reforms, restructuring, and privatisation of the formerly state-owned companies made it possible to create the largest and the most important stock bourse in Central and Eastern Europe, characterised by the largest IPOs, improved liquidity, transparency and investor protection (Jackowicz, Kowalewski, Kozłowski, & Roszkowska, 2017). In 2016 and 2017, the Warsaw Stock Exchange reported the third-highest number of IPOs among all European exchanges (PwC, 2018). Further, the Polish stock exchange was recently classified by the FTSE Russell as a ‘developed market’ in its FTSE Global Equity Index Series—an improvement from the ‘advanced emerging’ category (FTSE, 2017). Despite its regional importance, research on the IPO effects in the Polish stock market is still scarce.

We use a unique dataset of IPOs listed in the two equity markets for the years 2005-2017. We combine multiple databases and hand-collect missing data to arrive at the most comprehensive sample ever used for the Polish stock market[3]. The abundance of observations in our dataset yields an opportunity to investigate IPO underpricing effects in the Polish stock market. We find strong support for both the IPO underpricing and long-term underperformance effects observed previously in the U.S. and other advanced markets.

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[1] The NewConnect market is less formal than the main market. It takes less time for a firm to go public there due to simpler and faster conditions for introducing the company to trading, there are lighter information obligations (e.g. the lack of audited semi-annual reports) and low debut costs. However, the market microstructure is the same in the main and alternative market (e.g. regarding how trade is done, the role of market makers, etc.)

[2] Exchange rates are as per the official Central Bank of Poland statistics; retrieved on February 23, 2019, from https://www.nbp.pl/homen.aspx?navid=archen&c=/ascx/TabArchAen.ascx&n=17a125en for June 30, 2017.

[3] Jelic and Briston (2003) study the average market-adjusted first-day returns and report the average underpricing of 27.37% from 1991 to 1998, while Jewartowski and Lizińska (2012) report the average underpricing of 13.95% for the subsequent ten years of 1998–2008. However, both studies are rather confined because they encompass short periods, a limited number of stocks and they focus only on the main equity market of the Warsaw Stock Exchange.

We evaluate the impact of the government-forced changes to the AUM and investment policy of the Polish pension funds on the IPO pricing and aftermarket performance using various approaches and robustness checks. First, we compare the mean outcomes in the pre- and post-treatment periods, separately for the WSE and NC markets. The two-sample *t*-test provides first insights. We find that the mean IPO size decreased by 142 million PLN (over 45 million USD[4]) in the WSE market post-treatment. At the same time, there was no statistically significant effect in the control (NC) sample. Then, we find no statistically significant changes in either the mean IPO first-day returns or in the mean long-term underperformance. Robustness tests based on bootstrapping confirm our results.

Further, we look for the causal effect between the regulatory treatment of the Polish pension funds and changes in the IPO pricing using the difference-in-differences research design. Such an identification is possible because our data consists of two subsamples: WSE and NC IPOs. Pension funds in Poland only invest in the former market, and not in the latter due to regulatory constraints. Therefore, we can compare the treated (WSE) and nontreated (NC) IPO outcomes, and ascribe any differences to the effect of the regulation. We first report a notable and statistically significant change in the IPO size: average IPO proceeds decrease by almost 107 million PLN after the regulatory treatment of the Polish pension funds. We attribute this finding to the institutional investors' decreased aggregate demand and more (practical) independence in making investment decisions (release from the internal benchmark as a performance measure). We argue that there is a new equilibrium established in the IPO market after the reform where listing companies can raise relatively less money on average (no more 'pension premium'). We discard a possibility of big companies cancelling IPOs for the reason of not being able to raise as much money as before the reform because we see neither a decline in the post-treatment number of new listings nor an increase in the number of IPO withdrawals. To check for the robustness of our results, we employ a placebo test where we alter the date of the event. We run difference-in-differences regressions as in the original specification using the simulated 'event time' as each day from the +/- 3-year period around the original event, and we analyse the *p*-values of the difference-in-differences estimator. Our results are robust to placebo testing. Only around the original date of the regulatory change do the *p*-values fall close to the zero-level. We conclude that a demand shock resulting from the pension fund reform of 2014

[4] Exchange rates are as per the official Central Bank of Poland statistics; retrieved on February 23, 2019, from https://www.nbp.pl/homen.aspx?navid=archen&c=/ascx/TabArchAen.ascx&n=14a022en, for February 3, 2014.

effectively decreased the benefits of newly listed companies regarding the amount of money raised in the public stock market.

Additionally, our cross-sectional tests based on IPO size reveal that the effects of regulation are concentrated among larger issuers. This observation is consistent with a blue-chip bias, whereby Polish pension funds prioritize larger firms in their equity portfolios (Zalewska, 2006). The effect of the reform on the IPO size is statistically significant in the larger two IPO size quantiles, but not in the smaller two quantiles. Moreover, we report a statistically significant decrease in the IPO underpricing. This decrease is because underpricing is a means of compensating institutional investors for the critical role they play in supporting IPOs. First, the book building literature posits that IPO underpricing is used by investment bankers to elicit revelation of information about the market value of an IPO during the book building process (e.g., Benveniste & Spindt, 1989). Second, underwriters use IPO underpricing to further compensate institutional investors for supporting IPOs with weaker post-issue demand. Hereto, institutional investors support IPOs by agreeing to hold their allocations for a longer period (Chemmanur, Hu, & Huang, 2010). Against this backdrop, we expect to see a decline in the level of underpricing in response to the above-described demand shock on the institutional investor side. Indeed, for the largest IPO quartile, where Polish pension funds primarily invest, we see a statistically significant decline in the first-day returns following the reform. We conclude that the reform not only affected the amount of pension funds' AUM, but it also had a detrimental effect on how much they make on average when investing in IPOs as the first-day returns of large IPOs decreased.

The contribution of this paper is three-fold. First, our paper provides empirical evidence as to whether the IPO underpricing and long-term underperformance effects are present in each of the equity markets in Poland. Second, by exploiting a quasi-experiment, we shed new light on the indirect impacts of the 2014 pension funds reform. We show that the decreasing net cash flows to pension funds AUM and abandoning the internal benchmark as a measure of funds' performance (a shock on the demand side) result in lower IPO proceeds. At the same time, we report a significant decrease in the first-day returns for the subsample of the largest issuers, consistent with book building theories of IPO underpricing. These findings are of interest to both market participants but also legislators if they consider similar reforms in the future. Lastly, by providing insights about the changed levels of IPO first-day returns and IPO long-term underperformance in the reform aftermath, we contribute to the literature on the factors that drive IPO underpricing effects and to the literature on asset pricing reaction to demand shocks on the investor side. We argue that the insights and contributions provided in this study are of importance to the international community, given the scale of IPO activity in the Warsaw Stock Exchange. Our tests have been run on a sample of 706 IPOs for the years 2008-2017, while in the U.S.—the biggest stock market in the world—there have been 990 IPOs in a respective period (Ritter, 2019).

The rest of the paper is organised as follows. Section 2 depicts the regulatory change to the pension fund AUM and investment policy. In Section 3, we develop research hypotheses. Section 4 explains the research methodology. Section 5 details the data and provides descriptive statistics. Section 6 presents the main results and robustness tests. Section 7 concludes.

2. Pension Fund Reform of 2014: The Setting

The contemporary version of the Polish pension system follows the World Bank's multi-pillar design since its establishment in 1999 (Holzmann, 2000). Voronkova and Bohl (2005) provide a comprehensive description of the Polish pension fund sector, which consists of the two obligatory pillars: (1) a pay-as-you-go system managed by a state-owned entity (the Social Insurance Institution), (2) open pension funds run by private managing companies, and one voluntary pillar: (3) privately funded pension security schemes.

In time, however, this pension system demonstrated severe shortcomings. Most importantly, the system turned out to be unsustainable, which manifested itself in the growing fiscal problem of not enough funds allocated to the first pillar for covering state pension obligations. System unsustainability was the main reason for the 2014 policy change that resulted in several alterations to the pension system. The key changes included: (i) a one-off asset transfer of T-bonds (written-off immediately), government-guaranteed bonds, municipal bonds, and cash in the amount of 51.5% of their total assets from the second pillar (pension funds AUM) to the first pillar; (ii) lower contribution rate to the second pillar: 2.92 per cent instead of 3.5 per cent of income, and (iii) gradual transfer of assets of persons ten years before the retirement age from the second pillar to the first pillar (IMF, 2014; Journal of Laws 2013, Item 1717, 2013). The legal act was signed in December 2013, while legal changes were enforced on February 3, 2014. The one-off transfer took effect on the same date. From the pensioners' standpoint, the value of their contributions only changed hands from privately-managed to a state-managed institution. The transfer of debt securities did not have any effect on pension funds' equity investments either. However, future contributions and AUM were to decrease, which could affect equity investments. Moreover, effective February 2014, pension funds were restricted from investing in government-backed securities, while a minimum of 75% of their assets under management was to be invested in equities (with a ban to invest in the NC market).

Additionally, until July 2014 pension funds were required to guarantee a minimum rate of return on their investments, which was set every quarter based on the average performance of all pension funds (an internal benchmark mechanism). In case of not meeting the minimum return, a fund covered the shortfall from its special reserves, own assets, joint pension funds' reserves, or from the government's funds in the last resort. The reason for imposing such investment policy restrictions on pension funds was to protect informed and uninformed contributors (KNUiFE, 2000). This mechanism gave rise to herding. Voronkova and Bohl (2005) and Kominek (2012) document strong herding behaviour by Polish pension funds, much more intense than in other developed markets, and particularly evident regarding investment in large-capitalization stocks. In July 2014, the internal benchmarking mechanism stopped serving as a performance measure for the Polish pension funds. Thus, the herding should no longer exist and should not determine pension funds' investments.

3. Background and Hypotheses Development

Our hypotheses are first motivated by macro-finance literature. Roldos (2004) underlines the importance of loosening the regulatory framework for pension funds. Due to herding practices among these entities, changes in investing sentiment usually result in widespread changes. Therefore, even a marginal policy change may have a substantial effect on the capital market as a whole. In the Polish context, policymakers prioritise political and social outcomes of pension fund reforms (Góra, 2014). The caveat of such behaviour is that it potentially comes at the expense of capital markets prosperity. Up to date, the literature did not consider how these policy changes may affect IPOs and thus the private sector.

3.1. Pension Fund Reform and IPO Pricing

A few theoretical developments and stylized facts motivate us to study the relationship between the described regulatory change and IPO pricing. Primarily, we are intrigued by untested practitioners' claims that after the regulatory change of 2014, pension funds started pushing for a reduction in the price of offers (Kucharczyk, 2016) and that post-reform there should no longer be a 'pension premium' of higher valuations in the WSE (Cienski, 2014).

We argue that there are two specific elements of the regulatory change that drive such expectations. First, as a result of the lower contribution rate to the second pillar and a gradual transfer of pension assets of persons ten years before the retirement age to the first pillar, there would be a decrease in the net cash inflow to the open pension funds. Coval and Stafford (2007) provide evidence that changes in capital flows to institutional investors' AUM create price pressures in equity markets. Funds that experience significant outflows have no choice but to sell some of their holdings to cover redemptions ('asset fire sales'). When many funds sell at the same time, one should expect a negative price pressure in the market. We anticipate this mechanism to work also in our setting: reduced inflows to all pension funds decrease their aggregate demand for new shares, which negatively affects the amount of money raised by listing companies.[5]

Our expectation is conditional on pension funds having economic power enough to affect equity prices. As reported in the stock exchange statistics, the second pillar—open pension funds—has been a significant equity investor in the Warsaw Stock Exchange. Between end-2013 (around the regulatory change) and mid-2017 (end of our sample), pension funds' equity investments have been growing from 100 billion PLN (about 33 billion USD) to over 148 billion PLN (nearly 40 billion USD)[6]. This corresponds to almost 12% share of the joint WSE and NC market capitalization end-2013, and 11.2% end-2017 (GPW, 2018). More importantly, Polish pension funds are a repeated 'price maker' during the book building process, determining the ultimate offering price. Zalewska (2006) provides evidence that pension funds take significant stakes in the new listings[7]. Hereto, there is an apparent blue-chip bias: the average free float taken over by the pension funds in the highest IPO-size quartile (large cap companies) equals to 36.4%, successively decreasing to 26% in the lowest IPO size quartile (small cap companies), which is still a significant stake being taken up by only a few institutions. Additionally, as confirmed in our consultations with pension fund managers and investment bankers,[8] pension

[5] In addition, Chemmanur et al. (2010) argue that pension funds are contributing to both the demand at the time of the issue and the post-issue demand.

[6] Exchange rates are as per the official Central Bank of Poland statistics; retrieved on February 23, 2019, from https://www.nbp.pl/homen.aspx?navid=archen&c=/ascx/TabArchAen.ascx&n=13a251en for December 31, 2013, and from https://www.nbp.pl/homen.aspx?navid=archen&c=/ascx/TabArchAen.ascx&n=17a125en for June 30, 2017.

[7] Zalewska (2006) also gives examples of IPOs, where Polish pension funds jointly took over 60% of the newly issued shares, e.g., 12 open pension funds absorbed 63.8% of the free float of GTC's IPO in 2004.

[8] In the research process, we have consulted our hypotheses and results with pension fund managers in Poland. Consultations have been made in form of phone calls conducted between October 2018 and January 2019, and again between November and December 2019. The interviewees were asked unstructured questions about the practice of their investing, about the effect of the reform on their trading decisions, and about their perception of the reform and its outcome on the equity market in Poland. There has been no specific questionnaire followed in those interviews. The consulted fund managers represent four of the seven biggest pension funds in Poland (AUM-wise), and all wished to remain anonymous. We refer to practitioners' insights a few more times throughout the paper.

funds are the first subject to be approached during the roadshow when an investment bank assesses the overall demand for the shares of the new issue.

Second, the departure from the minimum rate of return as a measure of performance reduces incentives for herding. Blake and Timmermann (2002) suggest that when the evaluation benchmark for fund managers is a weighted average (as it used to be in Poland before 2014), the safest investment strategy is to follow market leaders. Lakonishok, Shleifer, and Vishny (1992) document herding behaviour in a large sample of U.S. pension funds. Similar findings have been reported for the behaviour of the Polish pension funds before 2014 (Kominek, 2012; Voronkova & Bohl, 2005). Zalewska (2006) provides indirect evidence for herding while taking up shares of new listees, especially in cases of big IPOs. She reports that out of 15 open pension funds active in Poland, the average number of pension funds investing in the new shares of the companies from the highest IPO size quartile was 11.4, and it decreased successively towards 2.6 in the lowest quartile. This decline is in line with fund managers' accounts of their behaviour in the primary market: before the regulatory change, they tried to mimic each other's equity engagements in new, especially more prominent listees to stay close to the internal benchmark with their portfolio performance. In practitioner's views, by declaring demand for shares of a listing company primarily because other funds declared their demand, a fund contributes to an artificial 'pension premium.'

Marrying these two immediate outcomes of the 2014 pension fund reform: the decrease in net cash flows to pension funds with eliminating the main driver for herding behaviour, we argue that the regulation, aimed initially at covering the state deficit in the pension system, was also a negative demand shock for equities in the primary stock market. Our premise is that as a result of this negative shock on the investors' side, pension funds bid less during the book building process[9] (no more 'pension premium') and, as a result, companies raise less money at an IPO. We test our prediction with the following hypothesis:

Hypothesis 1: *Government-forced changes to the AUM and investment policy of the Polish pension decreased the average IPO size.*

[9] Noteworthy, the IPO process in the Warsaw Stock Exchange follows typical characteristics and stages of IPOs in other markets, for instance in the U.S. as explained by Ellis et al. (2000) or Brau and Fawcett (2006). The main practical difference is that in the U.S., most IPOs are done within a firm-commitment agreement with the underwriter (the underwriter purchases the whole offer and resells the shares to investors, while the issuing company is guaranteed that a particular sum of money will be raised), while in Poland, most transactions happen on the best efforts agreement (the underwriter does not guarantee the amount raised for the issuing company, rather it's role is limited to sells the securities on behalf of the issuing company).

We state the alternative hypothesis that government-forced changes to the AUM and investment policy of the Polish pension funds did not affect the average IPO size.[10] Failing to reject the null hypothesis will imply that lower net cash flows to pension funds and eliminating formal incentives for herding behaviour reduced the amount of money listing companies raise on average at an IPO. The counterfactual reasoning, in this case, is that the event was not meaningful for the money raised at an IPO. Thus, IPO size did not change after the event. Accordingly, the time trend of the average yearly IPO size in the WSE post-reform should be mirroring the time trend of an average annual IPO size in the NC market. We do not envision a positive scenario, under which the average IPO size increases in the reform aftermath, as this would be both surprising and counterintuitive. There is no empirical evidence suggesting that after a demand shock on the investor side, equities were priced higher on any stock exchange.

3.2. Pension Fund Reform and IPO Underpricing Effects

In a seminal study, Stoll and Curley (1970) observe a 'remarkable price appreciation' between the IPO offering price and the first-day closing market price. Soon after that, Logue (1973) and Ibbotson and Jaffe (1975) provide empirical evidence for this phenomenon that is currently known as the 'IPO underpricing effect'—a situation when a stock is priced below the value anticipated by the market, leading to abnormal returns in the first day of trading. Recent evidence suggests that at an aggregated level, stocks listed in the U.S. exchanges experienced an average first-day price appreciation of 7.2%, 14.8%, 64.6%, and 14% during the periods of 1980–1989, 1990–1998, 1999–2000 and 2001–2016, respectively (Ritter, 2017). Similar IPO underpricing is observed in other capital markets.

In this paper, our investigation extends onto examining whether the exogenous treatment of pension funds that originated a demand shock on the investors' side, caused any changes in the level of the first-day IPO returns in the stock market of Poland. Our analysis is motivated by the fact that pension funds are materially invested in IPOs, especially in the most significant IPOs. If the 2014 pension reform was also a negative demand shock for equities in the secondary stock market, then the lower activity of institutional investors could lead to higher first-day returns due to sentiment (Ljungqvist, Nanda, & Singh, 2006) and information asymmetries

[10] Because the two changes coincided in time and IPO observations are only observed at discrete intervals, it is impossible to test for the two driving forces separately.

(Field & Lowry, 2009)[11]. Alternatively, any decline in the first-day returns would be consistent with the book building literature (Aggarwal, Prabhala, & Puri, 2002; Benveniste & Spindt, 1989; Chemmanur et al., 2010; Rock, 1986) that posits that IPO underpricing is a mean of compensating institutional investors for truthfully disclosing value-relevant information during the book building process. Specifically, institutional investors have access to private information on the fair IPO value, and investment bankers use underpricing as a tool of eliciting that information during the pre-selling period. Accordingly, there is a positive association between the level of underpricing and the level of institutional investors' participation in the book building process. In our setting, this story can be recast through the following mechanism: in the pre-reform equilibrium, *ceteris paribus*, there was a certain average level of IPO underpricing that was 'used' by underwriters to compensate institutional investors. In the immediate post-reform period, however, there would be less underpricing as a larger fraction of the market is taken up by retail investors, who are less likely to possess and disclose value-relevant information, and who do not need to be compensated with extra underpricing. Given these two potential alternative scenarios, the ultimate effect of the negative demand shock on the investors' side on the IPO underpricing is *ex-ante* unclear. Therefore, we state the following null hypothesis pertaining to the aftermarket first-day IPO returns as:

Hypothesis 2: *Government-forced changes to the AUM and investment policy of the Polish pension funds did not affect the degree of the first-day IPO underpricing.*

The null hypothesis is consistent with the reasoning that pension funds are not affecting the first-day closing price in the treatment period to counterbalance the consequences of the reform because, in general, they invest for a long-term horizon and, according to practitioners, they do not rebalance their portfolios with newly listed shares shortly after an IPO (especially during the first day of trade). If the null hypothesis is rejected, the sign of the estimated relationship could be either positive or negative due to offsetting mechanisms affecting the first-day returns. Accordingly, the alternative hypothesis states that the government-forced changes to the AUM and investment policy of the Polish pension funds affected the first-day IPO returns.

After the first-day abnormal price appreciation, there is a tendency for the subsequent downward-oriented stock price adjustment (Miller, 1977). Ritter (1991) names this

[11] Ljungqvist et al. (2006) argue that the smaller the involvement of institutional investors, the more sentiment-driven investment decisions, and thus higher underpricing. Field and Lowry (2009) add that institutional investors have an information advantage, therefore due to their investments, stocks tend to underperform less in the long-term.

phenomenon an ‘IPO underperformance effect’ and defines it as the stock’s underperformance in the years following an IPO. Miller (1977) explains this tendency through the combination of heterogeneous expectations (pessimistic vs optimistic investors) and short-sales constraints. In this paper, we investigate whether the regulatory change of 2014 affected the level of the IPO underperformance. The decrease in the activity of institutional investors after the treatment could exacerbate the IPO underperformance by leaving more room for sentiment driven investment and information asymmetries (Field & Lowry, 2009). Thereupon, we state the following null hypothesis pertaining to the effect of the reform on the IPO underperformance within a short-term (30-day) time window following an IPO:

Hypothesis 3: *Government-forced changes to the AUM and investment policy of the Polish pension funds did not affect the degree of short-term IPO underperformance.*

The alternative hypothesis states that the government-forced changes to the AUM and investment policy of the Polish pension funds did affect the short-term underperformance. We do not specify the counterfactual more precisely because the short-term window also includes the first-day returns.

Notably, the positive association between institutional trading of IPOs in the aftermarket and the long-term IPO performance decays over time (Chemmanur et al., 2010). This is because after the company goes public, a substantial amount of new information is provided to the market, which reduces the information advantage of institutional investors. Accordingly, we state the following null hypothesis testing for the effect of the reform on the IPO underperformance within a short-term (30-day) time window following an IPO:

Hypothesis 4: *Government-forced changes to the AUM and investment policy of the Polish pension funds did not affect the degree of long-term IPO underperformance.*

The counterfactual reasoning could be that the 2014 reform did affect the degree of the long-term IPO underpricing. Still, we do not find previous empirical evidence in support of the alternative hypothesis.

Lastly, we explore a theoretical possibility of an alternative explanation of our results: given the new market conditions, where pension funds have less money to invest, firms’ managers evaluate the potential IPO proceeds against the costs of raising capital, and because they anticipate lower demand for their company’s shares (fear of not raising enough funds), they choose not to do an IPO at all (Lowry, 2003). Firms that still decide to go public under new conditions would need to 'compensate' investors for the unfavourable financial environment with higher underpricing. For example, Dorsman and Gaunopoulos (2013) show that the European sovereign crisis resulted in Dutch IPOs 'leaving more money on the table'.

Our identification relies on the specific design of the Polish stock market. As previously described, the government-forced changes to pension funds' AUM and investment policy serve as an exogenous shock to IPO pricing and stocks' subsequent performance. Importantly, this is only true for one of our markets under research: the WSE, because pension funds do not invest in the NC market. Before 2014, Polish pension funds' investments in the NC companies were close to zero, while after 2014, their investment in the alternative market was forbidden. Thus, the shock originating from the regulatory change had the potential to affect IPO pricing and stock's subsequent performance only in the WSE subsample. By exploiting a quasi-experiment, we can test for the changes in IPO pricing and the degree of IPO underpricing effects as a consequence of the 2014 regulatory change.

4. Methodology

To test for IPO underpricing and long-term underperformance, we calculate the first-day IPO returns and post-IPO 30-day and 360-day returns. For n-th stock, we measure the first-day return ($RET_{n,1}$) as the difference between the closing price of the first day of trading, $P_{n,1}$, and the IPO offer price, $P_{n,0}$, divided by that closing price, as in Equation (1):

$$RET_{n,1} = \frac{P_{n,1} - P_{n,0}}{P_{n,0}} . \qquad (1)$$

We calculate short- and long-term returns as compounded stock price return over a k-day period following the equity issuance (i.e. for 30 and 360 calendar days, respectively). As in Barber and Lyon (1997), to calculate the non-market-adjusted returns for the period starting at the IPO date (t=1) and ending after k days, we use the buy-and-hold (holding period) rate of return ($HPR_{n,k}$), as in Equation (2):

$$HPR_{n,k} = \prod_{t=1}^{k}\left(1 + RET_{n,t}\right) - 1, \qquad (2)$$

where $RET_{n,t}$ is the return on day t for stock n, calculated using adjusted prices. In the average long-term cumulated return calculations, an individual stock forms a single portfolio, and both size- and equal-weights are used. If any data is missing for a company, we exclude it from calculations for all research periods starting with the first missing data point.

To evaluate performance, we adjust returns using the Capital Asset Pricing Model (CAPM) market factor. The CAPM-assumed market portfolio does not occur in reality, and its betas, as systematic risk measures, are not sufficient to explain the cross-section of expected returns (Roll, 1977). Additionally, the model's empirical robustness is questionable (see Roszkowska and Langer (2016b) for evidence of its poor performance in the Polish stock market and Fama and French (2015) for the similar conclusions for the global markets). Nonetheless, a single-factor model allows for convenient estimation of risk-adjusted returns in an everyday setting. Thus, we calculate the abnormal return for an *n*-th stock for each day k ($AR_{n,k}$) as in Equation (3):

$$AR_{n,k} = RET_{n,k} - \left[Rf + \beta_j \times MKT_n\right], \tag{3}$$

where Rf is the risk-free rate, β_j is the respective industry *j*-th beta, and MKT_n is the single market risk factor premium for the *n*-th stock. Then, following Ritter (1991), we compute the cumulative abnormal returns (CAR). For the *n*-th stock and the period from the IPO date (*t*) to the *k*th day after an IPO, we calculate abnormal return as in Equation (4):

$$CAR_{n,k} = \sum_{t=1}^{k} AR_t. \tag{4}$$

The one-month Warsaw Interbank Offer Rate (WIBOR) proxies the risk-free rate, given the lack of data on short-term treasury bills in Poland. Roszkowska and Langer (2016a) test for the merits of using WIBOR as the risk-free rate and find no significant differences between WIBOR and the periodically available time series of short-term treasury bills. For the beta coefficients, we use average levered industry beta estimates for the period of 2011–2016 from the Damodaran database (Damodaran, 2018). We use this database for Europe as it includes Polish stocks. We proxy the market risk factor premium using the adjusted return on the Warsaw Stock Exchange Index (WIG). This total return index includes dividends and pre-emptive rights of all companies listed in the primary market, excluding foreign companies and investment funds.

Finally, we study the impact of an exogenous shock in the form of government-forced changes to pension funds' AUM and investment policy on the IPO pricing. We begin by testing for the differences in mean IPO returns (first-day returns, short- and long-term performance) and mean IPO size between the pre- and post-treatment subsamples. Hereto, we use a two-sample *t*-test.

We split the sample by the IPO date, using February 3, 2014, as a cut-off date. We separately test for the differences in means in the WSE and NC subsamples. We do not log-transform IPO size because we expect the results to be driven by large IPOs, where pension funds heavily invest as confirmed by our interviewing of pension fund managers, and as previously evidenced by Zalewska (2006).

We test for the robustness of these results using a nonparametric bootstrap methodology. In doing so, we address (and try to eliminate) the possibility that our results are due to chance, i.e. smaller IPOs randomly occur in the post-treatment period, in the treated sample. To generate possible outcomes due to chance, we randomly sample IPOs with replacement and calculate the difference between mean IPO size in the pre-treatment period and mean IPO size in the post-treatment period To test for significance of the differences we use the *p*-value calculated based on the bootstrap sample. Specifically, we calculate *p*-value as in Equation (5):

$$p_{boot} = \frac{1}{N}\sum_{i=1}^{N} I\left(d_i > \bar{d}\right) \tag{5}$$

where N is the number of bootstrap iterations, d_i is the pre-post difference in the iteration i, and $\bar{d}$ is the observed pre-post difference.

We use February 3, 2014, as our cut-off date, and the symmetric 3-year period around the cut-off when sampling the IPOs (short-time window). The number of IPOs in every bootstrap sample, in both pre- and post-intervention periods, matches the total number of IPOs in our sample, in both pre- and post-intervention periods. We report the density based on 10,000 iterations, and compare the differences estimated in our sample to the distribution of differences based on the sampling results. We repeat the same procedure separately for the treated (WSE) and control (NC) group.

Next, we implement the difference-in-differences regression methodology. This methodology compares the effect of an event on groups affected thereby (treated) with those that are unaffected (control). The treated group corresponds to the WSE new listings, in which pension funds (directly affected by the legal change) are materially invested, and the control group corresponds to the NC new listings, in which pension funds do not invest at all. Thus, any change to the pension funds' AUM or investment policy could only have a potential effect on the IPO pricing in the WSE, and not in the NC market. Noteworthy, the difference-in-differences strategy eliminates the bias that comes from other unobserved changes that could have affected the treated group.

Accordingly, using IPO-level data, we study the response to the exogenous shock regulatory change using the cross-sectional regression specification as in Equation (6):

$$Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \gamma_t + \delta_j + \varepsilon_i. \qquad (6)$$

The dependent variable Y_i is either the first-day return, short- or long-term underperformance, or the IPO Size for the i-th stock. WSE_i is a binary indicator equal to 1 for the IPO listings in the treated WSE market, and 0 otherwise. $Post_i$ is a binary indicator equal to 1 for the IPO listings after the pension fund reform came into force, and 0 otherwise. Thus, the interaction term $WSE_i \times Post_i$ captures the difference-in-differences effect. γ_t are the time fixed effects to control for unobservable differences between the IPOs of firms listed in different years, t, and δ_j are the industry fixed effects to control for time-invariant differences between the IPOs of firms from different industries, j.[12] ε_i is the error term. We omit the $Post_i$ variable because the pre-post firm-invariant differences are subsumed by year fixed effects.[13] We also omit firm-specific characteristics because, unlike fixed effects, they can be endogenous to our treatment assignment, and hence bias our coefficient of interest (e.g., Roberts & Whited, 2013). Following Bertrand et al. (2004), we cluster standard errors to avoid overstating the significance of our difference-in-differences estimates. We cluster our errors at industry-year levels to ensure a sufficient number of clusters.

To ensure the robustness of our results, we augment our baseline regression specification with additional control variables. First, we include the natural logarithm of total assets to control for the firm size (e.g., Barth, Landsman, & Taylor, 2017; Liu & Ritter, 2010; Loughran & Ritter, 2004). Second, we include the return on assets (ROA) to control for the firm profitability, or more broadly for the firm quality (e.g., Barth et al., 2017; Chang, Chiang, Qian, & Ritter, 2016). Finally, we also include VC backing status as an additional control (e.g., Barth et al., 2017; Liu & Ritter, 2010; Loughran & Ritter, 2004). Accordingly, we use the regression specification as per Equation (7):

[12] In an additional robustness check, we further add interactions between industry fixed effects and the post indicator to directly control for industry-specific differences between pre- and post-treatment periods that could affect our results. Our main coefficients remain statistically significant. Table C10 in Appendix C reports our results using this alternative specification.

[13] In the alternative specification of the difference-in-differences model we also consider a simpler model that excludes year fixed effects but includes a post-treatment indicator. Table C6 and Table C11 in Appendix C report our results using this alternative model without and with additional controls, respectively.

$$Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \beta_3 log\ (Assets)_i + \beta_4 ROA_i + \beta_5 I(VC)_i + \gamma_t + \delta_j + \varepsilon_i. \tag{7}$$

log $(Assets)$ is the total assets in the last fiscal year before the IPO. This variable controls for the confounding effects attributable to the firm size. ROA is the return on assets in the last fiscal year before the IPO, using the net income in the numerator and the end-of-period value of total assets in the denominator. I(VC) is an indicator of whether an IPO was Venture Capital-backed.

We perform our tests using the symmetric three-year window around the cut-off date of February 4, 2014. Thus, we have the pre-treatment period of February 4, 2011, to February 3, 2014, and the post-treatment period of February 4, 2014, to February 3, 2017. We complement our analysis with an implementation strategy that uses long-time window based on the full data, where the pre-treatment period is January 1, 2008, to February 3, 2014 (six years one month), and the post-treatment period is February 4, 2014, to June 30, 2017 (three years five months). Whereas the narrower time window might help establish the relationship on the treatment group, longer series alleviates concerns related to short-term autocorrelation (see Hausman & Rapson, 2018). We trim the time-series in January 2008, due to data availability: the NC market launched in 2007 and pension funds' detailed holdings are available from 2008.[14]

To reinforce our result, we run a robustness check in the form of a placebo test. Specifically, we re-estimate our regression based on the different cut-off dates from our sample. We again use a symmetric window of three years, conditional on data availability. In the final analysis, we reassess the significance of the coefficient on the interaction terms depending on the cut-off date.

5. Data and Descriptive Statistics

5.1. Data

We obtain data on IPOs in the Polish stock market from the following providers: *GPW* (gpw.pl), *Stooq* (stooq.pl), *Bossa* (bossa.pl), and *Bloomberg*. *GPW* data include a complete list of IPOs and a summary of the most relevant IPO data (i.e. IPO date, size, and offer price). We retrieve the adjusted daily stock prices from the *Stooq* database. Adjustments account for stock splits, dividends and other distributions, rights offerings, and denominations. We retrieve non-adjusted

[14] Because 2017 is incomplete in our dataset, we also test for the robustness of our results in the sample that excludes the year 2017. We confirm that our results are robust to exclusion of the year 2017. In fact, when we use a long-period estimation window, the coefficient on the $WSE \times Post$ variable is higher in magnitude and statistical significance.

prices from the *Bossa* database. *Bloomberg* is used to obtaining any missing time series data, particularly the daily adjusted and non-adjusted time series of the stock prices for delisted, acquired, and liquidated companies. The missing IPO data is hand-collected from companies' prospectuses. We obtain financial statements data and company identifiers from Notoria. Finally, we obtain additional IPO information from Dealogic. Appendix A provides a detailed description of our data.

Our data spans from January 2008 to June 2017. We trim the time-series, excluding observations prior to January 1, 2008, due to data availability: the NC market was only launched in 2007 and pension funds' detailed holdings are available from 2008. Additionally, Lyn and Zychowicz (2003) provide evidence that the small number of companies, ambiguous macroeconomic conditions, lack of proper structure and legislation, and a shortage of skilled brokers led to numerous pricing abnormalities in the WSE during the 1990s and early 2000s. This could bias the patterns of IPO pricing for the time series before 2008. Another reason for limiting the sample is the anomaly reported by Lyn and Zychowicz (2003): suspiciously high IPO underpricing (54.45%) for a sample of 103 IPOs in Poland from 1991 to 1998 with respectively high standard deviation (82.23%).

Our sample includes 706 IPOs: 237 listings in the WSE and 469 listings in the NC.[15] It is important to note that the sample diminishes in size when we analyse post-IPO short- and long-term underperformance. Some companies lacked the time series of returns after an IPO, either because they were no longer active (e.g., delisting within the first year after an IPO) because the IPO was too recent or because there was a delay in data availability (further described in Appendix A). Our dataset distinctly outnumbers previous studies: 195 observations in Jewartowski and Lizińska (2012), 211 in Jelic and Briston (2003), and 103 in Lyn and Zychowicz (2003). Noteworthy, Jewartowski and Lizińska (2012) explicitly exclude delisted companies in their research of IPO underpricing effects. We argue that such treatment introduces severe survivorship bias into research on IPO underpricing. In our dataset, there are 178 companies with a market status other than 'currently active', including 58 delisted companies in the WSE and 120 delisted companies in the NC. We include IPO data and subsequent return of the delisted companies in all of our analyses to mitigate the survivorship

[15] In Figures 1-3, as well as in the expanded time-series and cross-sectional statistics for the first-day returns reported in Tables B2 and B3 in Appendix B respectively, we include more data for descriptive purposes. This data goes back to January 2005 and predates the inception of the NC market. The total sample therein includes 874 IPOs: 387 listings in the WSE and 469 listings in the NC. This data starts in 2005 because (1) the stock exchange does not provide a comprehensive IPO list before 2005 and (2) prospectuses for earlier IPOs are often not available.

bias. Otherwise, the results would be biased towards the more successful IPOs. In rare cases of re-listings, we use the data of the chronologically first IPO. We are also the only scholars thus far to include observations from the NC market for analysing IPO underpricing effects. We point to the advantages of our dataset not to claim the superiority of our results over previous research. Instead, we strive to establish a viable reference point to communicate the reliability of our dataset.

We winsorize the data at 0.5% and 99.5% of observations to limit the influence of outliers. We examine winsorization choices on a case-by-case basis and test the sensitivity of our results (see Leone, Minutti-Meza, & Wasley, 2019; Wilson, 1997). We do not winsorize returns due to negative skewness, but we do winsorize IPO size to reduce the impact of the extreme outliers. Table C5 in Appendix C reports our main results corresponding to three different winsorization choices: 0% (no winsorization), 0.5% (winsorizing at 0.5% and 99.5%) and 1% (winsorizing at 1% and 99%). The sensitivity analysis confirms that our results are robust to the alternative winsorization choices.

5.2. IPO characteristics and IPO distribution by sector and in time

Table 1 presents descriptive statistics for the main variables, separately for the sample of companies listing in the WSE and the NC. For the period of January 1, 2005, to June 30, 2017, we report an average IPO size of 177 million PLN in the WSE and 2.6 million PLN in the NC. Table C1 in Appendix C reports the expanded summary statistics with additional percentiles of the empirical distributions of our main variables.

Table 2 Panel A summarises the sample distribution across economic sectors, and Panel B—in time, separately for listings in the WSE and the NC. We classify firms into industries using Global Industry Standard Classification (‘GICS’).[16] We observe no material differences in sector distribution between the two markets under research. In Panel B, we see an evident drop in the number of IPOs in both markets in the aftermath of the global financial crisis, particularly in the year 2009. The highest total number of IPOs (184) appears in 2011. During that period, the European Union targeted significant subsidies towards small and medium enterprises in Poland. Part of this support program was intended to finance equity issuances in the NC and covered up to half of the total IPO cost for the participating companies. Because the subsidy

[16] Bhojraj et al. (2003) compare alternative industry classifications (including SIC, NAICS, or Fama and French classifications) and argue that GICS outperforms other taxonomies in identifying comparable companies.

program ended in 2011, most firms went public by the end of the year, explaining an increased number of IPOs in that year and a subsequent decline.[17]

[Insert Table 1. here]

5.3. IPO underpricing and long-term underperformance

Figure 1 presents unadjusted first-day returns of Polish IPOs over time. The visual inspection of the plot reveals that most of the sample IPOs yield positive first-day returns. Further, the figure indicates that except for the year 2007, the NC first-day returns are consistently higher than the WSE returns. NC sample is also associated with the higher dispersion of the IPO returns. Moreover, there is a noticeable drop in the number of IPOs and an increase in the dispersion of returns during the global financial crisis, particularly in the year 2009. The highest number of IPOs is observed in the post-crisis years 2010 and 2011, which is consistent with the descriptive evidence.

[Insert Figure 1. here]

Table 1 presents empirical estimates for the IPO underpricing and long-term underperformance of equities listed in the two Polish stock markets. The equally-weighted average of the unadjusted first-day return for WSE (NC) is approximately 5.8% (40.0%), indicating a high degree of underpricing, especially in the NC market. The dispersion of NC first-day returns varies from -90.6% to as much as 2,900%. The corresponding minimum and maximum values for WSE first-day returns are -74.07% and 242.0%, respectively. This evidence aligns with Miller's (1977) theory of divergence of opinions among investors and reaffirms the inherent characteristics of the NC, a stock market dedicated to innovative, technology-related and less mature companies (Jackowicz et al., 2017).

The use of IPO size weights does not materially affect the interpretation of the previous results: the average first-day returns for WSE and NC equal to 5.5% and 25.4%, respectively. Given that equally-weighted returns are higher than size-weighted returns and that NC returns are higher than WSE returns, our evidence is consistent with the evidence of smaller IPOs

[17] We also look at the industry composition over time. To the degree that industry membership proxies for significant differences in firm characteristics, this is informative about whether important time trends in the IPO characteristics drive our results. Upon investigation, we do not see significant changes in the industry composition of our IPO sample. We report the industry composition over time in Appendix C Figure C1. One exception is that we see the decline in the number of IPOs in the consumer discretionary sector. Nevertheless, our regression results are robust to the exclusion of the IPOs from that sector. We report an additional robustness check using the alternative sample that excludes IPOs from that sector in Appendix C Table C7.

generating relatively higher IPO returns. This result is also in accordance with the idea of a size premium.

[Insert Table 2. here]

In line with empirical evidence from the U.S. market (e.g., Ritter, 1991), as well as from the foreign developed and emerging markets (e.g., Boulton, Smart, & Zutter, 2011) cumulative market-adjusted returns for newly floated stocks drift downwards after the initial first-day leap. In the Polish stock markets, average unadjusted returns also drop in the initial period following an IPO, which is consistent with Miller (1977). Comparing the average returns for 30 days following an IPO, we observe a downward trend in stocks' performance: an average WSE issue (in an equally weighted portfolio) delivers a 1.8% return, whereas an average NC issue delivers a 43.1% return. Therefore, compared to first-day returns of 5.8% and 40.0% for the WSE and the NC, respectively, an average stock accrues an incremental return of -4.0 percentage point and 3.1 percentage point respectively over the period from day 2 to day 30 after an IPO. Once we consider the IPO-size-weighted returns, stocks in both markets yield lower returns than over the first day of listing only: 4.9% and 21.6% in the WSE and NC, respectively.

Figures 2 and 3 show the average short- and long-term performance of the WSE and NC IPOs. Figure 2 presents the unadjusted, IPO-size-weighted (Panel A) and equally weighted (Panel B) returns accumulated until the 360th day after an IPO. The NC IPO-size-weighted stock returns drift downwards immediately after the upsurge on the first day. Their WSE counterparts do not follow that exact path: the WSE IPO-size-weighted stock returns rise gradually throughout the period, with a slight indication that they are sloping downwards around day 330. The NC stock returns experience numerous erratic movements, with frequent changes from high to low. Such variability in returns may be related to the specific characteristics (i.e. immaturity) of stocks listed in this market, especially when compared to the WSE companies.

[Insert Figure 2. here]

Next, we look at the market-adjusted returns. Figure 3 plots IPO-size-weighted (Panel A) and equally-weighted (Panel B) WSE and NC returns accumulated until the 360th day after an IPO. Despite numerous fluctuations, the returns indisputably descend over time, confirming previous findings. Further, NC portfolio returns are considerably more volatile than WSE portfolio returns (Panel A). This observation is consistent with differences in the risk-return profiles of WSE and NC stocks, and is further confirmed by lower standard deviations of the WSE returns compared to their NC counterparts (see Table 2). The cumulative market-unadjusted equally-weighted returns indicate flat to slightly increasing trends (Panel B).

[Insert Figure 3. here]

We also conduct a cross-section and time-series analysis of the first-day returns. Partitioning the sample by years and industries enables a more thorough understanding of the IPO underpricing effect, which does not seem to be driven by the year or industry under study. Time-series analysis indicates some cyclicality in the first-day returns. Yet, the relationship is not completely clear as the relative consequences of the global financial crisis were rather mild for the Polish financial sector. Cross-section analysis of the first-day returns reveals considerable cross-sectional variation in the IPO returns. Ultimately, the estimates for the decomposed returns reinforce earlier evidence of the IPO underpricing effect (Stoll & Curley, 1970). We provide the detailed results and more thorough analysis thereof in Appendix B, Tables B1 and B2 for the time-series and cross-section of first-day returns, respectively.

6. Empirical Results

Figure 4 adds to the practical motivation to our identification strategy and predictions. We compare the cumulated percentage changes in the aggregated pension funds' AUM, in the equity part of the aggregated pension funds' AUM and the Warsaw Stock Exchange WIG index. The cumulated percentage changes are scaled relative to the yearend of 2013.

[Insert Figure 4. here]

In 2014, we observe a material drop in the AUM due to the one-off transfer of funds from the second to the first pillar of the Polish pension system. However, this transfer encompassed only government-guaranteed debt securities and cash. A more critical observation is what happened with pension funds' equity investments. Before 2014, pension funds' equity holdings have been decreasing less and increasing more than the WIG index. It is due to the internal benchmark mechanism that funds have been following each other's equity investments rather than pursuing active returns, which Voronkova and Bohl (2005) classify as herding and provide evidence thereof in the Warsaw Stock Exchange. Consequently, pension funds' performance has been mirroring the WIG index returns, while additional growth in the equity holdings relative to changes in the WIG index is to be attributed to the positive and high net cash inflows of contributions (which funds have been investing in equities listed in the WSE). Interestingly, pension funds' equity holdings have not been surpassing the WIG index after February 2014. Instead, changes in pension funds' equity holdings have been strictly following the WIG changes. This is suggestive of not much additional net capital inflows, and non-increasing equity investments. If that is true that changes in equity holdings are only due to changes in assets prices rather than to incremental investments, this observation might also indicate the continuation of herding behaviour among funds when it comes to their equity engagement. We are inclined to think that the regulatory change of June 2014 related to departing from the internal benchmark mechanism as a measure of funds' performance did not have a considerable impact on their investment strategies. Providing empirical evidence thereto is, however, out of the scope of the current paper. Conversely, the above observations help us develop a prior that once there is no more incentive for herding, there will be less pressure for higher IPO valuations ('pension premium' in IPO proceeds) because not all pension funds would invest in IPOs that other pension find attractive.

Based on the casual observation of Figure 4, from the year 2014 onwards we expect lower aggregated equity investments in newly listed stocks, which should negatively impact IPO proceeds (IPO size). Since it is difficult to make a precise prediction upon graphical illustration only, the effect of the government-forced changes on the IPO pricing and performance becomes the focus of further analysis.

6.1. Regulation-Induced Changes in the IPO Market

We begin by testing for the differences in means between the pre- and post-treatment subsamples. We split the sample by the IPO date, using February 3, 2014 as the cut-off date. We run a two-sample *t*-test on the IPO first-day returns, on the return proxies for the short- and long-term underperformance, and on the IPO size. We perform a separate test for WSE and NC new listings.

[Insert Table 3. here]

The regulatory changes affected both the size of pension funds AUM and the way they invest capital (investment policy). Therefore, we are expecting this treatment to impact IPO pricing and the subsequent performance of the newly listed stocks. Table 3 posits that there is an economically and statistically significant difference in the mean IPO size between the pre-and post-treatment subsamples in the WSE. This finding does not depend on whether we use short- or long-time window implementation strategy. In both cases, the pre- and post-treatment mean IPO size in the WSE sample decreases materially, i.e. by 142 million PLN when using a short-time window, and by 128 million PLN when using long-time window. Both estimates are significant at 5% level. In terms of economic magnitude, the post-intervention decline corresponds to around 70% of the pre-intervention mean IPO size. At the same time, the difference in the pre- and post-treatments mean IPO size in the NC market lacks statistical and economic significance. These results provide first empirical evidence of lower IPO valuations associated with the regulatory change in the stock market with a strong presence of pension funds. The effect is otherwise unobservable (in the NC subsample).

We fail to uphold similar conclusions for the differences in the mean first-day IPO returns and the mean short- and long-term underperformance (cumulated IPO returns) in either of the stock markets. We conclude that the regulatory change did not affect the degree of IPO underpricing. The null result is consistent with the signalling model (Ritter & Welch, 2002). If the underpricing is understood as a tool to signal quality to investors, there is no reason to believe that the demand shock on the investors' side would affect underwriters' decision to underprice to a different degree. This result is also in line with practitioners' views. According to the fund managers consulted in our research process, pension funds actively invest at IPOs but do not tend to rebalance portfolios in the short-term. The latter insight means that they hardly ever trade stock on the first day after an IPO nor change their positions during the 360 days after an IPO.

Incidentally, we note that the negative demand shock on the investors' side could also affect overall market performance, especially the IPO risk-adjusted long-term performance. In Table C3 in Appendix C, we test for the difference between pre- and post-treatment sample average unadjusted returns in the WSE that are independent of the market performance. This difference shows as insignificant, regardless of the time horizon used. Therefore, we discard the possibility that the results for market-adjusted returns reported in Table 3 are driven by a combination of the offsetting effects on IPO returns and overall market performance. In Table C4 in Appendix C, we also test whether our conclusion is driven by inflated IPO returns in the years of the global financial crisis and we report that excluding years 2008 and 2009 from our sample does not have any effect on the degree of IPO underpricing. Table C8 in Appendix C reports our regression results in the subsample that excludes years 2008 and 2010 of the global financial crisis.

Figure 5 shows the visualization of our result. The figure plots mean annual IPO size for the WSE (Panel A) and NC (Panel B) markets in 3 years preceding and following the event (parallel trend lines). As it transpires, there was a significant drop in the mean IPO size following the regulatory change in early 2014. In fact, in the years 2014, 2015 and 2016, the WSE IPO valuations were consistently at their lowest levels in terms of the average offering size. This is not the case in the NC sample, what validates our expectations. This finding also eliminates potential concerns that the changes in IPO pricing may result from unobservable events or other changes in macroeconomic and market conditions.

[Insert Figure 5. here]

Next, we show the results of the nonparametric bootstrap test. We aim to formally assess the statistical significance of our finding without making any distributional assumptions. Figure 6 shows the distribution of bootstrapped pre-post differences in the mean IPO size for the WSE (Panel A) and NC (Panel B) market. In bootstrapping tests, we sample IPOs with replacement and calculate the difference between the mean IPO size in the pre-treatment period and the mean IPO size in the post-treatment period. We use February 3, 2014 as our cut-off date and symmetric 3-year period around the cut-off in the IPO sampling procedure (the short-time window). We plot the density based on 10,000 sampling iterations. The value of sample the difference in our original sample is illustrated with the solid vertical line. Figure 6, Panel A (WSE) shows the original sample difference of -142.04 million PLN. We argue that this is not driven by chance (is statistically significant). In fact, our observed original sample difference of -142.04 million PLN corresponds to the t-statistic of 2.07, as only 1.9% observations lie below the observed mean difference. In contrast, Figure 6, Panel B (NC) reports the original sample difference of -0.31 million PLN, which is no different from the mean difference resulting from the bootstrapping. In this case, the t-statistic equals to 0.34, as 37.0% of observations lie below the observed mean difference. We repeat the same analysis using full data from January 1, 2008 to June 30, 2017 (the long-time window). This implementation strategy yields similar results. Therefore, based on the random sampling testing procedure, we confirm a statistically and economically significant decline in the IPO size in the sample where pension funds are heavily invested (WSE), but not in the sample that is independent of changes in pension funds investment policies (NC). We also run a permutations test to arrive at similar conclusions. We show the distribution of permutated pre-post differences in the mean IPO size for the WSE and NC market in Figure C2 in Appendix C. We additionally provide a summary of all nonparametric test estimates in Table C2 in Appendix C.

[Insert Figure 6. here]

6.2. Event Study

We implement the difference-in-differences estimator to test for the impact of the 2014 regulatory change to the pension funds' AUM and investment policy on the IPO pricing and stock's subsequent performance. We estimate the average response of the IPO size, market-adjusted first-day returns, market-adjusted short- (30-days) and long-term (360-days) underperformance to the regulatory change on the demand side. As per difference-in-differences specification, we use the WSE IPOs as the treated group and the NC IPOs as the control group. Table 4 shows the estimates of the average dependent variables' response to the exogenous shock based on the model as per Equation 6. To control for unobservable industry-level and year heterogeneity, we use industry and year fixed effects in all regressions.

[Insert Table 4. here]

In Table 4, we report material and statistically significant change in the IPO size (winsorized at 0.5% and 99.5%). Our results indicate that the average IPO size in the treated group went down by 128 million PLN (almost 41 million USD[18]) after the regulatory change. The *p*-value of the estimate is 5.9%. The estimate increases to 152 million PLN if we use unwinsorized data (Table C5 in Appendix C). These results are reported for the short-time window implementation strategy. When we implement a long-time window strategy, our findings prevail. The point estimate for the effect of the regulatory change on IPO size is -107 million PLN, significant at the 5% level (*p*-value of 4.9%). Furthermore, irrespective of the implementation strategy, the statistical significance of the *WSE* estimate at the 1% level validates the difference in cross-market characteristics: IPOs in the WSE are on average 76 times bigger in their valuations than their NC counterparts (as per descriptive statistics in Table 1). Against this backdrop, we argue that government-forced changes to pension funds AUM and investment policy of 2014 resulted in lower IPO proceeds. We also provide first robust empirical evidence in support of the practitioners' claim that the described regulatory changes suppressed the level of IPO valuations, since historically pension funds have been crucial buyers of shares responsible for the so-called 'pension premium.'

As reported in Table 4, we find some evidence that government-forced changes on the demand side affected IPO first-day returns. Specifically, the difference-in-differences estimate for the first-day risk-adjusted returns shows a decline of 17 percentage point, significant at 5% level.

[18] Exchange rates are as per the official Central Bank of Poland statistics; retrieved on February 23, 2019, from https://www.nbp.pl/homen.aspx?navid=archen&c=/ascx/TabArchAen.ascx&n=14a022en for February 3, 2014.

This decline is consistent with the book building literature which argues that IPO underpricing is a mean of compensating institutional investors for revealing value-relevant information during the book building process (Aggarwal et al., 2002; Benveniste & Spindt, 1989; Chemmanur et al., 2010). Furthermore, we see no impact on the long-term IPO performance, which is consistent with the fact that the association between institutional trading and long-term IPO performance decays over time (Chemmanur et al., 2010).

For robustness purposes, we additionally test whether our results for the effect of regulatory change on the IPO short- and long-term underperformance could have been biased by improper treatment of IPO observations that listed during the pre-treatment period but which 30-day and 360-day interval for aftermarket performance measurement are also in the range of the post-treatment period. In other words, we investigate whether the short-term performance of stocks that debuted between January 3, 2014 and February 2, 2014, and long-term performance of stocks that debuted between February 3, 2013 and February 2, 2014 capture some impact of the policy change. To test this, we adjust the threshold so that the post-intervention period includes all the returns that could have been potentially affected. The results remain unchanged. We report these results in Table C9 in Appendix C.

In Table 5 we report the results from estimating the difference-in-differences regression that includes additional controls for firm characteristics. We control for size, profitability, and VC-backing status, as described earlier.

[Insert Table 5 here]

The results presented in Table 5 confirm our main results. Specifically, we report material and statistically significant change in the IPO size (winsorized at 0.5% and 99.5%). Our results indicate that the average IPO size in the treated group decreased by 96 million PLN. In terms of economic magnitude, this estimate is smaller than the estimate in the regression that omits controls (128 million PLN) but remains statistically significant. This result supports our conclusion that the government-forced changes to pension funds AUM and investment policy of 2014 resulted in lower IPO proceeds.

An alternative (and simpler) specification of our model is when instead of the year fixed effects, we include the post-treatment indicator, $Post_i$. Our results are robust to this alternative specification (see Table C6 and Table C11 in Appendix C for the results in the models without and with the control variables, respectively). Furthermore, explicitly including the $Post_i$ variable allows us to make linear predictions. As per table C6 in Appendix C, we report the overall decrease in the IPO size equal to 132 million PLN and the decrease in the degree of the first-day underpricing of 11 percentage points in the WSE. Although pension funds hold only 12% of the equity in the WSE, their effect on IPO pricing is undoubtedly material.

We next supplement our analysis with a cross-sectional test based on the IPO size. Zalewska (2006) documents a blue-chip bias among Polish pension funds, whereby pension funds prioritize issuers with the largest capitalization in their equity portfolios.[19] Since pension funds' investments are concentrated around the local blue-chips, we expect our results to be particularly evident in the subsample that includes the largest IPOs. We motivate our prior with the fact that the impact of the regulatory change was strongest for the firms with the highest pension funds' participation. Accordingly, each year we sort all IPOs based on the quartiles of IPO size, separately for the treatment and control sample.[20] We then rerun our difference-in-differences regressions across the four size partitions. We report the results in Table 6. For brevity, we only show the coefficients on the interaction term ($WSE \times Post$), which is our main coefficient of interest.

[Insert Table 6. here]

The table provides strong support for our hypothesis. Larger issuers drive the negative impact on the IPO size. The difference-in-differences estimates are statistically significant for the two biggest partitions (quartile 3 and 4), but insignificant for the smallest two partitions (quartile 1 and 2).[21] Specifically, using the short-window test, the decline is 24.5 million PLN in the third size quartile, and 669.3 million PLN in the fourth quartile, both estimates statistically significant. The decrease in the largest size quantile remains significant in the long-time window test. This result is consistent with our prediction that the decline in the IPO size that we observe in the full WSE IPOs sample is a result of the 2014 regulatory treatment of the Polish pension funds, which invest primarily in the larger companies (Field & Lowry, 2009).

Furthermore, Table 6 reveals a statistically significant decline in the IPO first-day returns following the regulation in the largest-IPO-size partition. Specifically, the average IPO first-

[19] The large-cap bias is not confined to the Polish stock market only. It is also a significant regulatory challenge in the developed markets (see Fleming, 2017).

[20] By sorting stocks separately for WSE and NC samples our quantile test compares the IPOs based on the relative size. To make sure that the differences in firm size between treatment and control firms in the given portfolios are not driving this result, we additionally rerun this analysis using an alternative specification that controls for firm size, in addition to profitability and VC backing status. We report those results in Table C12 in Appendix C.

[21] We also report the coefficients on the log size, as the distribution of IPO size is less asymmetric in the narrowly-defined size partitions. Otherwise, the log transformation introduces the attenuation bias. This is because the larger IPOs have higher Pension funds ownership and hence contribute the most to our results.

day returns decreased by 30.0 percentage points when we used a short-term window implementation strategy, and 24.1 percentage points when we used a long-term window, both statistically significant. This decline is consistent with the book building literature (e.g., Benveniste & Spindt, 1989), which posits that IPO underpricing is a mean of compensating institutional investors for revealing value-relevant information during the book building process (Chemmanur et al., 2010). If there is a positive association between the institutional investors' participation in the IPO book building process and the level of IPO underpricing, then an exogenous shock to institutional investors' participation, that creates a negative demand shock, results in a lower degree of IPO underpricing.

Then, we report weak evidence of a negative impact of the 2014 reform on the IPO long-term underperformance. In the largest-IPO-size partition, there is a statistically significant decline in the 30-day market returns of 69.4 percentage point when using the short-time window and of 63.1 percentage point when using the long-time window. This result is consistent with the information asymmetry literature (Field & Lowry, 2009), which suggests a positive association between institutional trading and IPO long-term performance. Also, there are no significant effects in the 360-day post-IPO period. This is in line with the stylized fact that the association between institutional trading and IPO long-term performance decays over time (Chemmanur et al., 2010).

Lastly, we discard the possibility of more firms deciding to stay private after the reform. First, we do not observe a reduction in the number of IPOs after the reform (in contrast to the U.S. market as reported by Gao, Ritter, & Zhu, 2013). As per Table 1, we report that in the years following the 2014 reform, IPO activity proxied by the number of IPOs did not decrease as compared to the years preceding the reform. We also look at the number of IPO withdrawals for each year in our time-series. Based on the Bloomberg data, the withdrawal rate was 11% in 2013 (the year before the reform), 3.4% in 2014 (event year). Between 2005 and 2013, an average yearly withdrawal rate was 10%, while between 2014 and 2017, it was 3% on average. In our study, we also report lower underpricing in the period after the pension reform in contrast to Dorsman and Gaunopoulos' (2013) findings. Against this backdrop, we conclude that it is unlikely that the pension reform resulted in fewer companies going publish with equity in the stock market of Poland. Conversely, we argue that after the reform market participants constituted a 'new market equilibrium', where listing companies accept new market conditions (i.e. the fact that pension funds have less money to invest) and they go on with an IPO despite the realization that their IPO proceeds would be lower than if they listed before 2014.

6.3. Robustness

Finally, we examine whether our results are robust with regard to the date of the regulatory change. Specifically, we would like to disprove an argument that IPO size changes multiple times in our sample period, and the regulatory change only happened to be one of such events. We are also interested whether the precise date of the event is the very moment of changes in IPO proceeds. The alternative view is that changes in the IPO pricing might have been occurring gradually in anticipation of the new regulation (pension funds demand for new equities decreased already when there were rumours about the oncoming changes to the Polish pension system). Figure 7 plots the *p*-values of the difference-in-differences coefficient in the regression where the IPO size is the response variable, relative to the alternating cut-off date. A visual inspection enables us to validate that February 2014 was indeed a critical timing for material changes in the average IPO size in the WSE treated group. Only at the beginning of the year 2014, the *p*-value of the difference-in-differences coefficient falls to the close-to-zero level, which means that the estimate becomes statistically significant. This result is particularly pronounced when we use the symmetric bandwidth of 3 years around the threshold (short-time window implementation strategy). Moreover, the plot indicates that the result is relatively insensitive to the exact cut-off date: the difference-in-differences estimate is statistically significant at the 5% level as long as we choose a date around the implementation of the regulatory change. We argue that this is driven by the nature of our data (IPOs do not happen every day). A lack of observations for each trading day, while running placebo testing using each trading day as the cut-off, results in the *p*-values being significant for the period around the event rather than for an individual point in time. However, there is no trait of IPO size being lower in the period preceding the regulatory change when it was being prepared and consulted. Hereto, we do not report statistically significant *p*-value estimates for the period of one year preceding the event. Further, we also reject the argument that the event might have been expected by companies that hastened the IPO process in anticipation of the law change to avoid lower IPO proceeds. In the data we do not see a tendency to go public earlier to avoid declining demand. The number of IPOs in the WSE market in 2014 is higher than its counterpart in 2013 (see Table 1 for the number of IPOs each year). The placebo also addresses the concerns that our results are directly attributable to the Global Financial Crisis because most of the simulated windows are contained entirely in the post-crisis period.

[Insert Figure 7. here]

Accordingly, we argue that the estimates for the causal effect of regulatory changes on the IPO size presented in Table 4 are robust to the placebo testing related to the event date.

7. Conclusion

The Polish pension system was established in 1999, and in time proved unsustainable. In early 2014, a pension fund reform was implemented to solve the country's fiscal problems. The main changes of the 2014 reform encompassed lower contributions to the pension plans managed by open (private) funds, followed by the removal of the internal benchmark mechanism as a measure of pension fund's performance. These changes had a direct effect on pension funds AUM and their investment policy. In this paper, we go one step further and investigate the indirect effects of the 2014 regulatory change on capital markets, and specifically on the IPO pricing and stock's subsequent performance.

Using a novel data sample, we provide robust evidence for the IPO underpricing effects in the two equity markets in Poland. We find that between 2005 and 2017 an average WSE (NC) IPO yields a 10.59% (37.56%) market-adjusted first-day return when equal weights are used and 8.26% (24.1%) first-day market-adjusted return when IPO-size weights are used. Significant differences in returns between equally- and IPO-size-weighted portfolios indicate the superior performance of small IPOs. The estimates for short- and long-term performance of stocks listed in both equity markets (cumulated over 30 and 360 days after an IPO) confirm the long-term underperformance effect. The average 360-day market-adjusted return for an IPO debuting in the WSE (NC) is 11.61% (6.49%) when equally-weighted, and -0.31% (-10.53%) when IPO-size-weighted. Notably, the test for long-term IPO underperformance is dependent upon the risk adjustment of the returns.

The unique opportunity to exploit a quasi-natural experiment arising from the Polish government's regulatory treatment of pension funds enables us to shed new light on what factors influence new issues' market. We show that the decreasing net cash inflows to pension funds' AUM and abandoning internal benchmark as a measure of funds' performance result in less demand for new equities that manifests itself in significantly smaller IPO size. In the reform aftermath, firms that list in the WSE market, raise substantially less money in the process: 142 million PLN less on average using a simple *t*-test approach and estimating the difference between the pre- and post-treatment mean IPO size. Regression results reinforce our conclusions and confirm the causal effect—the incremental impact of the regulatory-forced changes to pension funds AUM and investment policy on the IPO size is estimated at -107 million PLN and is statistically significant. This result is robust to using different bandwidth around the cut-off date placebo testing. There is no such effect in the NC market, where pension funds are not invested. We leave the question of whether the impact of the reform on valuations reverses in the long-term as a potential avenue for further research.

Testing for the differences in the IPO underpricing effects before and after the government-forced changes to pension funds' AUM and investment policy reveals that an exogenous shock had some impact on the degree of underpricing. The effect was especially among larger firms, where pension funds are primarily invested. We report no impact of the reform on the degree of IPO long-term underperformance. Our findings contribute to the literature on the drivers of IPO underpricing. We report a significant decrease in the IPO underpricing for the subsample of issuers where the Polish pension funds invest the most. This decline is consistent with the rich book building literature which states that IPO underpricing is a mean of compensating institutional investors for revealing value-relevant information during the book building process (Baron, 1982; Beatty & Ritter, 1986; Benveniste & Spindt, 1989; Chemmanur et al., 2010; Rock, 1986).

We can interpret our results in a negative tone, that is after the regulatory change companies raise less money in the primary stock market, or in a positive tone, that is the reform eliminated an artificial 'pension premium' of higher valuations, and thus contributed to market efficiency. In any case, reforming pension system for the sake of solving fiscal problems turned out to have spillover effects on capital markets. Specifically, a government trying to fix the unsustainable pension system by limiting net cash inflows to the privately-managed pension funds directly contributed to a decrease in the average IPO size. By providing empirical evidence thereof, we deliver novel insights into the existing IPO literature.

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Figures and Tables

Figure 1. Unadjusted first-day returns of WSE and NC IPOs from January 2005 to June 2017.

The figure plots the first-day return of all IPOs listed on WSE and NC in the period from January 2005 to June 2017. The first-day return is calculated as the difference between the closing price of the first day of trading and the IPO offer price, divided by that closing price. For presentation purposes, we exclude 7 observations that have the first-day returns exceeding 300%.

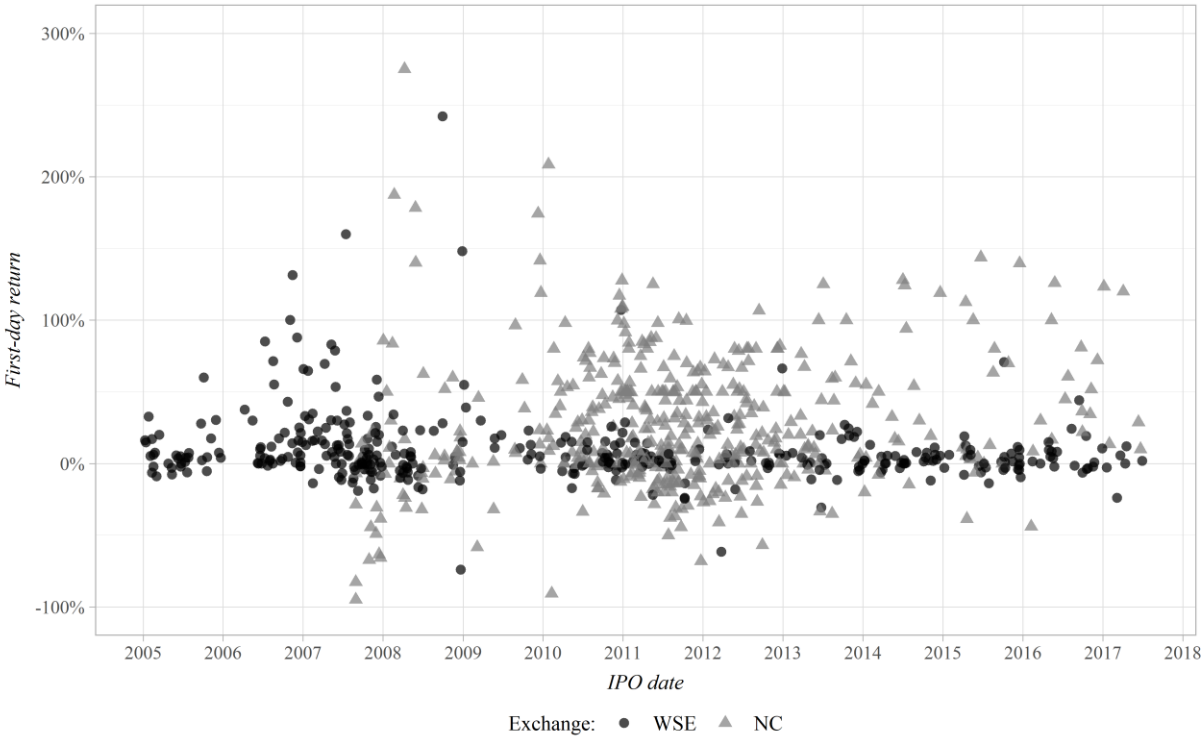
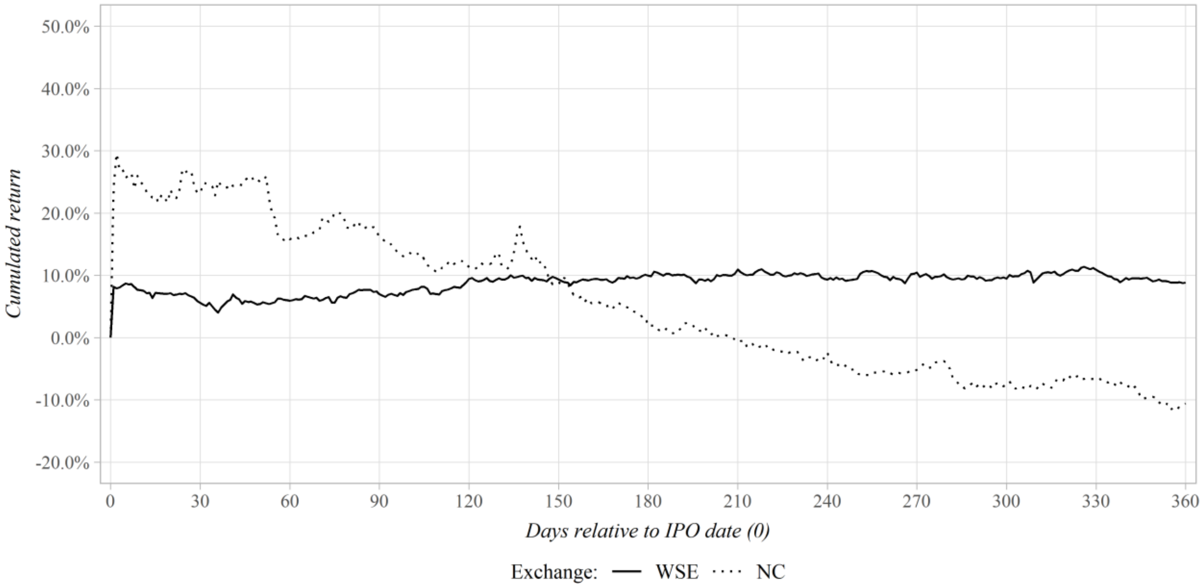


Figure 2. Unadjusted short- and long-term performance of WSE and NC IPOs from January 2005 to June 2017.

Panels A and B show the performance of stocks following an IPO. Returns are cumulated over the relevant period using the CAR procedure. The solid line depicts the performance of the WSE IPO portfolio, while the dotted line represents the performance of the NC IPO portfolio. Portfolios are formed using 99% winsorized IPO size weights (Panel A) or equal weights (Panel B). A company is included in a portfolio unless, for any day, any data regarding this company is missing. If data is missing, the company is excluded from further calculations. The sample for the first-day return calculations includes 385 (486) companies for the WSE (NC). For 30-day and 360-day returns, the sample size is 347 (395) and 332 (382) for the WSE and NC, respectively due to firms being acquired or delisted before the 30-day or 360-day period finishes.

Panel A (SW)



Panel B (EW)

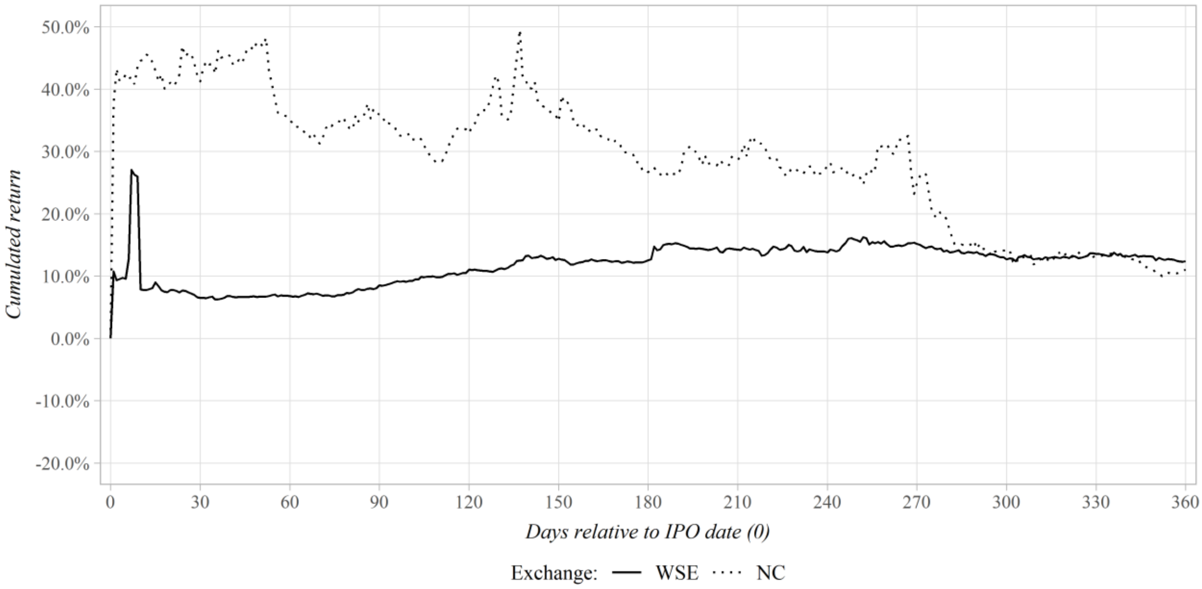
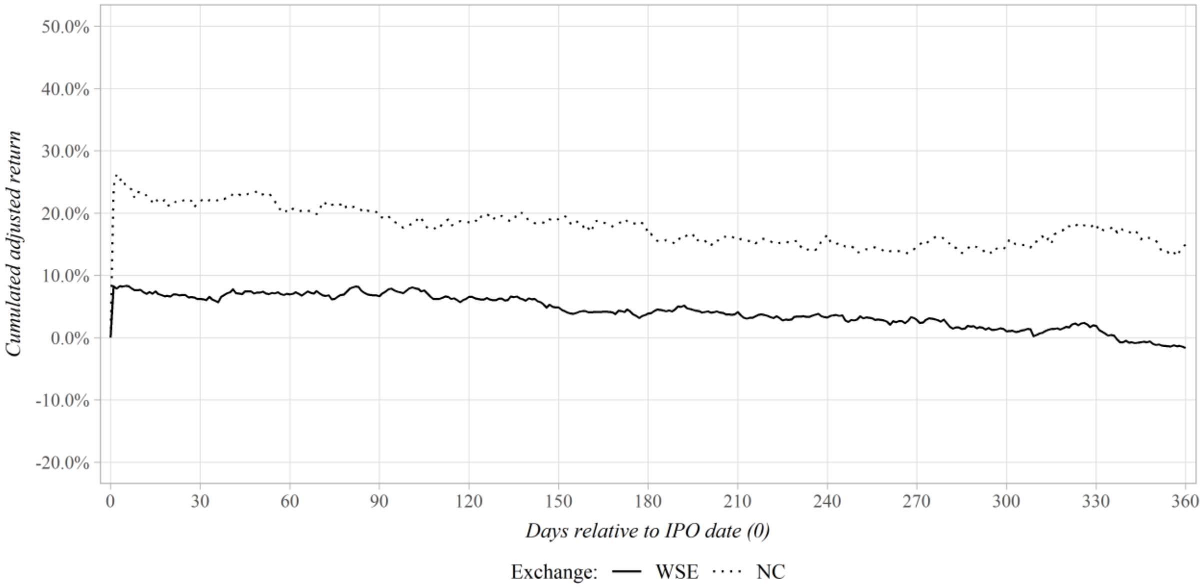


Figure 3. Market-adjusted short- and long-term performance of the WSE and NC IPOs from January 2005 to June 2017.

Panel A and B show the stock performance following an IPO. Returns are market-adjusted and cumulated over the relevant period using CAR procedure. The solid line depicts the performance of the WSE IPOs portfolio, while the dotted line represents performance of portfolios of the NC IPOs portfolio. Portfolios are formed using IPO-size weights winsorized at 0.5% and 99.5% (Panel A) or equal weights (Panel B). A company is included in a portfolio unless for any day, any data regarding this company is missing. If there is data missing, a company is excluded from further calculations. The sample for the first-day return calculations includes 385 (486) companies for the WSE (NC). For 30-day and 360-day returns, the WSE (NC) sample size is 347 (395) and 332 (382), respectively.

Panel A (SW)



Panel B (EW)

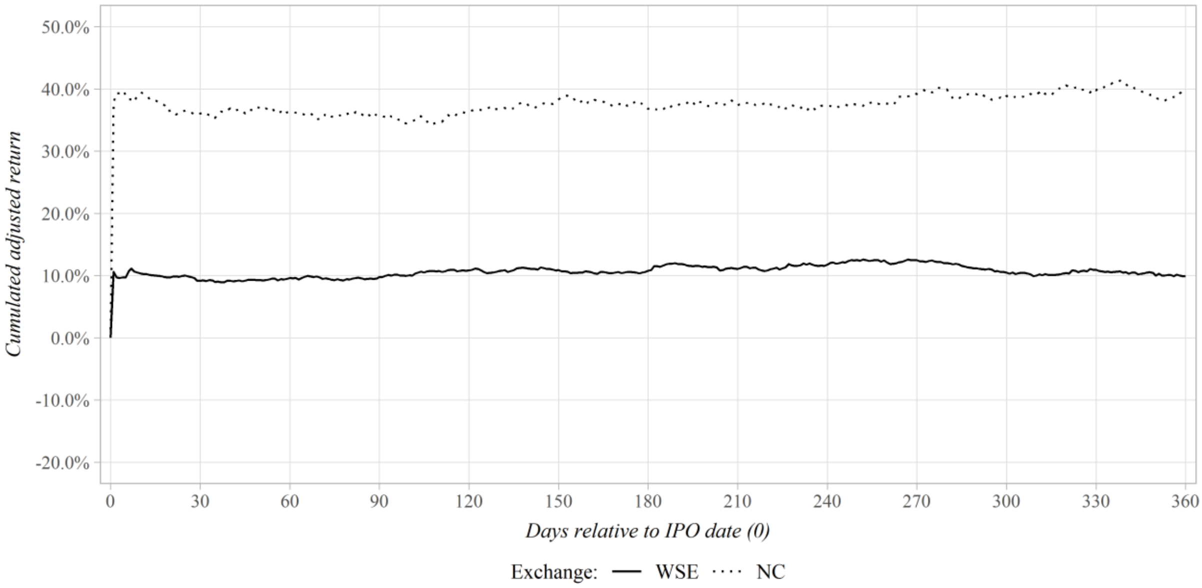


Figure 4. Cumulated percentage changes in the aggregated pension funds' AUM, in the equity part of the aggregated pension funds' AUM and in the Warsaw Stock Exchange WIG index.

Consecutive plots in the figure depict the cumulated percentage changes in the aggregated pension funds' AUM, in the equity part of the aggregated pension funds' AUM and in the Warsaw Stock Exchange WIG index. The changes are scaled relative to December 31, 2013.

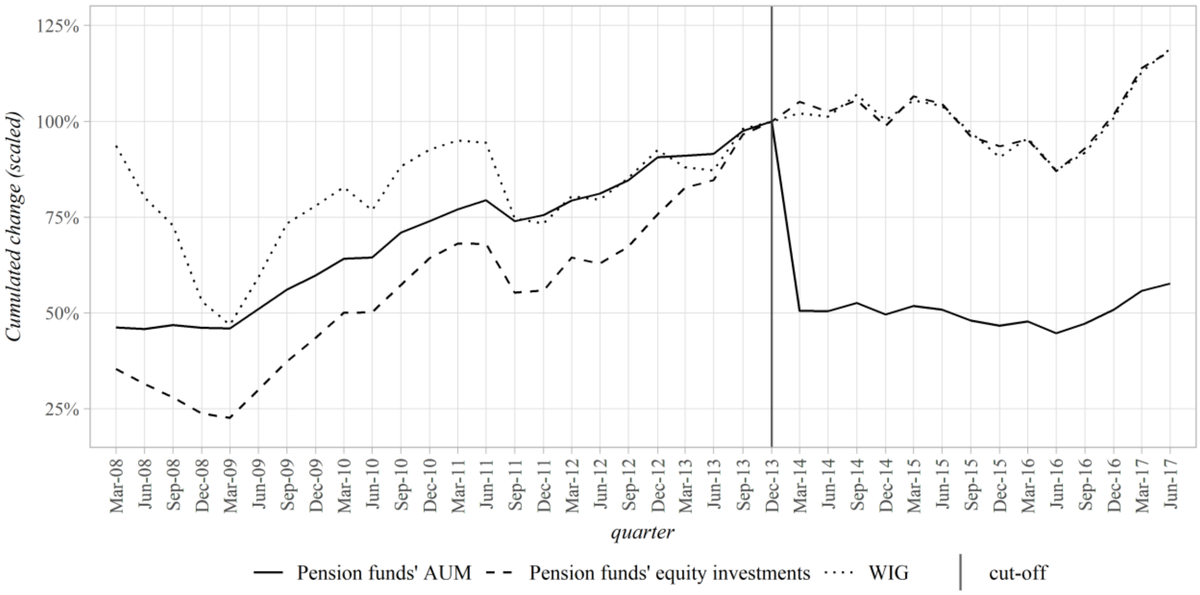
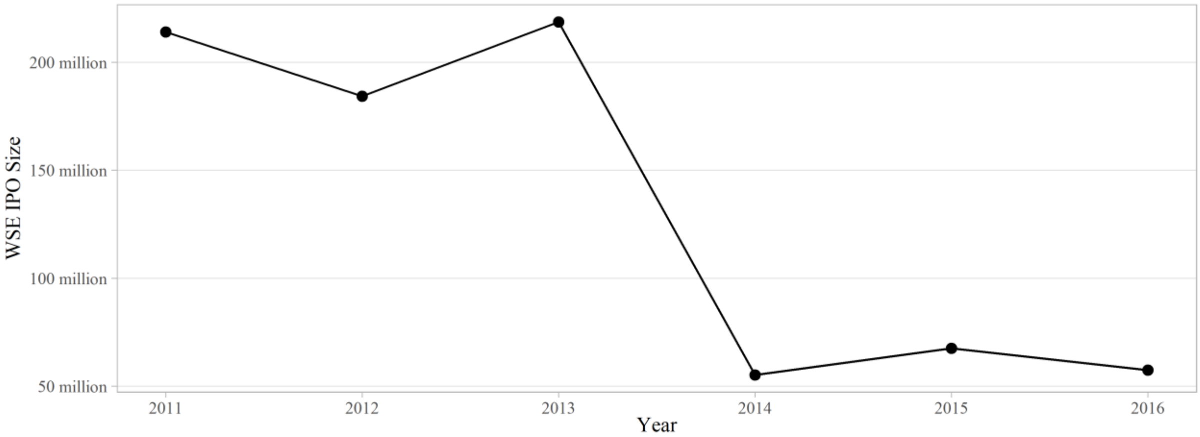


Figure 5. Trend line of the mean IPO size by year in the WSE and NC from January 2005 to December 2016.

Figure 5 shows the changing mean IPO size from year on year. Panel A and B show annual mean IPO size for the WSE and NC listings, respectively. The IPO size is defined as number of shares issued multiplied by the IPO price. The IPO size are reported in million PLN. The data frame spans from January 2011 to December 2016 (Panel A) and from January 2007 to December 2016 (Panel B). Year 2017 is omitted because the sample period ends in June 30.

Panel A (WSE)



Panel B (NC)

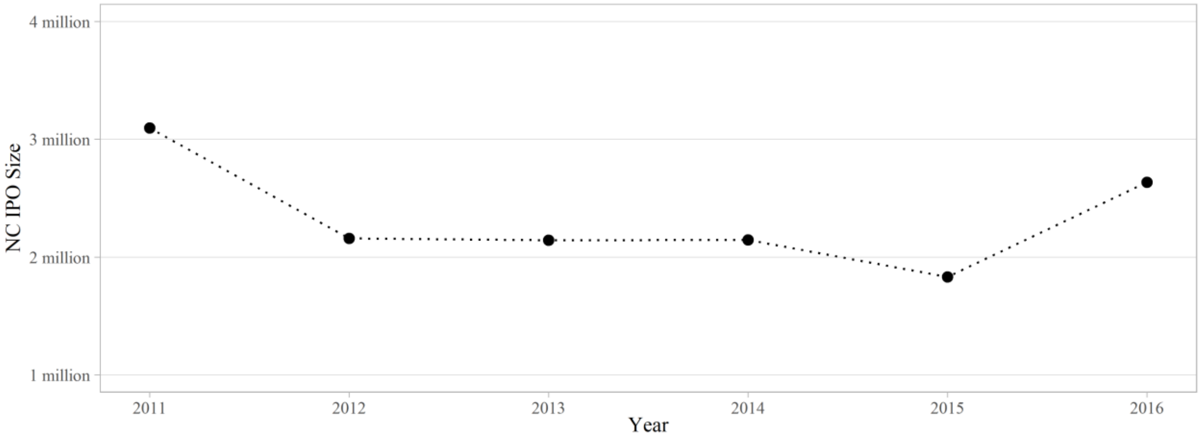
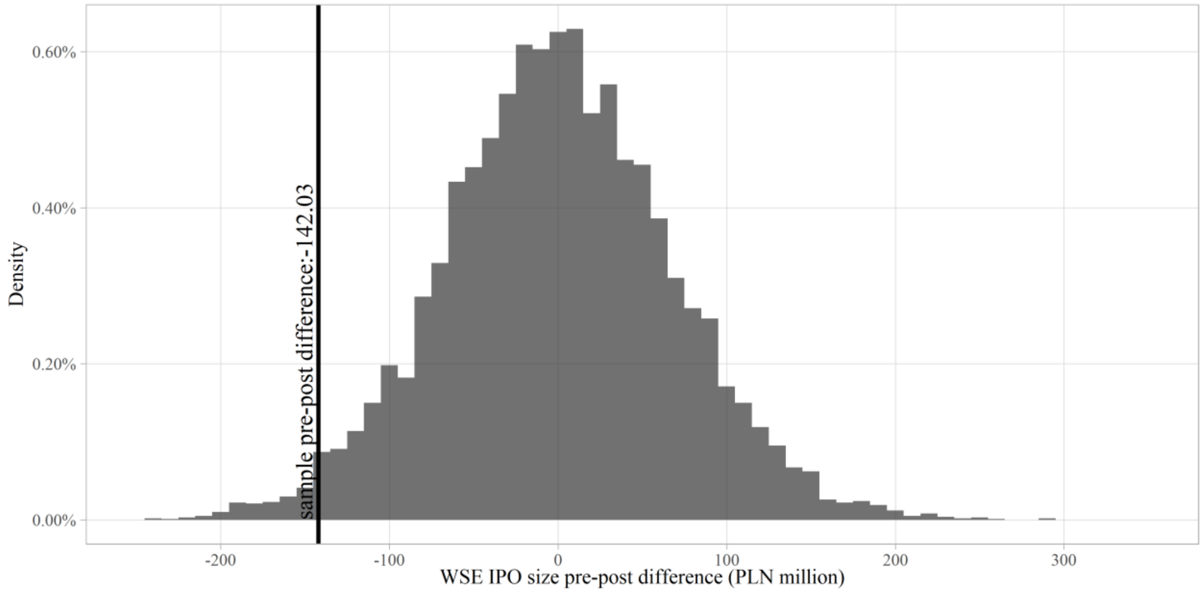


Figure 6. Bootstrapped pre-post differences in IPO size in the WSE and NC.

Figure 6 shows the empirical distribution of bootstrapped pre-post differences in IPO size for the WSE (Panel A) and NC (Panel B). The IPO size is defined as number of shares issued multiplied by the IPO price. The IPO size is reported in million PLN. For the purpose of bootstrapping, we sample IPOs with replacement and calculate difference between mean IPO size in the pre-treatment period and mean IPO size in the post-treatment period. We use February 3, 2014 as our cut-off date. We use a symmetric 3-year period around the cut-off in IPO sampling procedure (the short-time window). We plot the density based on 10,000 sampling iterations. The original sample differences are illustrated with the solid vertical lines.

Panel A (WSE)



Panel B (NC)

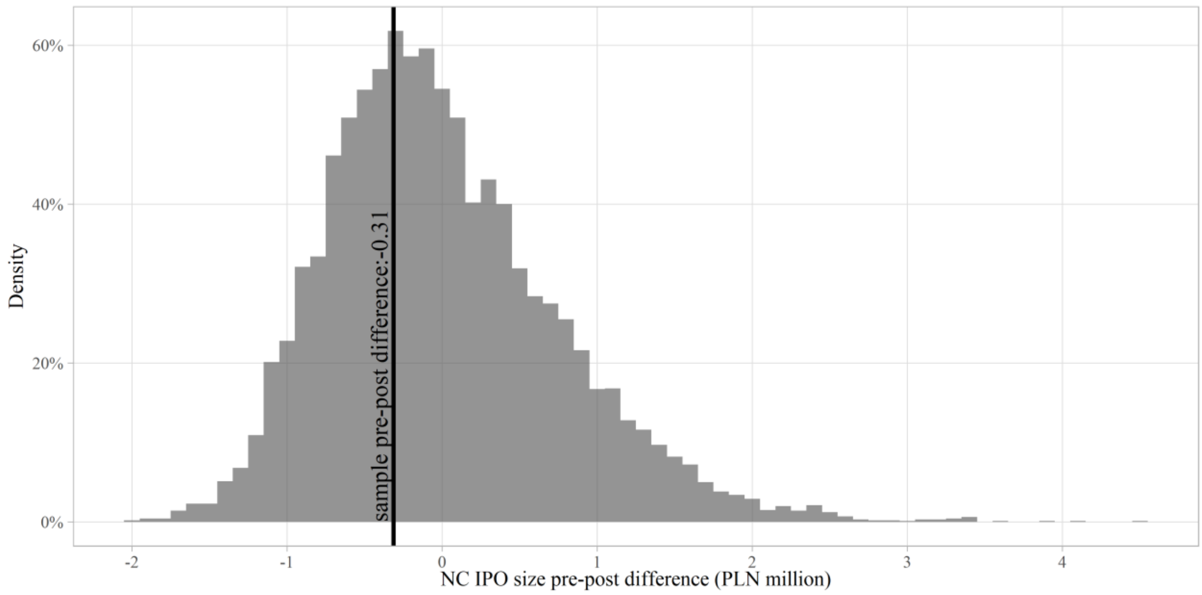


Figure 7. Placebo test for the robustness of the regulatory treatment effect on the IPO size.

Figure 7 shows the results of the placebo test of our main results. For each day (cut-off), we run the regression $Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \gamma_t + \delta_j + \varepsilon_i$. IPO size (in million PLN) is the response variable in the regression. WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the WSE IPO listings. $Post$ is an indicator variable that takes on a value of 1 for IPO listings post the cut-off date, and 0 for the pre-cut-off IPO listings. $WSE \times Post$ is the variable of interest which captures the difference-in-differences effect. All regressions include industry and time fixed effects. We plot the p-values of the coefficient of interest (β_2). In estimation, we use the symmetric period +/- 3 years relative to the cut-off date (short-time window; solid line) and full sample, starting on January 1, 2008 (long-time window; dotted line).

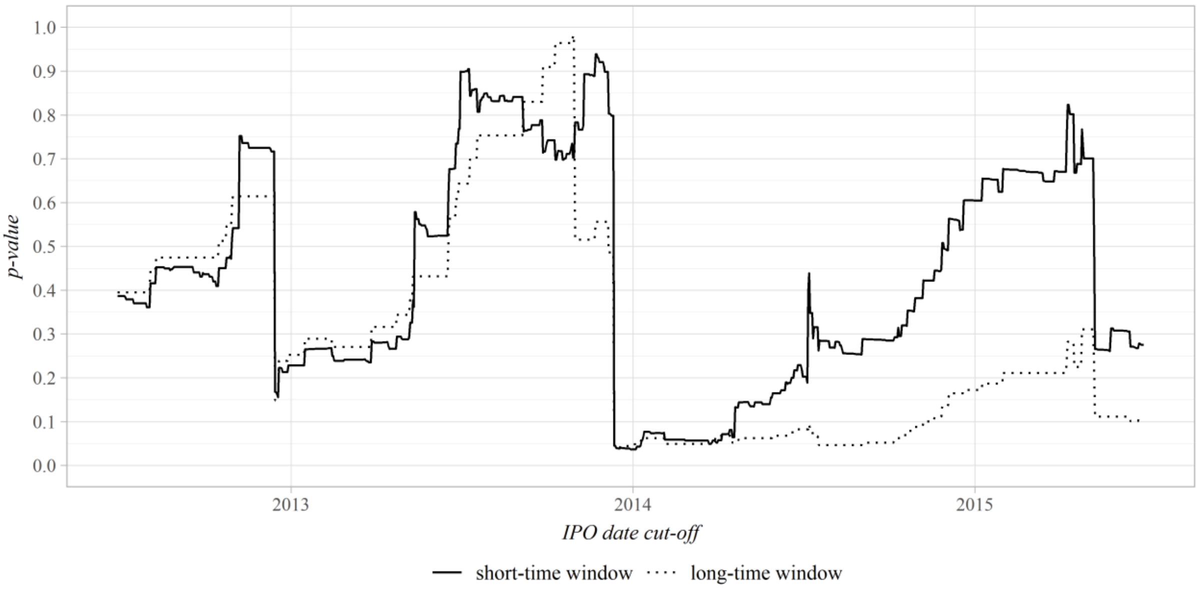


Table 1. Summary statistics for IPOs listed in the WSE and NC from January 2008 to June 2017.

Panel A contains descriptive statistics regarding IPO returns in the WSE, and Panel B contains the statistics for IPO returns in the NC. IPO size is the number of shares issued at an IPO multiplied by the IPO offer price. All returns are in decimal format; IPO size is in million PLN. Unadjusted returns are calculated using the HPR, whereas market-adjusted returns are computed using the CAR methodology. The average returns are equally weighted (EW average) or IPO-size-weighted (SW average). Our sample extends from January 2008 to June 2017. The sample size for the IPO size is 706, including 237 IPOs in the WSE and 469 IPOs in the NC. The sample diminishes in size when we analyse post-IPO short- and long-term underperformance. For reference, we report the number of observations in the last row of each panel.

Sample statistic	IPO Size	1st day return	1st day market-adjusted return	30-day return	30-day market-adjusted return	360-day return	360-day market-adjusted return
Panel A: Summary statistics for companies listed in the WSE.							
EW Avg.	177.001	0.058	0.056	0.018	0.017	0.038	0.012
VW Avg.[a]	-	0.055	0.056	0.049	0.048	0.121	0.042
Std. Dev.	549.801	0.241	0.242	0.205	0.211	0.574	0.524
Min	0.000	-0.741	-0.729	-0.716	-0.672	-0.899	-1.130
P25	0.000	-0.023	-0.023	-0.076	-0.085	-0.293	-0.325
P50	26.000	0.019	0.019	0.004	0.007	0.000	-0.027
P75	96.300	0.084	0.081	0.104	0.117	0.235	0.227
Max	3,652.757	2.420	2.448	0.783	0.912	3.451	3.250
Obs.	237	235	235	222	222	207	207
Panel B: Summary statistics for companies listed in the NC.							
EW Avg.	2.609	0.400	0.401	0.431	0.430	0.126	0.072
VW Avg. [a]	-	0.254	0.254	0.216	0.213	-0.075	-0.108
Std. Dev.	5.721	1.505	1.506	2.976	2.979	1.845	1.828
Min	0.000	-0.906	-0.916	-0.908	-1.016	-0.973	-1.373
P25	0.571	0.000	0.000	-0.200	-0.192	-0.574	-0.587
P50	1.090	0.185	0.187	0.006	0.013	-0.200	-0.253
P75	2.580	0.518	0.512	0.343	0.351	0.169	0.167
Max	85.138	29.000	28.992	51.000	51.010	25.331	25.132
Obs.	469	468	467	382	381	370	369

[a] IPO size weighting with IPO size winsorized at 0.5% and 99.5%.

Table 2. IPO Sector and Year Distribution

Panel A shows the sample distribution of the WSE and NC IPOs by the GICS sector. Panel B shows the sample distribution of the WSE and NC IPOs by the IPO year. Our sample extends from January 2008 to June 2017. The sample size is 706 IPOs, including 237 IPOs in the WSE and 469 IPOs in the NC. The sample diminishes in size when we analyse post-IPO short- and long-term underperformance. For reference, we report the number of observations in the last row of each panel.

Panel A: Sample distribution by the sector

Exchange	**WSE (no. and %)**		**NC (no. and %)**	
Energy	7	3.0%	2	0.4%
Materials	18	7.6%	16	3.4%
Industrials	45	19.0%	114	24.3%
Consumer Discretionary	44	18.6%	110	23.5%
Consumer Staples	17	7.2%	20	4.3%
Health Care	18	7.6%	30	6.4%
Financials	35	14.8%	61	13.0%
Information Technology	24	10.1%	74	15.8%
Telecommunication Services	1	0.4%	13	2.8%
Utilities	8	3.4%	10	2.1%
Real Estate	20	8.4%	19	4.1%
TOTAL	237	100.0%	469	100.0%

Panel B: Sample distribution by the IPO year

Exchange	**WSE (no. and %)**		**NC (no. and %)**	
2008	32	13.5%	48	10.2%
2009	13	5.5%	18	3.8%
2010	34	14.3%	78	16.6%
2011	34	14.3%	150	32.0%
2012	18	7.6%	83	17.7%
2013	23	9.7%	36	7.7%
2014	27	11.4%	19	4.1%
2015	29	12.2%	17	3.6%
2016	19	8.0%	14	3.0%
2017[a]	8	3.4%	6	1.3%
TOTAL	237	100.0%	469	100.0%

[a] As of June 30, 2017.

Table 3. Two-sample *t*-test of pre- and post-treatment mean returns and IPO sizes.

The table presents pre- and post-treatment mean values of the first-day, 30-day, and 360-day market-adjusted returns and IPO size for WSE and NC listings. In Panel A and B, we show the results using a short-time window implementation strategy (3-year symmetric windows; the period of February 4, 2011, till February 3, 2017) and a long-time window implementation strategy (January 1, 2008-July 30, 2017) respectively. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. For reference, we report the number of observations in the last row of each panel. Returns are market-adjusted and cumulated over the relevant period as CARs and are reported in percentages. The IPO size equals the IPO offer price multiplied by the number of shares issued and is reported in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

	IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.				
WSE				
Pre-mean	201.248	0.023	0.023	-0.041
Post-mean	59.220	0.035	0.022	0.060
Difference	-142.028 **	0.012	0.000	0.101
t-statistic	(2.133)	(-0.548)	(0.005)	(-1.086)
Obs.	148	148	143	132
NC				
Pre-mean	2.586	0.326	0.318	-0.106
Post-mean	2.273	0.449	1.050	0.194
Difference	-0.313	0.123	0.732	0.299
t-statistic	(0.614)	(-0.916)	(-1.292)	(-1.071)
Obs.	306	305	257	246
Panel B: Implementation strategy using a long-time window.				
WSE				
Pre-mean	220.200	0.067	0.019	-0.008
Post-mean	92.224	0.033	0.013	0.060
Difference	-127.976 **	-0.035	-0.005	0.067
t-statistic	(2.222)	(1.323)	(0.191)	(-0.920)
Obs.	237	235	222	207
NC				
Pre-mean	2.616	0.394	0.350	0.059
Post-mean	2.553	0.454	1.019	0.194
Difference	-0.064	0.060	0.669	0.134
t-statistic	(0.122)	(-0.579)	(-1.272)	(-0.470)
Obs.	469	467	381	369

Table 4. Effects of the 2014 regulatory change on the IPO underpricing, long-term underperformance and on the IPO size.

The table reports results for the regression $Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \gamma_t + \delta_j + \varepsilon_i$. The dependent variable is either: the IPO size (in million PLN), the first-day market-adjusted return, the 30-day market-adjusted return, or the 360-day market-adjusted return (all in percentages), and our unit of observation is a single IPO. WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the NC IPO listings. $Post$ is an indicator variable that takes on a value of 1 for post-event IPO listings, and 0 for the pre-event IPO listings, $WSE \times Post$ is the variable of interest that captures the difference-in-differences effect. All regressions include industry and year fixed effects. Year fixed effects indicate the year of the IPO. t-statistics are reported in parentheses. Data are for the period February 4, 2011-February 3, 2017 for Panel A, and January 1, 2008-July 30, 2017 for Panel B. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. For reference, we report the number of observations in the last row of each panel. Returns are in decimal format; IPO size is in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. We winsorize the observed IPO sizes at 0.5% and 99.5%.

	IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.				
WSE	184.960 ***	-0.213 ***	-0.154	0.077
	(2.797)	(-4.367)	(-1.433)	(0.702)
WSE x Post	-127.825 *	-0.174 **	-0.872	-0.219
	(-1.895)	(-2.060)	(-1.620)	(-0.870)
Industry FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R^2	0.136	0.040	0.032	0.060
Obs.	454	453	400	378
Panel B: Implementation strategy using a long-time window.				
WSE	194.130 ***	-0.317 ***	-0.271 ***	-0.135
	(4.039)	(-5.382)	(-3.309)	(-0.930)
WSE x Post	-106.689 **	-0.103	-0.737	-0.052
	(-1.969)	(-1.251)	(-1.463)	(-0.202)
Industry FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R^2	0.180	0.044	0.029	0.027
Obs.	706	702	603	576

Table 5. Effects of the 2014 regulatory change on the IPO underpricing, long-term underperformance and on the IPO size – alternative regression specification with additional control variables.

The table reports results for the regression $Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \beta_3 \log(Assets)_i + \beta_4 ROA_i + \beta_5 I(VC)_i + \gamma_t + \delta_j + \varepsilon_i$. The dependent variable is either: the IPO size (in million PLN), the first-day market-adjusted return, the 30-day market-adjusted return, or the 360-day market-adjusted return (all in percentages), and our unit of observation is a single IPO. WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the NC IPO listings. $Post$ is an indicator variable that takes on a value of 1 for post-event IPO listings, and 0 for the pre-event IPO listings, $WSE \times Post$ is the variable of interest that captures the difference-in-differences effect. log $(Assets)$ are the total assets in the last fiscal year prior to the IPO. ROA is the return on assets in the last fiscal year prior to the IPO, using the end-of-period value of total assets in the denominator. $\mathrm{I(VC)}$ is an indicator of whether an IPO was Venture Capital backed. All regressions include industry and year fixed effects. Year fixed effects indicate the year of the IPO. t-statistics are reported in parentheses. Data are for the period February 4, 2011-February 3, 2017 for Panel A, and January 1, 2008-July 30, 2017 for Panel B. The sample size varies depending on the length of the sample window as well as data requirements for the dependent variable. For reference, we report the number of observations in the last row of each panel. Returns are in decimal format; IPO size is in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. We winsorize the observed IPO sizes at 0.5% and 99.5%.

	IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.				
WSE	35.260	-0.068 *	0.052	0.089
	(1.405)	(-1.944)	(0.392)	(0.653)
WSE x Post	-96.477 *	-0.238 ***	-1.009 *	-0.235
	(-1.808)	(-3.876)	(-1.724)	(-0.895)
Controls	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R2	0.195	0.198	0.099	0.085
Obs.	415	414	371	351
Panel B: Implementation strategy using a long-time window.				
WSE	-39.346	-0.171 ***	-0.082	0.040
	(-1.214)	(-3.328)	(-0.911)	(0.384)
WSE x Post	-61.227	-0.172 **	-0.881	-0.208
	(-1.378)	(-2.487)	(-1.643)	(-0.833)
Controls	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R2	0.302	0.117	0.087	0.062
Obs.	643	639	556	531

Table 6. Issuer effects of the 2014 regulatory change on the IPO underpricing, long-term underperformance and on the IPO size across different IPO size sample partitions.

The table reports difference-in-differences estimates for the regression $Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \gamma_t + \delta_j + \varepsilon_i$. We estimate the regression for the IPO size partitions. To construct the size partitions, we sort the sample based on IPO size each year. Prior to sorting, we exclude issuers with IPO size unavailable or reported as 0 by the exchange. Small issuers are issuers in the smallest size quartile (bottom 25% issuers). Big issuers are issuers in the largest size quartile (top 20% issuers). We also report intermediate portfolios (labelled portfolio 2 and 3) corresponding to the second and third quartile of IPO size. The dependent variable is either: the IPO size (in million PLN), the first-day market-adjusted return, the 30-day market-adjusted return, or the 360-day market-adjusted return (all in percentages), and our unit of observation is a single IPO. Log IPO size is the natural logarithm of IPO size. In the regression, WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the NC IPO listings. $Post$ is an indicator variable that takes on a value of 1 for post-event IPO listings, and 0 for the pre-event IPO listings, $WSE \times Post$ is the variable of interest that captures the difference-in-differences effect. All regressions include industry and year fixed effects. Year fixed effects indicate the year of the IPO. We only report difference-in-differences estimates, which are the coefficients on $WSE \times Post$. t-statistics are reported in parentheses. Data are for the period February 4, 2011-February 3, 2017 for Panel A, and January 1, 2008-July 30, 2017 for Panel B. The sample size varies depending on the length of the sample window as well as data requirements for the dependent variable. Returns are in decimal format; IPO size is in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. We winsorize observed IPO sizes at 0.5% and 99.5%.

	IPO Size	Log IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.					
Small	-6.616	-0.483	-0.254 *	-0.203	-0.737
	(-1.602)	(-1.361)	(-1.865)	(-0.526)	(-1.196)
2	-6.390	-0.068	-0.173	-0.063	0.264
	(-1.102)	(-0.506)	(-1.122)	(-0.196)	(1.450)
3	-24.493 **	-0.596 **	-0.405	-0.416	-0.704
	(-2.250)	(-2.450)	(-0.956)	(-0.533)	(-1.113)
Big	-669.324 **	-0.870 **	-0.300 **	-0.694 *	0.086
	(-2.354)	(-2.285)	(-2.287)	(-1.664)	(0.276)
Panel B: Implementation strategy using a long-time window.					
Small	1.639	0.267	-0.214	-0.063	-0.388
	(0.482)	(0.663)	(-1.320)	(-0.174)	(-0.560)
2	1.542	0.248	-0.178	0.155	0.544
	(0.300)	(1.451)	(-0.843)	(0.525)	(1.074)
3	31.009	-0.082	0.010	-0.134	-0.220
	(1.199)	(-0.317)	(0.034)	(-0.246)	(-0.377)
Big	-634.297 **	-0.844 **	-0.241 **	-0.631 *	0.178
	(-2.343)	(-2.449)	(-2.114)	(-1.664)	(0.598)

Appendix A

In this appendix we describe the IPO datasets in greater detail in order to facilitate the replication and extension of our study.

IPO Data

We obtain data on IPOs in the Polish stock market from the following providers: GPW (*gpw.pl* and *newconnect.pl*), Stooq (*stooq.com*), Bossa (*bossa.pl*) and Bloomberg.

GPW data available at *gpw.pl*, the official website of the Warsaw Stock Exchange ('WSE'), is the publicly available source of information about securities listed in the WSE (*gpw.pl*). Similarly, the GPW also provides the data about securities listed in the New Connect ('NC') equity market. (*newconnect.pl*). GPW data includes the complete list of IPOs for both markets as well as the summary of the most relevant IPO data. The IPO data available directly from GPW are IPO date, IPO size, and IPO offer price. In rare cases (<1% of the sample), where the GPW provides two offer prices for the same IPO corresponding to different share series, we verify IPO prospectuses individually. We only use GPW's prices to calculate first-day returns but not to calculate IPO size, which is provided by GPW separately.

The IPO data from GPW also includes the transitions from the NC to the WSE market. In cases where the NC company opts for a direct listing in the WSE without new issuance, the GPW assigns it the IPO size of 0, and we retain this treatment as appropriate.

Stock Prices

We retrieve the adjusted daily stock prices from the Stooq database. Adjustments account for stock splits, dividends, and other distributions, rights offerings, and denominations. This data is publicly available at *stooq.com*. The same data was previously used in the studies of the Polish IPO market (e.g. Jewartowski & Lizińska, 2012). We note that in the Stooq database, IPO price definition is not consistent throughout the sample: sometimes first observation is not the IPO price, but a first-day close price. We use IPO offer prices retrieved from the GPW database to verify the starting data point in the Stooq database. We retrieve non-adjusted (raw) daily stock prices from the Bossa database. This is publicly available data provided by Brokerage Unit of Bank Ochrony Środowiska S.A. Bossa provides time-series data in the textual files, separately for each security. The data is publicly available at *bossa.pl*.

Bloomberg is used to obtaining any missing time series data, particularly the daily adjusted and non-adjusted time series of the stock prices for delisted, acquired, and liquidated companies.

In our dataset, there are 178 companies with a market status other than ‘currently active’, including 58 delisted companies in the WSE and 120 delisted companies in the NC. We do not rely on Bloomberg as our main data source because Bloomberg’s coverage of Polish IPOs starts with a few days delay relative to the actual IPO date and is not backfilled. As a result, early data is missing. Moreover, not all IPOs (especially from the NC) are included in Bloomberg’s data.

Financial Statements Data

We obtain financial statements data from Notoria Serwis S.A. (“Notoria”). The access to the Notoria database is not public, although the financial statements are available from companies’ websites. Notoria Financials offers high-quality financial statements data in the standardized format. The database also includes data from the reporting periods prior to the IPO (as reported by the company).

Moreover, we use Notoria to access header information such as the company’s name, trading ticker, or the International Security Identifier Number (ISIN). We manually merge IPO data with financial statements data using the company name. Historical names for the companies that experienced name change are stored in the underlying Notoria files, enabling us to ensure the accuracy of matches. We then use company trading ticker and ISIN to merge the remaining datasets.

Other Data

We use the Dealogic database to access the information about IPO’s Venture-Capital backing status.

We use a one-month Warsaw Interbank Offer Rate (WIBOR) as a proxy for the risk-free rate. We obtain WIBOR quotes from *stooq.com*. The rate is available with ticker PLOPLN1M.

We use WIG and NCIndex indices to track the performance of stocks in the main and the NC markets. We obtain WIG and NCIndex quotes from *stooq.com*. We use index levels to calculate returns, defined as a percentage change in the index value. The quotes are available with tickers WIG and NCIndex, respectively.

We use sector betas provided by Aswath Damodaran. The data is publicly available via Damodaran’s NYU website. We use the July 1, 2017 update of the database, which corresponds to the end of our sample period. We use a leveraged beta database for Europe, which includes Polish stocks. We link companies to the industries using the industry details file, available in *Data: Breakdown* section of Damodaran’s website.

Links

WSE IPOs list – https://www.gpw.pl/debiuty

NC IPOs list – https://newconnect.pl/debiuty-spolek

Adjusted prices – https://stooq.com/

Unadjusted prices – https://info.bossa.pl/notowania/metastock/

Sector betas – http://pages.stern.nyu.edu/~adamodar/

Appendix B

In this Appendix we provide times-series and cross-section of first-day IPO returns in the Polish stock market.

Time-Series of First-Day Returns

Table A2 shows the summary statistics for market-adjusted returns for companies listed on the WSE (Panel A) and NC (Panel B) for each year from January 2005 to June 2017. We note that the NC first-day returns are higher than the corresponding WSE returns, except in the year 2007, when the NC market launched and when average IPO first-day returns were plummeting. Loughran and Ritter (2004) suggest that general market conditions explain first-day returns. They argue that IPO underpricing effect is sample- and period-specific, and report that in the U.S., IPOs were underpriced by an average of 18% in the 1990s, 65% during the Internet bubble (1999–2000), while only 12% in the following years 2000–2003.

[Insert Table A2. here]

Cross-Section of First-Day Returns

Table A3 shows the summary statistics for the market-adjusted first-day returns for companies listed in the WSE (Panel A) and NC (Panel B), broken down into 11 GICS sectors.

[Insert Table A3. here]

Financials, Energy, and Healthcare sectors experience the highest degree of IPO underpricing. The average first-day market-adjusted return equals 17.67% for companies in the financial sector, 17.17% for those in the energy sector and 14.79% for those in the healthcare. When size weights are used, the returns are 13.37%, 20% and 18.23%, respectively. The high first-day returns for the energy and healthcare sector might be linked to the demand argument. According to the theory of market rotation proposed by Bodie et al. (2014), investors can see when an economy is at its peak and anticipate contraction or recession (which was likely, given the level of uncertainty during the global financial crisis). Thus, they tend to invest in noncyclical industries such as energy or healthcare.

The NC differs significantly from the WSE in terms of the cross-section of IPO returns. First, most of the estimates are statistically significant. Second, information technology (IT), materials and telecommunication are the most prominent sectors yielding the highest first-day returns in the NC market. This is not surprising given the NC profile. IT stocks earn the highest equally weighted market-adjusted first-day return, 79.73% (as compared to the IPO-size-weighted market-adjusted return, 45.93%). We relate this outperformance to the divergence of opinions. Many promising small firms take advantage of the low barrier to entry into the NC market and the fact that investing in technological start-ups is in fashion. However, there is little pre-IPO information on the performance of start-up companies, which causes high divergence of opinions and eventually results in higher underpricing, as stated by Miller (1977). Third, real estate and industrial companies are the least underpriced during the IPO process in the NC market.

Table B1. Summary statistics for market-adjusted first-day returns in the WSE and NC for each year from January 2005 to June 2017.

Returns are market-adjusted against the expected return determined by CAPM for each observation. We use Warsaw Interbank Offer Rate (WIBOR) rates as a proxy for the risk-free rate. We use main index, the Warsaw Stock Exchange Index (WIG), as a proxy for market portfolio. Our estimated betas are based on levered industry betas from Damodaran database.

Sample statistic	2005	2006	2007	2008	2009	2010	2011	2012	2013	2014	2015	2016	2017 [a]
Panel A: Summary statistics for companies listed in the WSE.													
No. companies	33	38	79	32	13	34	34	18	23	27	29	19	6
EW average	0.0796	0.3476	0.1498	0.1253	0.1730	0.0683	0.0057	0.0354	0.0414	0.0215	0.0319	0.0675	-0.0107
SW average [b]	0.1345	0.1681	0.0985	0.0153	0.1441	0.0891	0.0082	0.0040	0.0598	0.0278	0.0481	0.0587	0.0659
Median	0.0385	0.1083	0.0653	0.0103	0.1384	0.0231	0.0130	0.0066	0.0022	0.0198	0.0082	0.0439	0.0111
Minimum	-0.1007	-0.0210	-0.1773	-0.7286	-0.0410	-0.1497	-0.2913	-0.6237	-0.3107	-0.1181	-0.1286	-0.0522	-0.2283
Maximum	0.6123	4.7890	1.6211	2.4478	0.5283	1.0703	0.2767	0.6665	0.2646	0.1251	0.7016	0.4365	0.1071
1st quartile	0.0091	0.0266	-0.0203	-0.0665	0.0527	-0.0117	-0.0069	-0.0105	-0.0335	0.0018	-0.0504	-0.0133	-0.0259
3rd quartile	0.1480	0.2878	0.2325	0.1020	0.2466	0.1355	0.0540	0.0647	0.1723	0.0505	0.0591	0.0946	0.0575
Std dev	0.1435	0.8068	0.2764	0.5286	0.1629	0.2013	0.1100	0.2444	0.1367	0.0518	0.1475	0.1209	0.1178
t-statistic	3.1866	2.6556	4.8175	1.3410	3.8307	1.9801	0.3015	0.6152	1.4522	2.1566	1.1662	2.4321	-0.2228
p-value	0.0032	0.0116	0.0000	0.1897	0.0024	0.0561	0.7649	0.5466	0.1605	0.0405	0.2534	0.0257	0.8325
Panel B: Summary statistics for companies listed in the NC.													
No. companies			18	48	18	78	150	82	36	19	17	14	5
EW average			-0.2798	0.6159	0.9543	0.3295	0.2190	0.5821	0.2828	0.3938	0.4204	0.4559	0.5928
SW average [b]			-0.1031	0.2268	0.3028	0.2933	0.1199	0.4329	0.3506	0.5258	0.6761	0.3985	0.2294
Median			-0.2872	0.0711	0.3033	0.2499	0.1332	0.2018	0.2101	0.2959	0.1125	0.4051	0.2988
Minimum			-0.9295	-0.3209	-0.6045	-0.9156	-0.6894	-0.5686	-0.3539	-0.2008	-0.3982	-0.4310	0.0957
Maximum			0.2547	6.3110	9.0065	2.0977	1.2454	28.9923	1.2335	1.2909	1.4389	1.2515	1.2338
1st quartile			-0.5907	-0.0634	0.0360	0.0041	-0.0528	0.0029	0.0150	0.0277	0.0227	0.2370	0.1488
3rd quartile			0.0200	0.5985	1.1329	0.5725	0.4971	0.4936	0.5142	0.5393	0.8606	0.6790	1.1871
Std dev			0.3670	1.3426	2.1238	0.4466	0.3625	3.1948	0.3718	0.4689	0.5621	0.4172	0.5689
t-statistic			-3.2339	3.1784	1.9063	6.5158	7.3975	1.6499	4.5642	3.6607	3.0839	4.0888	2.3300
p-value			0.0049	0.0026	0.0737	0.0000	0.0000	0.1028	0.0001	0.0018	0.0071	0.0013	0.0803

[a] As of June 30, 2017.

[b] IPO size weighting with IPO size winsorized at 0.5% and 99.5%.

Table B2. Summary statistics for market-adjusted first-day returns in the WSE and NC, sorted by sector (aggregated industry).

Returns are market-adjusted against the expected return determined by CAPM for each observation. We use Warsaw Interbank Offer Rate (WIBOR) rates as a proxy for the risk-free rate. We use the main index, the Warsaw Stock Exchange Index (WIG), as a proxy for the market portfolio. Our estimated betas are based on levered industry betas from Damodaran database. For the purpose of market adjustment (beta adjustment), we assign each company to one of 94 sectors. For convenience, here we aggregate the industries into 11 sectors using the GICS.

Sample statistic	Consumer discret.	Consumer staples	Energy	Financials	Health care	Industrials	IT	Materials	Real estate	Telecom.	Utilities
Panel A: Summary statistics for companies listed in the WSE.											
No. companies	80	37	11	43	24	80	41	25	28	5	11
EW average	0.0947	0.0413	0.1717	0.1767	0.1479	0.0937	0.1225	0.0063	0.1809	0.1174	0.0271
SW average [a]	0.0860	0.0569	0.2000	0.1337	0.1823	0.1195	0.1253	0.0048	0.0436	0.1410	0.0424
Median	0.0437	0.0439	0.0290	0.0314	0.0642	0.0309	0.0272	0.0050	0.0489	0.1430	0.0140
Minimum	-0.1773	-0.7286	-0.2913	-0.1497	-0.0634	-0.2253	-0.1286	-0.2847	-0.6237	-0.0573	-0.0693
Maximum	0.9892	0.3768	1.6211	4.7890	1.3252	1.4650	0.7096	0.2767	2.4478	0.3430	0.1384
1st quartile	-0.0070	-0.0124	0.0016	-0.0087	0.0105	-0.0053	-0.0460	-0.0311	-0.0060	-0.0056	-0.0128
3rd quartile	0.1148	0.1174	0.2117	0.1058	0.1970	0.1577	0.1607	0.0397	0.1559	0.1637	0.0659
Std dev	0.2072	0.1644	0.5114	0.7400	0.2745	0.2292	0.2473	0.1076	0.5327	0.1576	0.0636
t-statistic	4.0861	1.5271	1.1132	1.5659	2.6391	3.6562	3.1726	0.2932	1.7965	1.6652	1.4123
p-value	0.0001	0.1355	0.2916	0.1249	0.0147	0.0005	0.0029	0.7719	0.0836	0.1712	0.1882
Panel B: Summary statistics for companies listed in the NC.											
No. companies	112	21	2	62	30	118	78	16	22	14	10
EW average	0.3576	0.2337	0.1545	0.3421	0.4192	0.1941	0.7973	0.4222	0.1742	0.3659	0.2296
SW average [a]	0.2612	0.1958	0.2085	0.1517	0.1699	0.1617	0.4593	0.4250	0.0698	0.4075	0.2571
Median	0.1754	0.2262	0.1545	0.1522	0.0729	0.0871	0.3540	0.3782	0.1668	0.3210	0.1610
Minimum	-0.6638	-0.8159	0.0682	-0.6216	-0.4310	-0.9156	-0.9295	-0.6894	-0.3680	-0.3332	-0.0973
Maximum	3.4600	0.8293	0.2408	6.3110	9.0065	1.7978	28.9923	1.2909	1.0036	1.2483	0.7456
1st quartile	-0.0012	0.0592	0.1113	0.0007	-0.0773	-0.0523	0.0140	0.2023	-0.0068	0.0981	-0.0538
3rd quartile	0.5873	0.5060	0.1976	0.4965	0.2960	0.4183	0.6899	0.7134	0.3177	0.5277	0.4784
Std dev	0.6180	0.3930	0.1220	0.8567	1.6563	0.4260	3.3177	0.4833	0.3273	0.4382	0.3127
t-statistic	6.1228	2.7249	1.7905	3.1442	1.3863	4.9485	2.1224	3.4941	2.4962	3.1240	2.3222
p-value	0.0000	0.0130	0.3243	0.0026	0.1762	0.0000	0.0370	0.0033	0.0209	0.0081	0.0453

[a] IPO size weighting with IPO size winsorized at 0.5% and 99.5%.

Appendix C

In this Appendix, we provide additional analyses and robustness checks, as referenced throughout the paper.

Figure C1. Sample composition by industry and year.
Figure C1 shows the composition of the sample IPOs by industry and year. Our sample extends from January 2008 to June 2017. The sample size is 706 IPOs, including 237 IPOs in the WSE and 469 IPOs in the NC.

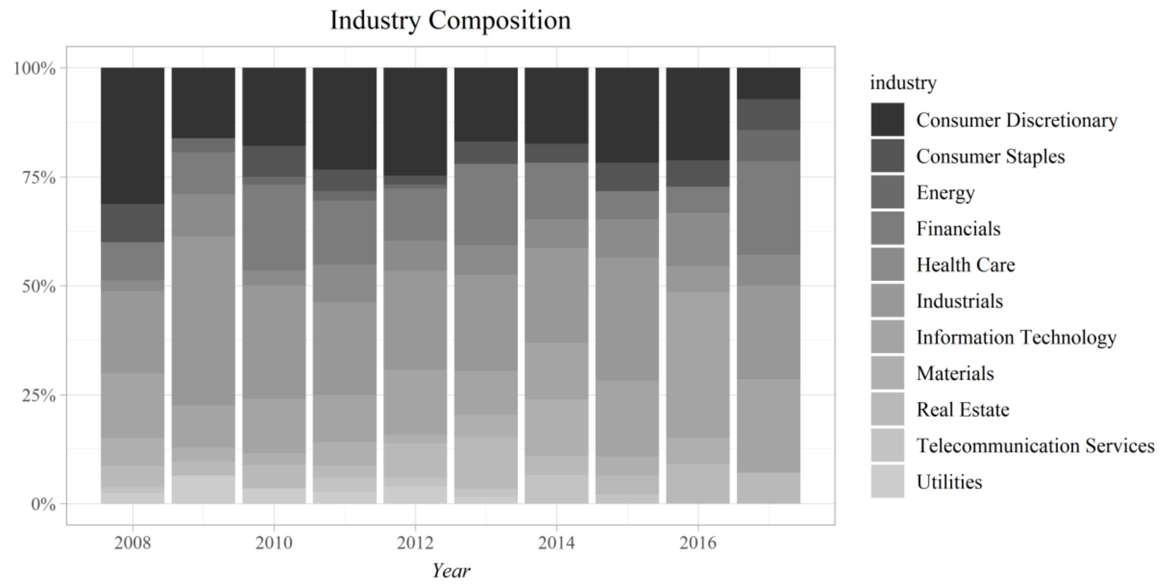
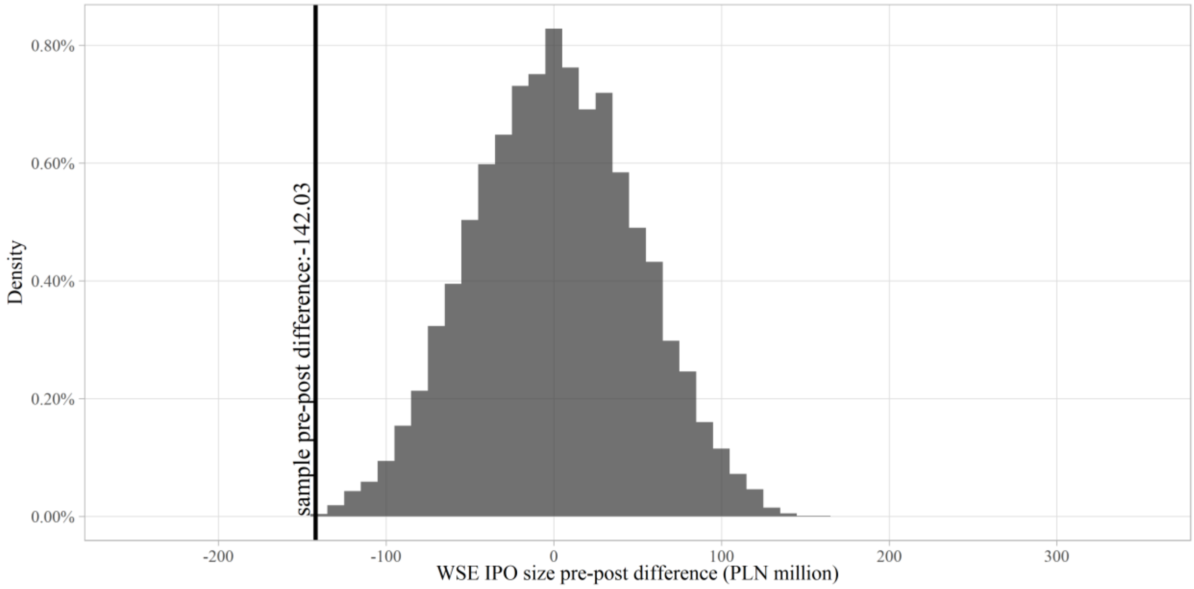


Figure C2. Pre-post differences in IPO size in the WSE and NC from the permutation test.
Figure C2 shows the distribution of pre-post differences in IPO size for the WSE (Panel A) and NC (Panel B) based on the permutation test. The IPO size is defined as the number of shares issued multiplied by the IPO price. The IPO size is reported in million PLN. For the purpose of permutations, we sample IPOs without replacement and calculate the difference between mean IPO size in the post-treatment period and mean IPO size in the pre-treatment period. We use February 3, 2014, as our cut-off date. We use a symmetric 3-year period around the cut-off in the IPO sampling procedure (the short-time window). We plot the density based on 10,000 sampling iterations. The original sample differences are illustrated with the solid vertical lines.

Panel A (WSE)



Panel B (NC)

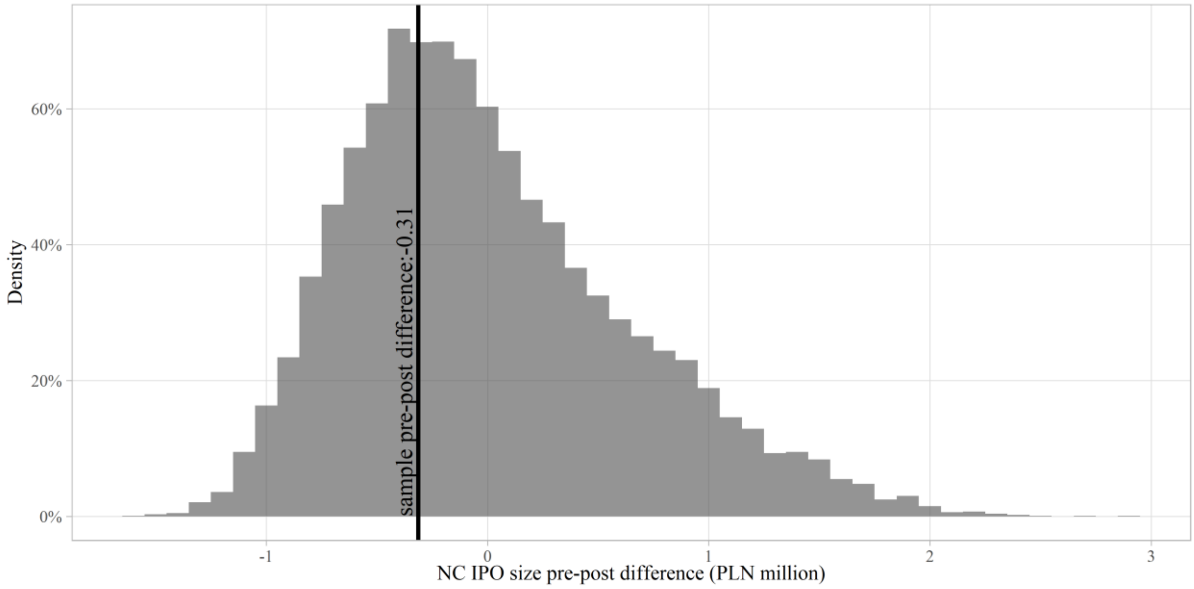


Table C1. Expanded Summary statistics for IPOs listed in the WSE and NC from January 2008 to June 2017.

Panel A contains descriptive statistics regarding IPO returns in the WSE, and Panel B contains the statistics for IPO returns in the NC. The sample for the first-day return calculations includes 385 (486) companies for the WSE (NC). For 30-day and 360-day returns, the WSE (NC) sample size is 347 (395) and 332 (382), respectively. Unadjusted returns are calculated using the HPR, whereas market-adjusted returns are computed using the CAR methodology. The average returns are equally weighted (EW average) or IPO-size-weighted (SW average). IPO size is the number of shares issued at an IPO multiplied by the IPO offer price. All returns are in decimal format; IPO size is in million PLN.

Sample statistic	IPO Size	1st day return	1st day market-adjusted return	30-day return	30-day market-adjusted return	360-day return	360-day market-adjusted return
Panel A: Summary statistics for companies listed in the WSE.							
EW Avg.	215.017	0.058	0.056	0.018	0.017	0.038	0.012
VW Avg.		0.055	0.056	0.049	0.048	0.121	0.042
Std. Dev.	827.895	0.241	0.242	0.205	0.211	0.574	0.524
Min	0.000	-0.741	-0.729	-0.716	-0.672	-0.899	-1.130
P1	0.000	-0.287	-0.304	-0.496	-0.517	-0.815	-0.912
P5	0.000	-0.137	-0.132	-0.294	-0.293	-0.697	-0.685
P10	0.000	-0.070	-0.074	-0.200	-0.212	-0.582	-0.561
P25	0.000	-0.023	-0.023	-0.076	-0.085	-0.293	-0.325
P50	26.000	0.019	0.019	0.004	0.007	0.000	-0.027
P75	96.300	0.084	0.081	0.104	0.117	0.235	0.227
P90	294.402	0.208	0.211	0.246	0.245	0.606	0.563
P95	693.232	0.282	0.268	0.333	0.349	0.926	0.781
P99	4,954.033	0.947	0.945	0.681	0.735	1.871	1.623
Max	8,068.543	2.420	2.448	0.783	0.912	3.451	3.250
N	237	235	235	222	222	207	207
Panel B: Summary statistics for companies listed in the NC.							
EW Avg.	2.609	0.400	0.401	0.431	0.430	0.126	0.072
VW Avg.	0.000	0.254	0.254	0.216	0.213	-0.075	-0.108
Std. Dev.	5.721	1.505	1.506	2.976	2.979	1.845	1.828
Min	0.000	-0.906	-0.916	-0.908	-1.016	-0.973	-1.373
P1	0.000	-0.463	-0.455	-0.773	-0.749	-0.934	-1.159
P5	0.102	-0.269	-0.269	-0.568	-0.582	-0.858	-0.987
P10	0.260	-0.157	-0.153	-0.423	-0.406	-0.782	-0.809
P25	0.571	0.000	0.000	-0.200	-0.192	-0.574	-0.587
P50	1.090	0.185	0.187	0.006	0.013	-0.200	-0.253
P75	2.580	0.518	0.512	0.343	0.351	0.169	0.167
P90	5.567	0.862	0.867	1.162	1.163	1.021	0.867
P95	9.558	1.197	1.195	2.033	2.076	1.752	1.663
P99	19.778	3.237	3.234	5.346	5.282	7.991	7.750
Max	85.138	29.000	28.992	51.000	51.010	25.331	25.132
N	469	468	467	382	381	370	369

[a] IPO size weighting with IPO size winsorized at 0.5% and 99.5%.

Table C2. Two-sample nonparametric tests of pre-post treatment differences in the mean IPO size.

The table presents the results of the nonparametric tests of pre-post treatment differences in the mean IPO size for the WSE and NC listings. In Panel A and B, we show the results using a short-time window implementation strategy (3-year symmetric windows; the period of February 4, 2011, till February 3, 2017) and a long-time window implementation strategy (January 1, 2008-July 30, 2017) respectively. We use February 3, 2014, as our cut-off date. To estimate *p*-values in the bootstrapping (permutations) test, we sample IPOs with (without) replacement and calculate the difference between mean IPO size in the post-treatment period and mean IPO size in the pre-treatment period. We estimate *p*-values as the frequency of differences in the sampled data that fall below the observed sample difference. Therefore, low *p*-values imply a high statistical significance of the observed differences relative to the random benchmark. We use a symmetric 3-year period around the cut-off in the IPO sampling procedure (the short-time window). We base our estimates on the 10,000 sampling iterations. Returns are market-adjusted and cumulated over the relevant period as CARs and are reported in percentages. The IPO size equals the IPO offer price multiplied by the number of shares issued and is reported in million PLN. The *t*-statistics corresponding to the estimated p-values are reported in parentheses.

	IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.				
WSE				
Bootstrapping	0.019	0.706	0.499	0.860
	(2.073)	(-0.541)	(0.002)	(-1.079)
Permutations	0.000	0.782	0.494	0.930
	(3.433)	(-0.778)	(0.014)	(-1.474)
NC				
Bootstrapping	0.369	0.841	0.873	0.897
	(0.335)	(-0.999)	(-1.139)	(-1.265)
Permutations	0.347	0.838	0.859	0.904
	(0.393)	(-0.984)	(-1.077)	(-1.304)
Panel B: Implementation strategy using a long-time window.				
WSE				
Bootstrapping	0.041	0.144	0.427	0.796
	(1.742)	(1.064)	(0.185)	(-0.826)
Permutations	0.006	0.072	0.412	0.861
	(2.531)	(1.462)	(0.222)	(-1.083)
NC				
Bootstrapping	0.548	0.756	0.888	0.741
	(-0.121)	(-0.695)	(-1.215)	(-0.647)
Permutations	0.566	0.790	0.880	0.751
	(-0.165)	(-0.806)	(-1.173)	(-0.678)

Table C3. Two-sample *t*-test of pre- and post-treatment mean unadjusted returns.

The table presents pre- and post-treatment mean values of the first-day, 30-day, and 360-day unadjusted (raw) returns for WSE and NC listings. In Panel A and B, we show the results using a short-time window implementation strategy (3-year symmetric windows; the period of February 4, 2011, till February 3, 2017) and a long-time window implementation strategy (January 1, 2008-July 30, 2017) respectively. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. For reference, we report the number of observations in the last row of each panel. Returns are cumulated over the relevant period as CARs and are reported in percentages.

	1st day return	30-day return	360-day return
Panel A: Implementation strategy using a short-time window.			
WSE			
Pre-mean	0.026	0.033	0.034
Post-mean	0.038	0.018	0.083
Difference	0.012	-0.015	0.049
t-statistic	(-0.542)	(0.467)	(-0.511)
Obs.	148	143	132
NC			
Pre-mean	0.325	0.314	-0.012
Post-mean	0.448	1.057	0.199
Difference	0.123	0.743	0.212
t-statistic	(-0.916)	(-1.312)	(-0.745)
Obs.	306	258	247
Panel B: Implementation strategy using a long-time window.			
WSE			
Pre-mean	0.069	0.021	0.020
Post-mean	0.035	0.010	0.083
Difference	-0.035	-0.011	0.063
t-statistic	(1.324)	(0.400)	(-0.803)
Obs.	235	222	207
NC			
Pre-mean	0.393	0.349	0.119
Post-mean	0.452	1.027	0.199
Difference	0.059	0.678	0.081
t-statistic	(-0.576)	(-1.290)	(-0.279)
Obs.	468	382	370

Table C4. Two-sample *t*-test of pre- and post-treatment mean returns and IPO sizes – alternative sample excluding 2008-2009 Financial Crisis.

The table presents pre- and post-treatment mean values of the IPO size, the first-day, 30-day, and 360-day market-adjusted returns for the WSE and NC listings. In Panel A and B, we show the results using a short-time window implementation strategy (3-year symmetric windows; the period of February 4, 2011, till February 3, 2017) and a long-time window implementation strategy (January 1, 2010-July 30, 2017) respectively. The sample excludes IPOs from 2008 and 2009 corresponding to the global financial crisis. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. For reference, we report the number of observations in the last row of each panel. Returns are market-adjusted and cumulated over the relevant period as CARs and are reported in percentages. The IPO size equals the IPO offer price multiplied by the number of shares issued and is reported in million PLN.

	IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.				
WSE				
Pre-mean	201.248	0.023	0.023	-0.041
Post-mean	59.220	0.035	0.022	0.060
Difference	-142.028 **	0.012	0.000	0.101
t-statistic	(2.133)	(-0.548)	(0.005)	(-1.086)
Obs.	148	148	143	132
NC				
Pre-mean	2.586	0.326	0.318	-0.106
Post-mean	2.273	0.449	1.050	0.194
Difference	-0.313	0.123	0.732	0.299
t-statistic	(0.614)	(-0.916)	(-1.292)	(-1.071)
Obs.	306	305	257	246
Panel B: Implementation strategy using a long-time window.				
WSE				
Pre-mean	233.175	0.038	0.017	-0.067
Post-mean	92.224	0.033	0.013	0.060
Difference	-140.951 **	-0.006	-0.004	0.127
t-statistic	(2.080)	(0.277)	(0.140)	(-1.645)
Obs.	192	190	181	166
NC				
Pre-mean	2.679	0.335	0.336	0.013
Post-mean	2.553	0.454	1.019	0.194
Difference	-0.126	0.119	0.683	0.180
t-statistic	(0.230)	(-1.098)	(-1.276)	(-0.622)
Obs.	403	401	326	314

Table C5. Effects of the 2014 regulatory change on the IPO size – winsorization choices.

The table reports results for the regression $Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \gamma_t + \delta_j + \varepsilon_i$. The dependent variable is the IPO size (in million PLN), and our unit of observation is a single IPO. Different columns correspond to different winsorization choices. We winsorize IPO size at 0% (no winsorization) at 0.5% (0.5% and 99.5%) and at 1% (1% and 99%). WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the NC IPO listings. $Post$ is an indicator variable that takes on a value of 1 for post-event IPO listings, and 0 for the pre-event IPO listings, $WSE \times Post$ is the variable of interest that captures the difference-in-differences effect. All regressions include industry and year fixed effects. Year fixed effects indicate the year of the IPO. t-statistics are reported in parentheses. Data are for the period February 4, 2011-February 3, 2017 for Panel A, and January 1, 2008-July 30, 2017 for Panel B. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. For reference, we report the number of observations in the last row of each panel. IPO size is in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

	IPO Size (winsorized at 0%)	IPO Size (winsorized at 0.5%)	IPO Size (winsorized at 1%)
Panel A: Implementation strategy using a short-time window.			
WSE	209.168 **	184.950 ***	147.984 ***
	(2.412)	(2.797)	(3.457)
WSE x Post	-151.769 *	-127.814 *	-90.945 **
	(-1.740)	(-1.895)	(-2.029)
Industry FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
R2	0.114	0.136	0.169
Obs.	454	454	454
Panel B: Implementation strategy using a long-time window.			
WSE	247.566 ***	194.113 ***	144.526 ***
	(3.271)	(4.040)	(4.910)
WSE x Post	-163.203 **	-106.672 **	-56.522
	(-2.007)	(-1.968)	(-1.468)
Industry FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
R2	0.129	0.180	0.210
Obs.	706	706	706

Table C6. Effects of the 2014 regulatory change on the IPO underpricing, long-term underperformance and on the IPO size – alternative regression specification with Post variable.

The table reports results for the regression $Y_i = \beta_1 WSE_i + \beta_2 Post_i + \beta_3 WSE_i \times Post_i + \delta_j + \varepsilon_i$. The dependent variable is either: the IPO size (in million PLN), a first-day market-adjusted return, a 30-day market-adjusted return, or a 360-day market-adjusted return (all in percentages), and our unit of observation is a single IPO. WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the NC IPO listings. $Post$ is an indicator variable that takes on a value of 1 for post-event IPO listings, and 0 for the pre-event IPO listings, $WSE \times Post$ is the variable of interest that captures the difference-in-differences effect. All regressions include industry fixed effects. t-statistics are reported in parentheses. Data are for the period February 4, 2011-February 3, 2017 for Panel A, and January 1, 2008-July 30, 2017 for Panel B. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. For reference, we report the number of observations in the last row of each panel. Returns are in decimal format; IPO size is in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. We winsorize the observed IPO sizes at 0.5% and 99.5%.

	IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.				
WSE	187.370 ***	-0.216 ***	-0.163	0.098
	(2.904)	(-3.364)	(-1.209)	(0.849)
Post	1.341	0.080	0.678	0.256
	(0.139)	(0.513)	(1.156)	(0.931)
WSE x Post	-133.946 *	-0.189 *	-0.878	-0.216
	(-1.961)	(-1.697)	(-1.560)	(-0.757)
Industry FE	Yes	Yes	Yes	Yes
Year FE	No	No	No	No
R2	0.134	0.032	0.024	0.040
Obs.	454	453	400	378
Panel B: Implementation strategy using a long-time window.				
WSE	194.675 ***	-0.280 ***	-0.258 **	-0.066
	(4.082)	(-4.401)	(-2.450)	(-0.499)
Post	18.514	0.041	0.637	0.140
	(1.631)	(0.357)	(1.197)	(0.501)
WSE x Post	-112.248 **	-0.151	-0.747	-0.094
	(-1.978)	(-1.551)	(-1.430)	(-0.335)
Industry FE	Yes	Yes	Yes	Yes
Year FE	No	No	No	No
R2	0.171	0.032	0.020	0.011
Obs.	706	702	603	576

Table C7. Effects of the 2014 regulatory change on the IPO underpricing, long-term underperformance and on the IPO size – alternative sample excluding Consumer Discretionary sector.

The table reports results for the regression $Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \gamma_t + \delta_j + \varepsilon_i$. The dependent variable is either: the IPO size (in million PLN), the first-day market-adjusted return, the 30-day market-adjusted return, or the 360-day market-adjusted return (all in percentages), and our unit of observation is a single IPO. WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the NC IPO listings. $Post$ is an indicator variable that takes on a value of 1 for post-event IPO listings, and 0 for the pre-event IPO listings, $WSE \times Post$ is the variable of interest that captures the difference-in-differences effect. All regressions include industry and year fixed effects. Year fixed effects indicate the year of the IPO. t-statistics are reported in parentheses. Data are for the period February 4, 2011-February 3, 2017 for Panel A, and January 1, 2008-July 30, 2017 for Panel B. The sample excludes IPOs from the Consumer Discretionary sector due to the observed decline in the number of IPOs in that sector and its large contribution to the full sample. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. For reference, we report the number of observations in the last row of each panel. Returns are in decimal format; IPO size is in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. We winsorize the observed IPO sizes at 0.5% and 99.5%.

	IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.				
WSE	217.085 ***	-0.212 ***	-0.143	0.144
	(2.694)	(-3.697)	(-1.136)	(1.266)
WSE x Post	-157.819 *	-0.207 **	-0.362 *	-0.304
	(-1.909)	(-2.194)	(-1.756)	(-1.177)
Industry FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R2	0.142	0.042	0.028	0.067
Obs.	353	353	316	298
Panel B: Implementation strategy using a long-time window.				
WSE	226.963 ***	-0.301 ***	-0.262 ***	-0.043
	(3.825)	(-4.271)	(-2.902)	(-0.378)
WSE x Post	-133.005 **	-0.159 *	-0.258	-0.141
	(-2.044)	(-1.689)	(-1.296)	(-0.562)
Industry FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R2	0.185	0.043	0.025	0.054
Obs.	552	549	476	453

Table C8. Effects of the 2014 regulatory change on the IPO underpricing, long-term underperformance and on the IPO size – alternative sample excluding 2008-2009 Financial Crisis.

The table reports results for the regression $Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \gamma_t + \delta_j + \varepsilon_i$. The dependent variable is either: the IPO size (in million PLN), the first-day market-adjusted return, the 30-day market-adjusted return, or the 360-day market-adjusted return (all in percentages), and our unit of observation is a single IPO. WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the NC IPO listings. $Post$ is an indicator variable that takes on a value of 1 for post-event IPO listings, and 0 for the pre-event IPO listings, $WSE \times Post$ is the variable of interest that captures the difference-in-differences effect. All regressions include industry and year fixed effects. Year fixed effects indicate the year of the IPO. t-statistics are reported in parentheses. Data are for the period February 4, 2011-February 3, 2017 for Panel A, and January 1, 2010-July 30, 2017 for Panel B. The sample excludes IPOs from the 2008 and 2009 corresponding to the Global Financial Crisis. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. For reference, we report the number of observations in the last row of each panel. Returns are in decimal format; IPO size is in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. We winsorize the observed IPO sizes at 0.5% and 99.5%.

	IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.				
WSE	184.950 ***	-0.213 ***	-0.154	0.077
	(2.797)	(-4.367)	(-1.433)	(0.702)
WSE x Post	-127.814 *	-0.174 **	-0.872	-0.219
	(-1.895)	(-2.060)	(-1.620)	(-0.870)
Industry FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R2	0.136	0.040	0.032	0.060
Obs.	454	453	400	378
Panel B: Implementation strategy using a long-time window.				
WSE	208.709 ***	-0.232 ***	-0.235 ***	-0.122
	(3.598)	(-6.349)	(-2.723)	(-0.719)
WSE x Post	-120.096 *	-0.172 **	-0.782	-0.087
	(-1.904)	(-2.403)	(-1.551)	(-0.320)
Industry FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R2	0.142	0.039	0.030	0.028
Obs.	595	591	507	480

Table C9. Effects of the 2014 regulatory change on the IPO underpricing, long-term underperformance and on the IPO size – alternative threshold for the long-term return variables.

The table reports results for the regression $Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \gamma_t + \delta_j + \varepsilon_i$. The dependent variable is either: the IPO size (in million PLN), the first-day market-adjusted return, the 30-day market-adjusted return, or the 360-day market-adjusted return (all in percentages), and our unit of observation is a single IPO. WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the NC IPO listings. $Post$ is an indicator variable that takes on a value of 1 for post-event IPO listings, and 0 for the pre-event IPO listings, $WSE \times Post$ is the variable of interest that captures the difference-in-differences effect. All regressions include industry and year fixed effects. We adjust the threshold so that the post-intervention period includes all returns that could have been potentially affected, that is, observations of firms that listed during the pre-treatment period but which 30-day and 360-day interval for aftermarket performance measurement is also in the range of the post-treatment period. Year fixed effects indicate the year of the IPO. t-statistics are reported in parentheses. Data are for the period February 4, 2011-February 3, 2017 for Panel A, and January 1, 2008-July 30, 2017 for Panel B. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. For reference, we report the number of observations in the last row of each panel. Returns are in decimal format; IPO size is in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. We winsorize the observed IPO sizes at 0.5% and 99.5%.

	IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.				
WSE	184.950 ***	-0.213 ***	-0.170	-0.174
	(2.797)	(-4.367)	(-1.517)	(-0.788)
WSE x Post	-127.814 *	-0.174 **	-0.809	0.180
	(-1.895)	(-2.060)	(-1.409)	(0.706)
Industry FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R2	0.136	0.040	0.033	0.028
Obs.	454	453	410	462
Panel B: Implementation strategy using a long-time window.				
WSE	194.113 ***	-0.317 ***	-0.267 ***	-0.154
	(4.040)	(-5.382)	(-3.216)	(-0.905)
WSE x Post	-106.672 **	-0.103	-0.727	0.029
	(-1.968)	(-1.251)	(-1.396)	(0.125)
Industry FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R2	0.180	0.044	0.029	0.027
Obs.	706	702	603	576

Table C10. Effects of the 2014 regulatory change on the IPO underpricing, long-term underperformance and on the IPO size – alternative specification with interacted industry and Post effects.

The table reports results for the regression $Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \gamma_t + \delta_j + \delta_j \times Post_i + \varepsilon_i$. The dependent variable is either: the IPO size (in million PLN), the first-day market-adjusted return, the 30-day market-adjusted return, or the 360-day market-adjusted return (all in percentages), and our unit of observation is a single IPO. WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the NC IPO listings. $Post$ is an indicator variable that takes on a value of 1 for post-event IPO listings, and 0 for the pre-event IPO listings, $WSE \times Post$ is the variable of interest that captures the difference-in-differences effect. All regressions include industry and year fixed effects. Industry fixed effects are additionally interacted with the $Post$ variable to control for the industry-specific impact of the regulation. Year fixed effects indicate the year of the IPO. t-statistics are reported in parentheses. Data are for the period February 4, 2011-February 3, 2017 for Panel A, and January 1, 2008-July 30, 2017 for Panel B. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. For reference, we report the number of observations in the last row of each panel. Returns are in decimal format; IPO size is in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. We winsorize the observed IPO sizes at 0.5% and 99.5%.

	IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.				
WSE	184.950 ***	-0.213 ***	-0.154	0.077
	(2.797)	(-4.367)	(-1.433)	(0.702)
WSE x Post	-127.814 *	-0.174 **	-0.872	-0.219
	(-1.895)	(-2.060)	(-1.620)	(-0.870)
Industry FE	Yes	Yes	Yes	Yes
Industry FE x Post	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R2	0.136	0.040	0.032	0.060
Obs.	454	453	400	378
Panel B: Implementation strategy using a long-time window.				
WSE	194.113 ***	-0.317 ***	-0.271 ***	-0.135
	(4.040)	(-5.382)	(-3.309)	(-0.930)
WSE x Post	-106.672 **	-0.103	-0.737	-0.052
	(-1.968)	(-1.251)	(-1.463)	(-0.202)
Industry FE	Yes	Yes	Yes	Yes
Industry FE x Post	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
R2	0.180	0.044	0.029	0.027
Obs.	706	702	603	576

Table C11. Effects of the 2014 regulatory change on the IPO underpricing, long-term underperformance and on the IPO size – alternative regression specification with additional control variables and Post variable.

The table reports results for the regression $Y_i = \beta_1 WSE_i + \beta_2 Post_i + \beta_3 WSE_i \times Post_i + \beta_4 \log{(Assets)_i} + \beta_5 ROA_i + \beta_6 I(VC)_i + \gamma_t + \delta_j + \varepsilon_i$. The dependent variable is either: the IPO size (in million PLN), the first-day market-adjusted return, the 30-day market-adjusted return, or the 360-day market-adjusted return (all in percentages), and our unit of observation is a single IPO. WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the NC IPO listings. $Post$ is an indicator variable that takes on a value of 1 for post-event IPO listings, and 0 for the pre-event IPO listings, $WSE \times Post$ is the variable of interest that captures the difference-in-differences effect. log $(Assets)$ are the total assets in the last fiscal year prior to the IPO. ROA is the return on assets in the last fiscal year prior to the IPO, using the end-of-period value of total assets in the denominator. I(VC) is an indicator of whether an IPO was Venture Capital backed. All regressions include industry fixed effects. t-statistics are reported in parentheses. Data are for the period February 4, 2011-February 3, 2017 for Panel A, and January 1, 2008-July 30, 2017 for Panel B. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. For reference, we report the number of observations in the last row of each panel. Returns are in decimal format; IPO size is in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. We winsorize the observed IPO sizes at 0.5% and 99.5%.

	IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.				
WSE	35.816	-0.049	0.089	0.093
	(1.530)	(-1.406)	(0.638)	(0.658)
Post	-15.446	0.248 ***	1.013 *	0.291
	(-0.953)	(3.590)	(1.892)	(1.074)
WSE x Post	-98.095 *	-0.286 ***	-1.059 *	-0.221
	(-1.861)	(-3.762)	(-1.789)	(-0.766)
Controls	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Year FE	No	No	No	No
R2	0.193	0.191	0.083	0.054
Obs.	415	414	371	351
Panel B: Implementation strategy using a long-time window.				
WSE	-35.017	-0.115 **	-0.035	0.064
	(-1.126)	(-2.244)	(-0.402)	(0.575)
Post	-7.696	0.176 **	0.892 *	0.270
	(-0.423)	(2.573)	(1.743)	(1.027)
WSE x Post	-68.249	-0.236 ***	-0.915 *	-0.201
	(-1.460)	(-2.995)	(-1.687)	(-0.745)
Controls	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Year FE	No	No	No	No
R2	0.292	0.091	0.070	0.039
Obs.	643	639	556	531

Table C12. Issuer effects of the 2014 regulatory change on the IPO underpricing, long-term underperformance and on the IPO size across different IPO size sample partitions – alternative specification with additional control variables.

The table reports results for the regression $Y_i = \beta_1 WSE_i + \beta_2 WSE_i \times Post_i + \beta_3 \log(Assets)_i + \beta_4 ROA_i + \beta_5 I(VC)_i + \gamma_t + \delta_j + \varepsilon_i$. We estimate the regression for the IPO size partitions. To construct the size partitions, we sort the sample based on IPO size each year. Prior to sorting, we exclude issuers with IPO size unavailable or reported as 0 by the exchange. Small issuers are issuers in the smallest size quartile (bottom 25% issuers). Big issuers are issuers in the largest size quartile (top 20% issuers). We also report intermediate portfolios (labelled portfolio 2 and 3) corresponding to the second and third quartile of IPO size. The dependent variable is either: the IPO size (in million PLN), the first-day market-adjusted return, the 30-day market-adjusted return, or the 360-day market-adjusted return (all in percentages), and our unit of observation is a single IPO. Log IPO size is the natural logarithm of IPO size. In the regression, WSE is an indicator variable that takes on a value of 1 for IPO listings in the WSE market, and 0 for the NC IPO listings. $Post$ is an indicator variable that takes on a value of 1 for post-event IPO listings, and 0 for the pre-event IPO listings, $WSE \times Post$ is the variable of interest that captures the difference-in-differences effect. $\log(Assets)$ are the total assets in the last fiscal year prior to the IPO. ROA is the return on assets in the last fiscal year prior to the IPO, using the end-of-period value of total assets in the denominator. $\mathrm{I(VC)}$ is an indicator of whether an IPO was Venture Capital backed. All regressions include industry and year fixed effects. Year fixed effects indicate the year of the IPO. We only report difference-in-differences estimates, which are the coefficients on $WSE \times Post$. t-statistics are reported in parentheses. Data are for the period February 4, 2011-February 3, 2017 for Panel A, and January 1, 2008-July 30, 2017 for Panel B. The sample size varies depending on the length of the sample window as well as data requirements with respect to the dependent variable. Returns are in decimal format; IPO size is in million PLN. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. We winsorize observed IPO sizes at 0.5% and 99.5%.

	IPO Size	Log IPO Size	1st day market-adjusted return	30-day market-adjusted return	360-day market-adjusted return
Panel A: Implementation strategy using a short-time window.					
Small	-6.608	-0.538	-0.277 *	-0.233	-0.720
2	(-1.521)	(-1.516)	(-1.891)	(-0.630)	(-1.111)
	-4.520	0.028	-0.187	-0.151	0.430
3	(-0.713)	(0.204)	(-1.021)	(-0.395)	(1.526)
	-24.930 **	-0.610 **	-0.189	-0.198	-0.544
Big	(-2.310)	(-2.529)	(-1.150)	(-1.017)	(-0.885)
	-522.823 **	-0.587 *	-0.311 **	-0.685	0.152
Small	(-2.304)	(-1.823)	(-2.094)	(-1.474)	(0.460)
Panel B: Implementation strategy using a long-time window.					
Small	1.395	0.234	-0.223	-0.171	-0.705
	(0.402)	(0.579)	(-1.414)	(-0.481)	(-1.154)
2	2.502	0.277 *	-0.200	0.135	0.411
	(0.491)	(1.865)	(-0.787)	(0.389)	(1.597)
3	29.954	-0.103	-0.049	-0.246 *	-0.343
	(1.122)	(-0.398)	(-0.328)	(-1.715)	(-0.669)
Big	-334.022 *	-0.397	-0.252 **	-0.617	0.186
	(-1.663)	(-1.303)	(-2.068)	(-1.444)	(0.596)