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Transaction costs and volatility on Warsaw Stock Exchange: implications for financial transaction tax

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Abstract

We present the concept of financial transaction tax (Tobin tax, FTT) and describe its potential consequences. We analyse the relation between transaction costs and volatility of prices by presenting empirical evidence from Warsaw Stock Exchange and exploiting the natural experiment of varying tick size. A higher tick size (and thus increased transaction costs) seems to be connected to higher volatility. Since increased transaction costs may be a proxy for the effect of FTT, our findings may be interpreted as an evidence against the stabilizing role of the tax.

Keywords: transaction costs, volatility, Tobin tax, financial transaction tax

JEL: D23, C33, G10, H30

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1. Introduction

The recent period, dominated by the consequences of the financial crisis, has caused a renewal of interest in the proposal of taxing financial transactions. Although this idea goes back to Tobin (1978) and Keynes (1936), it remained a purely theoretical concept for a long time. Although some mechanisms of taxing financial transactions have been functioning in selected countries, a universal tax with a broad geographical and product coverage has never been implemented. An attempt to do so was made after the last financial crisis. The European Commission proposed a project on taxing financial transactions, calling for its implementation in the entire European Union. The tax is meant to improve the functioning of the financial sector, unify the ways of its taxation throughout the European Union and increase its contribution to the public budgets. After some members of EU strongly opposed the initial proposal from 2011, the project was modified to include only 11 member states and is currently discussed at the European level.

Because of the vast number of transactions conducted every day, financial transaction tax (FTT) might significantly change the current picture of the financial system and thus is of great interest to the economists. Existing papers discuss the impact of the tax on the level and volatility of prices, volume of transactions and key macroeconomic variables. Although in some areas conclusions from the literature are consistent (e.g. the impact on trading volume), in others the evidence is mixed (e.g. the impact on volatility). Unfortunately, empirical evidence on the consequences of the tax is very limited and the existing examples often should be interpreted with caution (e.g. because of a significant change in market characteristics), which makes the inference very difficult. Nevertheless, some attempts to estimate the potential effect of FTT are made, often using some proxies for the tax mechanism, such as natural experiments provided by the market structure. Such an approach is also adopted in the empirical part of this paper.

This paper attempts to fill the gap in the Polish literature by presenting the vibrant topic of FTT and to enrich the existing empirical evidence on the impact of the tax on the volatility of prices. While we do not analyse any tax experiment directly, we try to estimate the effect of transaction costs increase on the level of volatility, arguing that this is a reasonable proxy for the effect of the tax. The rest of the paper is structured as follows. In Section 2 we describe the basic concept of FTT and provide a brief review of the existing literature. Section 3 narrows down the area of analysis to the impact of the tax on the volatility of prices, presenting selected theoretical approaches to the problem and the existing empirical evidence. Section 4 tries to enrich this evidence by presenting the analysis of the volatility of stock prices from Warsaw Stock Exchange. We exploit the fact that when the stock price is below 50 PLN, the minimum tick size is 0.01 PLN, whereas when the price exceeds 50 PLN, the minimum tick size rises to 0.05 PLN. This natural experiment is an exogenous increase in the transaction costs and thus may serve as a proxy for the mechanism of FTT. The analysis is conducted in a panel data setting. The work is concluded in the last, fifth part.

2. Review of the idea of FTT

FTT is often referred to as Tobin tax since James Tobin was the first one to present an elaborate proposal of taxing financial transactions (Tobin 1978). According to Tobin, taxing every foreign exchange transaction (the tax was meant to be implemented only in the foreign exchange market) at some small rate should not only collect some revenues, but also reduce the international mobility of capital, preventing investors from an abrupt withdrawal of all their funds from a selected country (as it was later the case in Mexico's 1994 crisis). The tax, therefore, was meant to be a macroeconomic device, stabilizing the capital market and freeing the hands of monetary policymakers.

The idea, however, did not get much attention for over a decade. The topic came back onto the academic agenda at the end of 1980s, with voices coming both from proponents (Stiglitz 1989; Summers, Summers 1989) and opponents (Kupiec 1996) of the tax. In the meantime, however, the idea of the tax has evolved thanks to non-government organizations which looked for a mechanism allowing to transfer capital from the rich financial world to the poor (therefore sporadically used name "Robin Hood tax"). Despite the ongoing debate and actions taken by the tax proponents, the tax has remained a theoretical concept. Although there have existed some examples of taxes on financial transactions, their mechanism is hardly comparable to the global, universal tax understood under the name of FTT.

As an aftermath of the financial crisis and the huge amount of public money received by financial institutions, the idea of taxing the financial sector has become very widely discussed. Following the call from the European Parliament, the European Commission examined potential mechanisms for taxing financial sector and pointed FTT as the best available solution (see EC 2011). The initial project from September 2011 was strongly opposed by some member states (mainly Great Britain) and its implementation (and introducing the tax in the entire European Union in general) has become unlikely. A new proposal was prepared, assuming that only 11 member states would introduce the tax. This project is currently discussed at the European level. Even though European Parliament approved the proposal in July 2013, the initial plan of introducing the tax in the beginning of 2014 failed and the project is still discussed on the European level (see Semeta 2014). In the meantime, France and Italy has introduced a unilateral tax on financial transactions. No published work assessing the French or Italian experience was available at the moment of preparing this article, though.

The tax proposal of the European Commission assumes taxing transactions with all the financial instruments at a small, but instrument-varying tax rate. For equities the rate would be 0.1%, whereas for derivatives 0.01%. Because of the numerous problems with constructing the tax mechanism (e.g. defining the tax base for options), those rates should be treated as a working assumption. While the first project of the tax was based on the residence criterion, i.e. the tax would be paid by all parties with headquarters in EU, the modified project extends the tax coverage to the cases when parties not located in FTT-jurisdiction trade an instrument issued in this jurisdiction. For more technical details on the tax functioning we refer to the proposal and accompanying impact assessment (EC 2011; 2013).

There is no doubt that the imposition of FTT can cause significant changes in the functioning of the financial market. We briefly review the potential consequences below.

2.1. The impact of FTT on the market functioning

Trading volume

Since FTT increases the transaction costs, it should at the same time decrease the number of transactions. There is little doubt about existence of this effect but the degree of the decrease of volume is arguable. Because there is not much theoretical work which can help here, the debate concentrates on the empirical evidence.

The classical paper documenting the effects of transaction tax on the stock trading volume is Umlauf (1993), who examines Swedish experience of introducing (and then modifying and finally abolishing) the transaction tax. The estimated elasticity of the turnover with respect to the tax rate (which was increased from 0.5% to 1%) is -0.6, which means that the increase of the tax by 10% leads to 6% decrease of turnover (i.e. the 100% increase of the tax rate caused the decrease of trading volume by 60%). However, in most European countries there is no tax in place, therefore any positive tax rate would translate into infinite change – the estimate of elasticity with respect to the tax rate is therefore not useful. What would be more interesting, however, is the elasticity with respect to the total transaction costs. This would require the transaction costs level to be estimated (so that we know what is the share of 0.5% tax increase in all transaction costs in Sweden at that time), which is not done by Umlauf. Further works (e.g. Westerholm 2003) estimate that the elasticity with respect to the level of transaction costs at the time of abolishing the Swedish tax lay in the interval (-1.3; -1). The data from Japan (Liu 2007) or China (e.g. Su, Zheng 2011) seems to confirm that the elasticity with respect to the total transaction costs is around -1. The Chinese experience, however, suggests that the elasticity may be asymmetric and the elasticity value for the decrease of the tax rate is estimated to be around -4, which is far from -1 estimated for tax increases. Since most of the available evidence is based on the experiences of tax decreases, it must be treated with caution when interpreted in the context of introducing FTT.

The elasticity on other markets, e.g. futures market, is likely to be higher. Based on the American futures data, Wang and Yau (2000) estimate the elasticity of the volume (with respect to the level of transaction costs) on the futures market to be between -2 and -1.2. Chou and Wang (2006), analysing the Taiwanese data, estimate the elasticity to be around -1.1.

Although the existing estimates should be interpreted with caution (since they are based on the data from the periods in which financial markets were less developed) they altogether suggest that the elasticity of trading volume with respect to the transaction costs lies around -1 in the stock market and may be a bit higher for the derivatives market. This means that if the rate of financial transaction tax constituted 10% of total transaction costs, its introduction would potentially cause 10% fall in the turnover. There are many empirical estimates of transaction costs. McCulloch and Pacillo (2011) provide a review, showing that a median estimate for equity market is slightly above 1 percentage point and is much smaller for foreign exchange market, where it amounts to 0.05 percentage point. Thus, proposed tax rates of 0.1% (equities) and 0.01% (derivatives) amount to ca. 10–25% of transaction costs.

Based on the findings from literature, the European Commission used a working assumption of -1 elasticity for the stock market and -1.5 for the derivatives market (EC 2013). While these assumptions are relatively uncontroversial, they only concern the decrease of trading due to the lower profitability of trading activities. The other possibility is the reduction of the trading activities by moving them to other markets or legally avoiding the tax in some way. While this cannot be done with all assets (e.g. a Chinese stock can only be traded on the Chinese exchange) and thus is often not reflected in the empirical data, it may constitute a significant problem, especially in the derivatives market. The European Commission assumes that the relocation parameter, capturing the fraction of trades shifted to other markets, is equal to 15% for securities and 75% for derivatives.[1] Assuming a high degree of relocation in the derivatives market is very reasonable, however there is no credible evidence which can help to determine the exact number – there is as much reason to assume 75% as to assume 95% or 60%.

Umlauf (1993) contains some evidence on the relocation in Sweden (where the trade moved to London), however it can hardly be representative. He presents official estimates from that time, which suggest that around 60% of the turnover of largest companies (which constitutes about 30% of the total turnover) migrated to London. On the one hand, migrating from Sweden to London is much easier than leaving the numerous countries in the European Union. On the other hand, however, today's technologies make any shift much more effortless.

It is important to notice that the size of the relocation strongly depends on the tax system design. In the Swedish experience the trade of bonds, which also were taxed, almost completely disappeared because it was very easy to slightly modify the contract and avoid the taxation. If FTT would only be applied to selected instruments (e.g. equities), trading could shift to some close substitutes, such as American depositary receipts (ADR), contracts for differences or other derivatives. While those instruments are not equity in the legal sense (and thus would not be taxed), their payoffs could be exactly the same or very close to equities' payoffs and thus investors would very likely prefer them to equities. This shows the need for the broad definition of the tax base, preferred by the European Commission. Nonetheless, there are numerous problems with a proper design of the tax system. For example, the taxation of investment funds units would cause the double taxation, since the funds will pay the tax when assembling their portfolio and then customers buying funds' units will also be taxed. On the other hand, excluding the units from the tax base may cause the creation of artificial investment funds which will hold simple assets and will be used for trading these assets tax-free. Also, choosing tax base for options is not obvious. While taxing the premium may seem reasonable, it punishes buying expensive in-the-money options and favors speculative options which are strongly OTM (out-of-the-money). Calculating the tax based on final payment obviously requires some other solutions for options which are never exercised. All these issues, if not solved properly, may allow for the successful tax avoidance.

Price level

Since paying the tax is an additional cost for an owner of a share, it should be discounted from the share's current value. Therefore, it might be expected that after the introduction of FTT equity prices will go down. If we are willing to assume a valuation model based on the discounted cash flows, e.g. the Gordon model, we may derive the theoretical drop in prices.

The standard price of a stock in the Gordon model is given by:

$$V = \frac{D}{r-g} \tag{1}$$

where:

D – the current dividend paid,
r – the interest rate,
g – the rate of growth of the dividend.

[1] In the initial project this was 10% and 90%, respectively.

When we impose the tax, every time the stock is traded there is an additional, negative cash flow. For a given frequency of trading the price of the stock is then given as (see Zator 2013):

$$V = \frac{D}{r-g}\frac{1-e^{-(r-g)n}}{1-(1-\tau)e^{-(r-g)n}} \tag{2}$$

where:

n – the average period of holding a share (the higher n, the lower the frequency of trades),
τ – the tax rate.

To get a rough idea about the magnitude of the change of price, we may assume r to be 5%, g to be 2%, n to be 6 months and increase the tax rate from zero to 0.1%. The theoretical price drop is then around 6%. If we lower n to 3 months, the decrease amounts to around 11.7%. The drop of the price is higher for the shorter holding period because it implies that the stock is traded more often and thus tax needs to be paid more often.

It might be noted that the theoretical value of price decrease depends on the assumption about the length of asset holding period and, what is more important, it may be only of limited use when applied to real markets. The Gordon model, being a simple approach to the valuation, definitely does not precisely reflect the market prices of shares. It is therefore meaningful to analyse the empirical data on price changes caused by transaction taxes.

The effect on the price level is analysed by Umlauf (1993) for the Swedish tax experience. He simply compares the behaviour of prices before and after the announcement of the tax reform and observes 5.3% drop of prices (the theoretical model forecasted the drop of 6.75%). For the abolishment of the tax, Westerholm (2003) observes positive rates of return: 13.9% in Sweden and 5.1% in Finland. However, in that period interest rates were very high and observed returns did not even exceed the risk-free rate.

Saporta and Kan (1997), who analyse the price effect of stamp duty in Great Britain,[2] compare the prices of shares to their corresponding ADRs (American depositary receipts), traded in the United States and thus not subject to stamp duty. They observe that when the tax rates were changed, the returns of stocks were significantly lower than the returns of ADRs (about 15%). However, the trading volume of ADRs was significantly lower than the volume of shares, which may be a part of the explanation of the observed effect. Bond, Hawkins and Klemm (2005), who also analyse British stamp duty, confirm that larger, more often traded companies exhibit a stronger price effect of the tax changes. Hu (1998), who analyses multiple Asian markets and Liu (2007), who analyses Japanese companies by comparing them to their ADRs (similarly to Saporta, Kan 1997), also confirm the implications of the theoretical model.

Therefore the effect of the tax on the price level is rather uncontroversial and relatively close to the implications of theoretical models. The evidence is less clear in the case of the impact of the tax on prices volatility. We postpone a deeper analysis of this issue to Section 3.

[2] Stamp duty is a fee paid in order to legally transfer the ownership of stocks. It has been present in Great Britain for several decades with the tax rate varying over time.

2.2. The impact on the overall economy

Because of the significant size and importance of the financial sector, any changes in its functioning may have a significant effect on the overall economy. Therefore it is reasonable to ask about the effects of imposition of FTT on output or unemployment. The introduction of FTT will probably cause the decrease of activity of financial institutions. This may lead to the decrease of the output, since value added generated by financial institutions is one of its components. The imposition of the tax may also increase the cost of capital. Investors who must pay an extra cost (e.g. a tax) for trading the asset require higher rate of return. Higher cost of capital, in turn, may lead to lower level of investment (see e.g. Oxera 2007 for the discussion of the effect of British stamp duties on the cost of capital).

The precise analysis of macroeconomic costs of FTT is very difficult. It is hard to judge the effect of the financial system reform on the overall economy if we have only a rough idea about the consequences of this reform within the financial system itself. Nonetheless, some attempts of making this assessment were made. The first one was prepared on the request of the European Commission and published in the impact assessment, European Commission (2011). The employed model suggests that after introducing 0.01% tax, output decreases by 0.17%, investment by 0.51% and employment by 0.03%. When the tax rate amounts to 0.1%, output decreases by 1.76% and employment by 0.2%. At the same time the observed effects on the prices of financial instruments are roughly consistent with the theoretical framework presented above.

The impact assessment accompanying the new, revised project of FTT (EC 2013), presents an enhanced model assessing the influence of FTT. The assessment is made in the 40 years horizon and the potential loss of GDP in that period amounts to 0.28%. However, the authors argue that if the tax revenues were used for some productive activity, the net effect of the tax could be positive or close to zero.

Lendvai, Raciborski and Vogel (2013) employ a DSGE model allowing for the imposition of FTT. They calibrate the tax rate so that the collected revenues amount to around 0.1% of GDP, which is in line with the initial estimates of European Commission (the implied tax rate is 0.11%). Such a tax causes 0.2% long-term decrease of the output and 0.16% decrease of the consumption. The authors show that the imposition of FTT has therefore very similar macroeconomic effects as an increase of the corporate tax. Their results suggest that the effects of the tax would be less pronounced as compared to the estimates of European Commission (EC 2011).

2.3. Revenues estimates

The key question for the policymakers concerns the potential revenues which FTT can create. The estimates of the effects of FTT on tax revenues are highly sensitive to the assumptions about the behaviour of the market. McCulloch and Pacillo (2011) gather existing estimates of revenues and try to create a median estimate. Most works calculate the potential revenues with a simple formula similar to:

$$R_C = T \cdot V \cdot E \cdot (1+\frac{T}{C})^e \tag{3}$$

where:

T – the tax rate,
V – the annual volume of trading,
E – the relocation and tax evasion (E is defined as the fraction of turnover which is not relocated),
c – the size of transaction costs,
e – the elasticity of the turnover with respect to transaction costs.

Depending on the assumed tax rate T (varying from 10% to 50% of total transaction costs), assumptions about the elasticity (McCulloch, Pacillo 2011 use median estimates for different segments from the literature) and tax evasion (taken as 20%) as well as inclusion of OTC (over-the-counter) markets, the revenues from the global tax may amount to USD 147–1631 bn per year. These estimates are relatively high, mainly because of assuming relatively low evasion parameter (20% for all – equity, derivative and foreign exchange markets, which means that only 20% of all trading activities is shifted to non-taxed markets).

Estimates presented by the European Commission are more conservative. They employ similar methodology as McCulloch and Pacillo (2011), but use three different values of the elasticity (arguing that the median of all estimates depends on numerous old estimates which do not correspond to today's market characteristics) and different evasion parameter (10% for stocks but 90% for derivatives market). For the most likely parameters of the elasticity and tax rate, estimated revenues from the European market are around EUR 45–50 bn per year. Table 1 presents the estimates of revenues for different values of the tax rate and the elasticity. In the new impact assessment European Commission (2013) translates those calculations to the new project (by taking into account the relative size of the selected 11 economies and the importance of the financial sector in those countries). The total amount of revenues for the new proposal, including only 11 selected countries, is estimated to be around EUR 34 bn.

3. Impact of FTT on volatility

We analyse the impact of FTT on the volatility of prices in a separate section for two reasons. First, there is still little consensus in the existing literature regarding this issue. Second, we want to present the theory relevant for further empirical analysis, which will be presented in Section 4.

Proponents of FTT claim that the tax would curb the speculation and thus stabilize the markets by lowering volatility – and this is one of the most important arguments in favour of the introduction of the tax. Opponents, on the other hand, claim that introducing extra transaction costs would decrease market efficiency and thus increase the observed volatility. Both claims could be argued both on the theoretical and empirical grounds.

3.1. Theoretical models

A large part of the theoretical debate is based on the market models consisting of informed traders and so called noise-traders. Stiglitz (1989) argues that the imposition of a low tax will not significantly change the behaviour of informed traders, who own a fundamental information about the stocks and make decisions with a long-term horizon. Noise-traders, on the other hand, would be strongly affected by the tax. Since they base their decisions on technical analysis signals and make transactions more frequently, their profitability will be decreased by the tax and thus their activity will be lowered. Since their actions lead to the detach of prices from the fundamentals, their elimination should stabilize the market. A theoretical model showing the potential decrease of volatility as a result of imposition of FTT was also proposed by Frankel (1996). Working in the similar framework it is however possible to argue against FTT. Kupiec (1996) shows that although FTT indeed lowers the volatility of prices, it also lowers their level and thus may increase the volatility of returns (which are of the main interest).

Another branch of the theoretical research in this field make use of agent-based models. Starting from Westerhoff (2003), there were numerous attempts to model the financial market and assess the effects of the Tobin tax. The conclusions are, however, mixed. In fact almost all of the works show potential decrease of the volatility as a result of imposition of FTT, but they require some assumptions to be met. They concern market microstructure (e.g. Pellizzari, Westerhoff 2009) or specific tax rate (e.g. Ehrenstein, Westerhoff, Stauffer 2005). The results are therefore sensitive to the assumptions and may provide only a limited support for the volatility-lowering effect of the tax.

Yet another way to assess the FTT impact on the volatility is an experimental research. There were couple of attempts to do so, each getting in fact different results. While Kaiser, Chmura and Pitz (2007) show possible drop of volatility, Bloomfield, O'Hara and Saar (2009) report no significant effect and Hanke et al. (2010) show that the effect depends on the existence of tax heavens. Despite all the limitations of experimental research, further experiments may provide an interesting evidence on the effect of FTT.

3.2. Empirical evidence

Since there is no consensus at the theoretical level, there is an urgent need for the empirical evidence. There were several attempts to provide it, most of which used changes in the tax rates as a natural experiment. Umlauf (1993) reports no significant change in price volatility in the Swedish market (although it slightly increases after introducing the tax). Saporta and Kan (1997) show that British stocks had higher variance that their ADR counterparts. This effect, however, is also likely to be caused by significantly different trading volumes. Phylaktis and Aristidou (2007), who analyse transaction tax in Greece, show that the effect on volatility may depend on the market situation, with the tax increasing volatility during bull periods but decreasing it during bear periods. All these analyses, however, are based on assessing the volatility in the period following a single change of the tax. Potential pitfalls of such approach are illustrated with the results of Su and Zheng (2011). They analyse the Chinese market and report an increased volatility after increasing the tax rate, but also increased volatility after decreasing the tax rate! This may suggest that any change, being a new event which needs to be incorporated, may lead to a temporary increase in the volatility. On top of that, any analysis based on a single event may be misleading – observed changes attributed to the change in the tax rate may be in reality caused by numerous, unobserved effects.

3.3. Proxying for the effect of the tax

Limited number of tax experiments and limitations of their analysis give motivation for employing different methodologies. It may be observed that for a market participant the tax is perceived mainly as an increase in transaction costs. It is therefore reasonable to expect that the effect of the tax would be similar to the effect of other increases of transactions costs, e.g. increase of brokerage fees. Therefore several works tried to analyse the potential effect of FTT by looking at the volatility effect of transaction costs in general.

Jones and Seguin (1997) analyse deregulation (and thus decrease) of commissions paid when trading on New York Stock Exchange. Because other exchange, NASDAQ, was not influenced by the deregulation, they compare the behaviour of volatility on NASDAQ and NYSE in the following period. NYSE exhibits larger drop of volatility than NASDAQ, which suggests volatility-decreasing effect of commissions deregulation. Although the effect is not uniform across companies of different size, it generally supports the claim of FTT causing an increase of the volatility. Interestingly, however, Liu and Zhu (2009) analyse very similar deregulation in Japan and observe the opposite effect – volatility rises. One explanation of this phenomenon is that Japanese reform significantly lowered the cost of transactions for individual investors who are most likely to be noise-traders.

Another approach is taken by Hau (2006), who analyses the discontinuity of the minimum tick size on the French stock exchange. Shares with the price higher than 500 FRF (French franc) are subject to 1 FRF minimum tick size, whereas shares with the price below 500 FRF have 0.1 FRF minimum tick size. There is a large literature showing that the tick size is positively related to the size of spreads and thus to the level of transaction costs, see e.g. Porter and Weaver (1997). Therefore, as argued by Hau (2006), the difference in tick size should be a good proxy for difference in transaction costs and consequently, a proxy for financial transaction tax. Because there should be little difference between stocks trading at 490 FRF and 510 FRF, except for the tick size, Hau compares the volatility of neighbouring groups and concludes that the higher tick size leads to higher volatility, supporting the claims of FTT opponents. In the next section we conduct a similar analysis for the Polish market.

4. Empirical evidence from Warsaw Stock Exchange

We analyse the discontinuity in the minimum tick size on Warsaw Stock Exchange (WSE) which constitutes an exogenous increase of transaction costs and allows for examining the relationship between transaction costs and volatility.

There are four different tick regimes on WSE, presented in Table 2. The minimum tick size, i.e. the precision of quotation of the stock price, depends on the level of this price. The higher the price, the larger the minimum tick (i.e. the lower the precision of quotations). In our analysis we focus on the tick size change around 50 PLN as the most binding and pronounced one. Shares with the price just below this level are subject to 0.01 PLN tick size, whereas shares with the price above this level are traded with 0.05 PLN minimum tick size. We argue that the higher the tick size, the larger the transaction costs in relative terms. When the tick size is very small, it induces a rounding of the price which is almost insensible. When the tick size is large, however, rounding the price to the nearest multiple of the tick may constitute a non-negligible cost. Different tick regimes are characterised by different levels of transaction costs and therefore may be a proxy for an imposition of FTT.

4.1. Data and methodology

Our initial dataset consists of daily prices and trading volumes for all companies quoted on WSE in the period between January 2003 and June 2013. Even though WSE was reestablished in 1991, we discard the data before 2003 because of low liquidity and immaturity of the market.

Because we want to analyse the consequences of the minimum tick size, we first check whether this formal requirement is actually a real constraint. It may be the case that increasing the tick size to 0.05 PLN has no practical consequences because all transactions are conducted with no higher accuracy anyway. Table 3 presents the fraction of closing prices being a multiple of 0.05 PLN in different time periods and price intervals.

As we may observe, the fraction of prices being a multiple of 0.05 PLN is decreasing in time which suggests some connection to the market development stage. It therefore supports earlier discarding of the data prior to 2003. Moreover, it suggests further omission of 2003–2005 period since in those years almost all prices in (40; 50) interval are multiple of 0.05 PLN. Starting from 2006, the situation is better, but still far from fully satisfactory. Nevertheless, we may hope for finding some effects in the years 2006–2013 or 2009–2013.

We concentrate on shares which at some point were traded in the neighbourhood of 50 PLN. We exclude all the stocks which were not traded in (40; 60) interval for at least 30 consecutive days. The final sample consists of 95 companies and 1878 trading days (January 2006 – June 2013). We check the robustness of our results to the definition of the neighbourhood by analysing (45; 55) and (35; 65) intervals. This changes the sample size accordingly, see the details in Subsection 4.3.

We work in a long panel data setup, also known as multiple time series setup. Our response variable is the observed volatility of prices. We define it as daily log range of prices, i.e.:

$$LR_{i,t} = \log\left(\frac{p_{i,t}^{\max}}{p_{i,t}^{\min}}\right) \tag{4}$$

where $p_{i,t}^{\max}$, $p_{i,t}^{\min}$ are maximum and minimum prices for company i on day t.

Such a measure of volatility is argued to be a good proxy for the fundamental volatility (see Alizadeh, Brandt, Diebold 2002) and allows for comparison between different tick regimes. The most popular measure of volatility, standard deviation of returns, is not appropriate in this context because it is biased across different tick regimes. See Hau (2006) for an analysis of the bias of standard deviation of returns as a measure of volatility when different tick regimes are concerned. The same argument also applies to other popular measures, such as realised volatility.

The point of our main interest, i.e. trading in higher or lower tick regime, is captured by a dummy variable $D_{i,t}$:

$$D_{i,t} = \begin{cases} 1, P_{i,t} \in [50; 60) \\ 0, P_{i,t} \in [40; 50) \end{cases} \tag{5}$$

where $P_{i,t}$ is the average of open and close prices for company i on day t.

When the price is not in the interval [40; 60), the observation is excluded and the data point is treated as missing. This causes the panel to be highly unbalanced (around 80% of observations are missing), however in our opinion there are no visible problems with the sample selection and thus we may proceed with the analysis without any special adjustments.

Our argument is that trading in the lower or higher tick regime does not change any fundamental characteristics of the stock but its price. Therefore observed volatility should be affected only if it is somehow connected to the price or if the difference in the tick regimes has some impact on the volatility level. We regress observed volatility $R_{i,t}$ on the dummy variable $D_{i,t}$, controlling for the price level $P_{i,t}$. We also control for other variables which influence the volatility. We include the daily log range of WIG index (index of all companies quoted on WSE), $\ln LR_t$, as a proxy for the market-wide effects (which to some extent may also serve as time effects), as well as the trading volume of specific stock, $V_{i,t}$. It is possible that the trading volume causes some problems with simultaneity, therefore we check how its inclusion influences the parameter of the dummy variable.

There is not much theoretical guidance on whether the individual effects are correlated with explanatory variables in our case. We start from fixed effects model and then check the results of random effects model as well. The estimated equation is as follows:

$$LR_{i,t} = \alpha D_{i,t} + \beta P_{i,t} + \theta Z_{i,t} + \mu_i + u_{i,t} \tag{6}$$

where:

$LR_{i,t}$ – the daily log range of prices,
$D_{i,t}$ – the dummy variable indicating the lower/higher tick regime,
$P_{i,t}$ – the midprice,
$Z_{i,t}$ – the vector consisting of $V_{i,t}$ trading volume,
$LR_{i,t}$ – log range of the index and a constant.

4.2. Empirical results

We estimate the equation using Gretl. Although our panel dimension is 95×1878, due to the large amount of data treated as missing (either really missing or being out of considered price interval) 31 066 observations are used. We report Arellano standard errors (Arellano 1987) given a potential autocorrelation of error term. The results are presented in Table 4.

Hypothesis of the same group intercepts is unambiguously rejected with F statistic being close to 50. Obtained coefficients of control variables are reasonable. Around 70% of exchange-wide volatility is reflected in the individual volatility and a higher volume is related to a higher volatility. What is especially interesting, the coefficient of the dummy variable is positive and statistically significant at the level close to 5%. This would suggest that moving to the higher tick regime increases the volatility by about 0.23 percentage points. Since the mean of the dependent variable (volatility) is 0.032 and its standard deviation equals 0.025, the increase corresponds to around 7% of the mean and around 10% of the standard deviation. Although this is not very much, it is also non-negligible and of some economic significance. When the random effects are used, the coefficient of the dummy variable almost does not change (the new value is 0.00237) and its significance rises considerably (the p-value is close to zero).

The Hausman test (Hausman 1978) indicates that the random effects estimates are consistent (p-value equals to 0.17).

We check the estimates without including the trading volume (as to detect potential simultaneity bias) as well as after including its non-linear functions (squares and logs). We allow for non-linear functions of the index volatility as well. The coefficient of interest, i.e. that of the dummy variable, does not change significantly and in each case remains in the interval (0.019; 0.027), being significant at 5% or 10% level (fixed effects).

What is worth noting, however, is the negative coefficient of the price in our model. Although this coefficient is not significant using conventionally accepted levels, its p-value is relatively small (0.14) and the price level should not be excluded from the model (see Maddala, Kim 1998, p. 141; Kennedy 2008, p. 60 for a discussion of significance level used for testing down the model). What could be the possible interpretation for the observed negative relationship? Note that the higher the price, the smaller is the tick size measured in relative terms. 0.05 PLN tick size constitutes 0.1% of 50 PLN but only 0.083% of 60 PLN. The negative coefficient of the price, therefore, supports the theory of volatility-increasing effect of the higher tick size (higher transaction costs).

On the other hand, the price is obviously very highly correlated with the dummy variable (since the dummy variable is just a non-linear, non-decreasing function of the price). Since the price in the upper regime is on average 10 PLN higher than in the lower one, the linear effect of the price change between regimes amounts to -0.00193, which almost balances out the estimated effect of the higher tick regime. Indeed, when we exclude the price from the model, the estimated effect of the dummy variable is much smaller and economically insignificant (the change in volatility would be around 0.05 percentage point, i.e. 1.5% of current average volatility). It is also statistically insignificant under fixed effects, however it is significant at 10% level under random effects (and the hypothesis of random effects estimator consistency cannot be rejected); results are reported in Table 5. Please note that if individual effects are not correlated with explanatory variables (i.e. the assumption for consistency of random effects estimator is met), random effects estimator is more efficient than fixed effects estimator. Therefore it is not surprising that it leads to lower standard errors and thus higher significance of the parameters. However, if the individual effects are correlated with other regressors, random effects estimator is not consistent and therefore comparison with fixed effect estimates (in the spirit of Hausman's test; see Hausman 1978) is undertaken. For a general discussion of differences between fixed and random effects estimators see e.g. Wooldridge (2002).

It may be the case that due to the sample characteristic, in the presence of potential non-linearities of the price effect, the observed effect of the tick regime does not really correspond to the effect of increased transaction costs and it is artificially induced by the negative coefficient of the price. To address this issue and reinforce our results, the next subsection presents the results of some robustness checks.

4.3. Robustness checks

First and second half of the period

First, we divide our sample into two halves, corresponding to 2006–2009 and 2009–2013 periods. Obtained results do not change significantly when we analyse only the second half of the sample, when the market should be more mature. The coefficient of the dummy variable is positive and close to 0.02 when the price is included, but very small and even negative when the price is excluded from the model. With the fixed effects the dummy coefficient is insignificant (see Table 6), but with random effects it is significant (p-value equals to 0.015) and of very similar magnitude (0.00174). GLS estimates are consistent according to Hausman's test.

In the first half of the period the increase in volatility driven by the higher tick regime is even larger and significant at 5% level with fixed effects. The results are presented in Table 7. We may therefore conclude that our results are robust to limiting the sample size and the volatility-increasing effect of the higher tick size is present in both halves of the period. Observe that the large part of the years 2006–2009 was a period of bull market, whereas the years 2009–2013 were closer to bear market. Therefore a higher coefficient of the dummy variable in the first half of the period is in line with the findings of Phylaktis and Aristidou (2007).

Impact of stock size

It may be argued that the reported effect of the tax is caused by the differences in capitalization. It is known that smaller stocks exhibit larger volatility and if it happens that smaller stocks are more frequently in the higher tick regime, the reported effect of the tick size may be explained by the size of analysed companies. Using panel data fixed effect estimator should mitigate this problem since a generally understood size of the company do not change during analysed period (while it is possible that a company treated as "small size" in the beginning of the period has grown significantly and joined "large size" stocks in the end of the period, it is rather uncommon). The size effect should be therefore captured by fixed effects, but such a control method certainly is not perfect. In addition, the effect of the tick size may be different across companies of different size.

To get better picture of the effect and to preclude the possibility that the observed effect is caused by differences in stock capitalization, we divide the sample into three groups of small, medium and large stocks. The division is based on average market capitalization in the analysed period. Small stocks are the ones with average capitalization below PLN 250 mn and large stocks have capitalization larger than PLN 1000 mn, with medium stocks' capitalization lying inside this interval. We perform the separate regression for these three groups. In Table 8 we present the values of tick dummy coefficients together with their p-values for fixed and random effects. The effect of the tick dummy is in all cases positive, though in the case of medium stocks it is small and statistically insignificant. For small and large stocks coefficients have high statistical significance for random effects estimation. For fixed effects only large stocks coefficient is significant at 10% level. Observe, however, that random effects and fixed effects estimates are very close in all cases.

Although for some reason we see no significant proof of positive tick size effect in the middle group, the positive effect of the tick dummy is present both in the group of small and large stocks. Therefore the effect cannot be explained by differences in capitalization.

Potential nonlinearities and price effect

We include the non-linear components of the index log range ($I\ ndLR$) and the trading volume (V) and obtain very similar estimates of the dummy coefficient (as already mentioned in Subsection 4.2). Including the non-linear components of the price does not seem to be a good idea – a combination of non-linear functions of the price can easily mimic the effect of the tick regime and make the results insignificant. What we can do to check if some non-captured non-linearities do not influence our positive coefficient of the dummy variable, is to check whether the coefficient of the price is negative also outside (40; 60) price interval. If it is, it suggests that the higher tick regime really rises the volatility (and so balances out the volatility-lowering effect of larger price). If it is not, it may be an indicator that our results may be just a sample characteristic and are poor evidence for volatility-increasing effect of the higher tick size.

We employ the similar framework as in the original model, but this time we consider two different price intervals: (25; 50) and (50; 75). We check whether within these intervals there exists a negative relationship between the observed volatility and the price (please note that within the interval the tick size is constant). It turns out that the negative effect is present in both intervals, however smaller in magnitude than in the original model. The coefficients of the price for different model configurations are presented in Table 9. Please note that in each case we estimate the original model (equation (6) for the two subsamples) without the dummy variable (i.e. the dummy in both cases is constant so it has to be excluded to avoid multicollinearity). The effect of the price is surprisingly stable – in all cases it is close to -5.5e-5. Therefore the negative relationship between the price level and volatility seems to be there, however its magnitude is smaller than estimated in the original model (Table 4). If we believe that -5.5e-5 is a proper and universal measure of the effect of price on the volatility and fix the price coefficient at this value in the original regression (equation (6)), estimated coefficient of the dummy variable is still positive (significant with random effects, p-value equal to 0.16 with fixed effects), but visibly smaller (slightly above 0.001). It is not clear whether this smaller value is a better estimate of the real effect of the tick regime than the larger, original one which amounts to 0.0023.

Neighbourhood definition

Our results suggest that the reported positive coefficient of the dummy variable really represents the effect of the increased tick size and thus the increased transaction costs lead to higher volatility. We now need to check the robustness of our results to the definition of the neighbourhood of 50 PLN. We substitute the initial (40; 60) interval with a broader (35; 65), and a narrower one (45; 55). We estimate the original model from equation (6) for two samples selected using these intervals.

When we consider the broader interval, the number of companies rises to 113. The results are presented in Table 10. Presented coefficient of the dummy variable, obtained using fixed effects, is not statistically significant. If we switch to random effects, the coefficient remains basically unchanged (0.00125) but the p-value is smaller than 1%.

The results for the narrower interval (45; 55) are presented in Table 11. Once again, the coefficient of the dummy variable estimated using fixed effects is not significant but the value obtained by random effects is almost the same (0.00151), while the p-value drops below 1%. The coefficient obtained for the narrower interval is higher than the one obtained for the broader interval, which is in line with our expectations (the effect of the tick size affects the shares with price close to 50 PLN more strongly). At the same time the coefficient of the price is very close to those from Table 9. One may therefore claim that the value obtained using the narrower or broader interval (since they are very alike) is a better estimate of the real effect of the tick size increase. Anyways, considering different intervals supports our main message: the increase of the tick size leads to the increase in volatility.

Dynamic model

We try to validate the results by employing an additional specification which includes lags of the dependent variable. In such a setting the coefficient of the tick dummy captures only the short-term effect of the increased tick size, but algebraic transformation allows for recovering the long-term effect and comparing it to the one estimated with the static model.

Due to large number of observations the significance of lags turns out to be very high and even ninth lag may be significant. However, the magnitude of the coefficient is very small and inclusion of so many lags does not improve the fit in a significant way – the information criteria actually suggest the model with no lags. At the same time the total sum of lagged coefficients remains fairly stable, no matter what number of lags we choose. Results for model with 9 lags are presented in Table 12. Observe that the coefficient of the dummy variable went down to 0.0013 and became even more significant (we still report Arellano standard errors because of potential heteroskedasticity). As expected, the short-run effect is therefore lower than the long-term one. To recover the long-term coefficient one may multiply obtained coefficient by $(1-\sum_{i=1}^{9} b_i)^{-1}$, where b_i stands for a coefficient of ith lag. When we do so, the estimated long-term effect is equal to 0.0025 and is very close to the original estimate from the static model.

The final question concerns the magnitude of the observed volatility-increasing effect. Estimated coefficients in different models vary from 0.001 to almost 0.003, which translates to 4–11% of the average volatility and 4.5–12% of the standard deviation of the volatility. Since the difference in the volatility should be smaller when the price moves away from 50 PLN, we may expect that using relatively broad intervals around 50 PLN leads to an underestimation of the coefficient (narrower intervals cannot be used, though, because of limited size of the sample).

Therefore, we are willing to believe that the effect reaches the upper bounds of mentioned intervals and is not lower than 0.0015 (which corresponds to over 5% of the observed volatility). Hau (2006) reports 30% increase of volatility, however he analyses the tenfold increase of the tick size, whereas in our case the tick rises only 5 times. Our coefficient is therefore rather small, but non-negligible. The small size of the effect is understandable if we take into account the fact that the minimum tick size on WSE is only sporadically constraining market participants (see Table 3).

5. Conclusions

This paper provides a review of evidence concerning financial transaction tax, presented in the context of the proposal of European Commission. Since FTT is currently on the agenda of policymakers (the new proposal was issued in February 2013 and is currently discussed at the European level), we believe that the importance of this topic has increased recently. In February 2013 Polish Ministry of Finance declared that Poland awaits the final result of the negotiations at the European level and postpones its decision on whether to introduce the tax until then. Therefore, even though Polish experts do not actively take part in the preparation of the current proposal, the academic discussion about FTT in the Polish context would be very beneficial. We hope that this paper stimulates the interest in this topic among the members of academia.

Since financial transaction tax is still mostly theoretical concept, the consequences of its introduction are rather hard to assess. It is clear that FTT will decrease the trading volume and the level of prices, but the exact magnitude of these effects remains uncertain. Nevertheless, the literature provides some guidance in this field and we may hope that the decrease of trade will not be very sharp. The effects on the entire economy are even harder to assess. A couple of available studies predict that the macroeconomic impact of introducing FTT should not be very different from the impact of potential increase of the corporate tax.

There is still little agreement in the literature regarding the effect of the tax on the volatility of prices in the financial markets. While most of the theoretical models predict the decrease of volatility as a result of imposition of the tax, the evidence from the empirical data is mixed (with majority of works pointing to the increase of volatility as a result of the tax). In this work we provide an additional evidence from Warsaw Stock Exchange, which suggests that there is a positive relationship between the level of transaction costs and volatility. We analyse the minimum tick size change when the price of a share crosses 50 PLN and conclude that this exogenous increase of the transaction costs leads to a minor, but non-negligible increase of volatility. If the effect of FTT does not differ from the effect of general increase in transaction costs, our results support the opponents of the tax regarding the impact on volatility. Our framework is similar to Hau (2006) and different from most of the remaining studies. By analysing the panel data we avoid numerous problems arising in the analysis of a single change of the tax. It must be noted that, just like the majority of the existing literature, we assess the short-term, daily volatility. While assessing the long-term volatility is much more challenging, it may be of even higher significance for the economy.

The final decision on whether to adopt financial transaction tax is a complex one and requires very profound consideration. In our opinion the argument of improving market functioning is not very convincing. As shown in this article and in some other works, it is very doubtful that the tax will stabilize the markets. At the same time, however, we are aware that there are many different arguments used by the advocates of the tax, with collecting extra tax revenues from the financial sector being probably the most important one. In the current situation of public budgets in many European countries (including Poland) providing some extra revenues may be inevitable. Thus the question becomes "is FTT a better way to secure those revenues than other taxes?" instead of "is FTT desirable?" This considerably changes the shape of the debate. Even though FTT may have some drawbacks (e.g. may bring some moderate decrease in markets stability), other taxes do have them as well and we need to decide what solution will have less deteriorating consequences for the real economy.

We believe that further debate should discuss FTT in such a context, comparing it to other potential solutions. The key issue we should be concerned with, is the impact on the real economy and the feasibility of the tax. The current state of knowledge about potential relocation and tax avoidance is far from satisfactory and any research in this direction, although very challenging, would be very welcome.

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Appendix

Table 1

Estimates of revenues from FTT when taxing all instruments (all) or when exempting currency transactions (no CTT) for different values of elasticity and tax rates

	Tax rate 0.1%	**Tax rate 0.05%**	**Tax rate 0.01%**
		elasticity = 0	
FTT (all)	221.9	110.9	22.2
FTT (no CTT)	96.7	48.4	9.7
		elasticity = 1	
FTT (all)	136.8	75.5	19.6
FTT (no CTT)	88.6	43.3	9.4
		elasticity = 1.5	
FTT (all)	92.0	59.3	17.9
FTT (no CTT)	70.8	40.9	9.3

Note: all figures assume that the tax covers the entire area of EU-27 and are in billions of euro per year.
Source: EC (2011).

Table 2

Different tick regimes functioning on Warsaw Stock Exchange

Price level	**Tick size**
≤ 50 PLN	0.01 PLN
(50; 100] PLN	0.05 PLN
(100; 500] PLN	0.1 PLN
> 500 PLN	0.5 PLN

Table 3

The fraction of closing prices being a multiple of 0.05 PLN in different time periods and price intervals on Warsaw Stock Exchange (in %)

Price interval	**Period**		
	2003–2005	**2006–2008**	**2009–2013**
[0; 20]	50.99	47.06	36.77
(20; 30]	99.12	72.15	61.58
(30; 40]	89.24	67.79	58.54
(40; 50]	98.34	78.32	67.10
> 50	100.00	100.00	100.00

Table 4

Log range of prices explained by tick regime dummy, price, volume and index log range. Fixed effects and Arellano standard errors were used

	Coefficient	**Standard error**	***t*-ratio**	***p*-value**
Const	0.0295	0.00634	4.66	0.000***
D	0.00235	0.00121	1.94	0.052*
V	5.65e-9	2.51e-9	2.25	0.024**
ind LR	0.685	0.0551	12.43	0.000***
P	-0.000193	0.000131	-1.47	0.141

Notes: we used 31 066 observations, included 95 cross-sectional units; time-series length: minimum 34, maximum 1276. Mean dependent var 0.032, S.D. dependent var 0.025, sum of squared residuals 15.626, S.E. of regression 0.022, R2 0.197, adjusted R2 0.195, F(98, 30967) 77.584, *p*-value(F) 0.000***. Stars denote significance at usual levels of 10%, 5% and 1%.

Table 5

Model without controlling for the price. Random effects used

	Coefficient	**Standard error**	***t*-ratio**	***p*-value**
Const	0.0261	0.00142	18.37	0.000***
D	0.000488	0.000294	1.66	0.098*
V	5.67e-9	4.86e-10	11.67	0.000***
ind LR	0.686	0.0138	49.59	0.000***

Notes: we used 31 066 observations, included 95 cross-sectional units; time-series length: minimum 34, maximum 1276. Mean dependent var 0.0319, S.D. dependent var 0.025, sum of squared residuals 18.940, S.E. of regression 0.025. Hausman test for the null hypothesis that GLS estimates are consistent: asymptotic test statistic: $\chi^2(3) = 1.369$, with *p*-value = 0.713. Stars denote significance at usual levels of 10%, 5% and 1%.

Table 6

Estimates of the initial model (equation (6)) for the second half of the sample, 2009–2013 period. Fixed effects and Arellano standard errors used

	Coefficient	**Standard error**	***t*-ratio**	***p*-value**
Const	0.0272	0.00775	3.504	0.000***
D	0.00177	0.00137	1.295	0.195
V	4.15e-9	1.55e-9	2.676	0.007***
ind LR	0.646	0.073	8.844	0.000***
P	-0.0000204	0.000164	-1.245	0.213

Notes: we used 12 950 observations, included 52 cross-sectional units; time-series length: minimum 1, maximum 858. Mean dependent var 0.026, S.D. dependent var 0.020, sum of squared residuals 4.546, S.E. of regression 0.019, R2 0.129, adjusted R2 0.125, F(55, 12894) 34.678, *p*-value(F) 0.000***. Stars denote significance at usual levels of 10%, 5% and 1%.

Table 7

Estimates of the initial model (equation (6)) for the first half of the sample, 2006–2009 period. Fixed effects and Arellano standard errors used

	Coefficient	**Standard error**	***t*-ratio**	***p*-value**
Const	0.0369	0.0072	5.130	0.000***
D	0.003	0.00139	2.161	0.031**
V	1.12e-8	4e-9	2.801	0.005***
ind LR	0.622	0.0596	10.434	0.000***
P	-0.000286	0.000148	-1.929	0.054*

Notes: we used 18 116 observations, included 86 cross-sectional units; time-series length: minimum 1, maximum 701. Mean dependent var 0.036, S.D. dependent var 0.027, sum of squared residuals 10.778, S.E. of regression 0.024, R2 0.198, adjusted R2 0.194, F(89, 18026) 49.908, *p*-value(F) 0.000***. Stars denote significance at usual levels of 10%, 5% and 1%.

Table 8

Estimates of tick dummy coefficients for small, medium and large stocks together with their *p*-values for fixed and random effects

	Small	**Medium**	**Large**
Coefficient (FE)	0.00401	0.00059	0.00187
p-value (FE)	0.20	0.76	0.09
Coefficient (RE)	0.00415	0.00057	0.00186
p-value (RE)	0.00	0.56	0.01

Table 9

The coefficients of the spot price in the models estimated for price intervals (25; 50) and (50; 75)

Method	**Coefficient**	**Standard error**	***p*-value**
	(25; 50) interval		
Fixed effects	-5.37e-5	6.62e-5	0.418
Random effects	-5.64e-5	1.73e-5	0.001
	(50; 75) interval		
Fixed effects	-5.73e-5	5.24e-5	0.274
Random effects	-5.77e-5	2.06e-5	0.005

Table 10

The estimates of the original model (equation (6)) when the neighbourhood of 50 PLN was defined as (35; 65) – wider interval. Fixed effects and Arellano standard errors used

	Coefficient	**Standard error**	***t*-ratio**	***p*-value**
Const	0.0216	0.00468	4.621	0.000***
D	0.00124	0.00127	0.976	0.329
V	5.51e-9	2.46e-9	2.246	0.025**
ind LR	0.735	0.0471	15.599	0.000***
P	-1.98e-5	9.79e-5	-0.202	0.840

Notes: we used 49 303 observations, included 113 cross-sectional units; time-series length: minimum 42, maximum 1578. Mean dependent var 0.033, S.D. dependent var 0.027, sum of squared residuals 28.342, S.E. of regression 0.024, R2 0.201, adjusted R2 0.199, F(116, 49186) 106.817, *p*-value(F) 0.000***. Stars denote significance at usual levels of 10%, 5% and 1%.

Table 11

The estimates of the original model (equation (6)) when the neighbourhood of 50 PLN was defined as (45; 55) – narrower interval. Fixed effects and Arellano standard errors used

	Coefficient	**Standard error**	***t*-ratio**	***p*-value**
Const	0.0212	0.00537	3.942	0.000***
D	0.00149	0.00141	1.058	0.290
V	4.13e-9	1.61e-9	2.568	0.010**
ind LR	0.726	0.0531	13.679	0.000***
P	-5.84e-5	0.000111	-0.526	0.599

Notes: we used 38 500 observations, included 62 cross-sectional units; time-series length: minimum 101, maximum 1578. Mean dependent var 0.030, S.D. dependent var 0.023, sum of squared residuals 17.862, S.E. of regression 0.022, R2 0.150, adjusted R2 0.149, F(65, 38434) 104.415, *p*-value(F) 0.000***. Stars denote significance at usual levels of 10%, 5% and 1%.

Table 12
The estimates of dynamic model (original model from equation (6) expanded with 9 lags of dependent variable). Fixed effects and Arellano standard errors used

	Coefficient	**Standard error**	***t*-ratio**	***p*-value**
Const	0.0121	0.00334	3.63	0.000
D	0.00129	0.000594	2.18	0.030
V	3.97e-9	1.51e-9	2.63	0.009
ind LR	0.51	0.0378	13.48	0.000
P	-0.000106	7.1e-5	-1.50	0.134
LR-1	0.206	0.00756	27.31	0.000
LR-2	0.081	0.00894	9.07	0.000
LR-3	0.0497	0.00713	6.97	0.000
LR-4	0.0312	0.00739	4.22	0.000
LR-5	0.021	0.00834	2.51	0.012
LR-6	0.0262	0.00894	2.93	0.003
LR-7	0.0224	0.00646	3.47	0.001
LR-8	0.0319	0.00733	4.36	0.000
LR-9	0.0233	0.0069	3.38	0.001

Notes: we used 22 870 observations, included 95 cross-sectional units; time-series length: minimum 23, maximum 1100. Mean dependent var 0.030350, S.D. dependent var 0.021853, sum of squared residuals 8.054490, S.E. of regression 0.018811, R2 0.262495, adjusted R2 0.259028, F(107, 22762) 75.71513, *p*-value(F) 0.000***. Stars denote significance at usual levels of 10%, 5% and 1%.