Do Individual Index Futures Investors Destabilize the Underlying Spot Market?*

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This paper provides empirical evidence on the impact of the introduction of index futures in Poland on the volatility of the underlying stock markets. The dominance of individual investors on the Polish futures market enables us to investigate the destabilization hypothesis more directly than available studies do and to contribute evidence specifically about the influence of individuals trading in index futures on spot market volatility. Using a Markov-switching-GARCH approach which allows us to endogenously identify volatility regimes the empirical evidence suggests that the introduction of index futures does not destabilise effect on the stock market.

JEL Classification: C32, G10, G14, G20

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Abstract

This paper provides empirical evidence on the impact of the introduction of index futures in Poland on the volatility of the underlying stock markets. The dominance of individual investors on the Polish futures market enables us to investigate the destabilization hypothesis more directly than available studies do and to contribute evidence specifically about the influence of individuals trading in index futures on spot market volatility. Using a Markov-switching-GARCH approach which allows us to endogenously identify volatility regimes the empirical evidence suggests that the introduction of index futures does not destabilise effect on the stock market.

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1 Introduction

Through the process of arbitrage futures markets are closely linked to the underlying spot markets. An important concern is the impact of the introduction of a futures market on the underlying spot market volatility. Two main strands of arguments exist in the theoretical literature. On the one hand, futures trading destabilizes the underlying spot market by increasing its volatility due to the existence of uniformed investors. Because of high leverage badly informed investors induce noise in the price discovery process and lower the information content of prices. This implies higher spot market volatility compared to the situation without a futures market (Cox (1976), Cagan (1981), Figlewski (1981), Stein (1987), Hart and Kreps (1986)). On the other hand, researchers argue that futures markets have a stabilizing effect on the underlying spot market because futures trading improves price discovery, enhances market efficiency, increases market depth as well as information flow and contributes to market completion. Consequently, the introduction of futures trading lowers the volatility of the underlying spot market (Powers (1970), Danthine (1978), Bray (1981), Kyle (1985), Stoll and Whaley (1988)).

Due to the inconclusiveness of the theoretical literature, empirical investigations are necessary to discover additional insights on the impact of futures trading on spot market volatility. While a direct empirical test of the theoretical approaches is not possible, the empirical literature exploits the introduction of futures markets to quantify the effect on spot returns volatility. The majority of the recent empirical time-series investigations exhibit at least two common characteristics (Antoniou et al. (1998), Gulen and Mayhew (2000), Antoniou et al. (2005)): First, the studies implement GARCH-type models augmented by dummy variables. The dummy variables allow to distinguish the pre- and post-futures period and an investigation of the impact the introduction of index futures markets has on spot market returns volatility. Second, the available studies provide to a large extent empirical evidence for mature stock markets in which institutional investors are the predominant trader type.

The first characteristic can be challenged because the dummy variable approach relies on an exogenous determination of the shift in stock returns volatility. Moreover, the simple technique can only model an abrupt one step change in the volatility process which does not provide a complete picture of volatility changes. It would be preferable to model the shift in stock returns volatility endogenously and let the data speak for themselves. By construction, the one step dummy variable cannot capture a gradual adjustment to a new volatility regime and is not able to allow for the possibility of a transitory volatility change.

Concerning the second aforementioned characteristic the introduction of index futures markets in mature countries was mainly in the 80s when institutional investors were the dominant player in stock markets. Financial economists tend to view institutional investors as informed traders and individual investors as uninformed ones (for example, Lee et al. (1999), Cohen et al. (2002), Barber and Odean (2008), Kaniel et al. (2008)). Therefore, the characteristic of futures markets populated mainly by informed investors does not fit to the argumentation of the destabilization hypothesis. In the destabilization hypothesis uninformed investors induce noise in the price discovery process and lower the information content of prices. Consequently, the test approach should be built in an institutional framework in which uninformed investors play a dominant role.

In this paper, we take into account both aspects outlined above. Instead of a simple dummy variable approach we implement a modifiaction of the Markov-switching-GARCH model proposed by Gray (1996b). This technique provides empirical evidence of whether the introduction of a futures market changes the volatility structure of stock returns in the underlying spot market. The Markov-switching-GARCH technique allows endogenous specifications of volatility regime shifts and provides evidence whether the volatility structure has changed transitorily or permanently. Moreover, we exploit a unique institutional peculiarity of the index futures market in Poland which is in line with the destabilization hypothesis. In terms of trading turnover individual investors in the Polish futures market account for more than 80% in the first five years after the start of the futures market (1998 - 2002) and around 75% during the following three years (2003 - 2005). Across recent years (2006 - 2007) individual investors' proportion of trading volume is still well above 55%. It is this feature of presumably uninformed individual traders as dominant players in the Polish futures markets which enables us to investigate the destabilization hypothesis more directly than available studies do. If the destabilization hypothesis is valid and individual investors are uninformed traders, our findings for the Polish stock market should provide clear cut evidence in favour of a permanent increase of stock market volatility after the introduction of the index futures market segment.

The paper contributes to the well-known and voluminous literature on the impact of futures markets on the underlying spot market. More recently, studies in this area take explicitly into account the importance of the investor structure and, in particular, the role of individual investors. McMillan and Garcia (2008) investigate the impact of the introduction of the mini-futures contract for the Spanish Ibex index in November 2001 on overall market efficiency. The main purpose of introducing this kind of contract was to encourage individual traders' access to futures markets. The mini-futures contract has introduced greater noise into the dynamic relationship between spot and futures markets. Kurov (2008) investigates the US S&P 500 and Nasdaq-100 E-mini futures to examine whether futures traders use feedback trading strategies. The median trade size in both market segments is consistent with small individual traders accounting for a substantial proportion of trading. The empirical findings for both E-mini futures show that investors are positive feedback traders which buy after price increases and sell after price declines.

Also related to our study is the one by Bae et al. (2004) for South Korea. The authors analyse the effect of the introduction of KOSPI 200 index futures trading in May 1996 on returns volatility and market efficiency. Among other peculiarities in the KOSPI 200 futures market individual investors account for about 40% of the trading volume. Their results indicate that stock returns volatility and market efficiency increase. However, when compared to non-KOSPI 200 stocks, KOSPI 200 stocks show lower returns volatility after the introduction of futures trading. None of the studies summarized above relies on an econometric technique which allows for an endogenously determined volatility regime shift and a market setting with individual investors as the by far dominant investors' group. We, therefore, apply the Markov-switching-GARCH model to the Polish index futures.

In order to empirically investigate the impact of individuals' index futures trading on the spot market volatility in Poland, section 2 describes the institutional background of these two markets. Section 3 introduces the dataset and the Markov-switching-GARCH methodology, before section 4 presents and interprets the empirical results. Section 5 summarises our findings and concludes.

2 The Polish Spot and Index Futures Markets

The first stock exchange in Warsaw was founded in 1817. Having been closed during World War II and the communist era, the Polish stock market was re-opened on 16 April 1991.

Exactly three years later, the WIG20 stock price index was launched.[1] This index reflects the performance of twenty blue chip stocks listed on the main market of the Warsaw Stock Exchange (WSE). The mWIG40 (called MIDWIG until 18 March 2007), a mid-cap price index, followed on 21 September 1998, and the TechWIG price index, representing innovative technologies, on 31 December 1999.

Being a medium-size stock exchange in Europe, the WSE ranks first in market capitalisation among the exchanges in all Central and Eastern European Countries (CEECs). As of 2007, total market capitalization of the Warsaw Stock Exchange has reached USD 212 billion, which is far more than other Eastern European markets like Budapest (USD 46 billion) and Prague (USD 70 billion). In fact, its size rivals that of smaller Western European exchanges such as Vienna (236 billion USD) or Luxembourg (166 billion USD). The Polish stock market has been growing rapidly in part because formerly state-owned companies were privatised and listed on the WSE. The first foreign company (Bank Austria Creditanstalt AG) was listed at the WSE on 14 October 2003. Since 1 May 2004, the exchange's market structure has been complying with EU standards, i.e. securities trading has two segments, the main market and the regulated unofficial parallel market.

Initially, there was a spot market only. Futures contracts on the WIG20 have been traded at the WSE since 16 January 1998. This was the first derivative product introduced by the exchange, which quickly became very popular among Polish investors. On 1 August 2000, futures contracts on the TechWIG index were introduced, with contracts on the mWIG40 index following on 18 February 2002.[2] Since 1998 the index futures market in Poland has grown substantially. Figure 1 illustrates trading volume and open interest at the WSE for the years 1998 to 2007.

[Insert Figure 1 about here]

With a volume traded of 9.35 million contracts and a notional value of 122,191 million

[1] All information about the Warsaw Stock Exchange is taken from the annual fact books. The comparisons of the Warsaw Stock Exchange with other exchanges are based on the World Federation of Exchanges Annual Report 2007.

[2] For completeness: Futures contracts on individual stocks were first launched on 22 January 2001. Put and call options with the WIG20 as underlying were introduced on 22 September 2003, and stock options started trading on 17 October 2005. Moreover, trading has been suspended in futures on three individual stocks. Futures contracts on US$ debuted on 25 September 1998, followed by futures on the Euro on 1 May 1999. T-note futures were launched on 14 February 2005. Ordinary warrants started trading on 9 March 1998, with American-style warrants on WIG20 futures contracts joining in on 24 September 2001. Convertible bonds were first quoted on 25 April 2002.

US$, total index futures trading on the WSE in 2007 is comparable in volume to Western European markets such as Borsa Italia (6.74 million contracts; 1,428,831 million US$) or the spanish MEFF (11.30 million contracts; 1,774,694 million US$). At the same time, the Polish stock index futures market is considerably larger compared to other CEECs, as for instance Budapest (3.95 million contracts; 5,758 million US$), or even some Western European exchanges such as Athens (2.74 million contracts; 52,096 million US$) or Vienna (0.23 contracts; 21,561 million US$).

Derivatives are traded in the continuous trading system, whose trading hours were 10.15am to 4.00pm prior to the introduction of the quotation system WARSET on 17 November 2000. Thereafter, derivatives were traded from 9.00am to 4.10pm. Since 3 October 2005, the derivatives market has been closing at 4.20pm, with an auction being held at opening and closing. The contracts expire in March, June, September, and December. The last trading day of any given contract is the third Friday of its expiry month, or the last trading day prior to that Friday in case of public holidays.

More relevant to our research questions is the unique investor structure on the Polish futures market. Information on the investor structure is provided by the WSE. Figure 2 exhibits all details on the investor structure from the years 1997 to 2007. On the futures markets, individual investors are the dominating trader type, accounting for about 75% of turnover value on average over the past eight years, with domestic institutions contributing 20%, while the remaining 5% were allocated to foreign investors. During the last 10 years, the spot turnover shares of domestic private, domestic institutional and foreign investors have been relatively equal.

[Insert Figure 2 about here]

This dominance of individual investors on the futures market is mainly due to three factors: First, small transactions can be settled at the Polish derivatives market. For example, the value of an index futures contract equals the product of the multiplier and the price of the underlying. The former was set to only 10 zl, which currently equals about 2.75 US$. This small multiplier makes WIG20 index futures affordable for small investors. Second, individual investors who wish to trade on the Polish futures market can easily register to do so without formal barriers. Third, Polish pension funds are not permitted to trade in derivatives.

3 Data and Methodology

The dataset comprises time series of daily close price observations on the WIG20, mWIG40, and TechWIG stock price indices. In order to control for international influence on the Polish stock market, we further include daily close prices of the S&P500 index in our dataset. The time series for the WIG20, mWIG40, and TechWIG are obtained from Warsaw Stock Exchange (WSE). Data for the S&P500 index is taken from Thomson Financial Data. For the WIG20 the sample period starts on 1 November 1994, which is the first complete month with five trading days per week, and covering 3291 trading days it ends on 31 December 2007. The sample period for mWIG40 covers 2503 trading days from 21 September 1998 to 31 December 2007. 2004 trading days are included in the sample for the TechWIG covering the period between 31 December 1999 and 31 December 2007.[3] The index return is defined as $R_t = 100 \ln(\text{Index}_t/\text{Index}_{t-1})$.

An appropriate econometric technique for analyzing stochastic volatility shifts is provided by Markov-switching-GARCH models. Apart from some early methodological contributions to Markov-switching models scattered in the literature, their modern formal foundation is due to Hamilton (1988, 1989). In our analysis we make use of a Markov-switching-GARCH model as developed in Gray (1996b), but modify his framework in two respects. First, we adapt Gray's model for t-distributed index returns within each regime and second, we incorporate a GARCH-dispersion specification as proposed by Dueker (1997).[4]

The idea of an univariate Markov-switching model is that the data generating process of the variable of interest—here of the daily stock returns of the WIG20 index—may be affected by a non-observable random variable S_t which represents the state the data generating process is in at date t. In our analysis, the state variable S_t differentiates between two volatility regimes and consequently takes on two distinct values. $S_t = 1$ indicates that the data generating process of the WIG20 index returns is in the high-volatility regime whereas

[3]The TechWIG price index was introduced on 19 May 2000 and back-calculated to 31 December 1999. We include in our sample TechWIG data from 31 December 1999.

[4]The use of t-distributed rather than normally distributed returns within each regime is motivated by the 'fat-tail'-property of stock index returns (Bollerslev (1987)). Alternatively, any other heavy-tailed parametric distribution like the *Generalized Error Distribution* (GED) suggested by Nelson (1991) could be specified to govern the tail thickness of the index returns. However, as will be argued below, the t-distribution constitutes a good empirical model for our dataset.

for $S_t = 2$ the generating process is in the low volatility regime.

To set up our Markov-switching-GARCH model, recall first the probability density function of a (displaced) t-distribution with ν degrees of freedom, mean μ and variance h:

$$t_{\nu,\mu,h}(x) = \frac{\Gamma[(\nu+1)/2]}{\Gamma[\nu/2]\sqrt{\pi \cdot (\nu-2)h}} \left[1 + \frac{(x-\mu)^2}{h(\nu-2)}\right]^{-(\nu+1)/2}, \tag{1}$$

where $\Gamma(z) \equiv \int_0^\infty t^{z-1} \mathrm{e}^{-t}\, dt,\ z > 0$, denotes the complete gamma function. Next, we will specify stochastic processes for the mean and the variance in regime i (μ_{it} and h_{it}, respectively) according to which the return at date t (denoted by R_t) is generated conditional upon the regime indicator $S_t = i, i = 1, 2$. Following Gray (1996b) Markov-switching framework, the conditional distribution of the returns can be represented as a mixture of two displaced t-distributions:

$$R_t|\phi_{t-1} \quad \sim \quad \begin{cases} t_{\nu_1,\mu_{1t},h_{1t}} & \text{with probability } p_{1t} \\ t_{\nu_2,\mu_{2t},h_{2t}} & \text{with probability } (1-p_{1t}) \end{cases}, \tag{2}$$

where ϕ_t represents the usual time-t information set and $p_{1t} \equiv \Pr\{S_t = 1|\phi_{t-1}\}$ denotes the so-called *"ex-ante probability"* of being in regime 1 at time t.

In our regime-dependent mean equations we explicitly take into account the possibility of first order autocorrelation in stock returns (by including R_{t-1}) and the interdependence of the Polish stock market with the international stock market. For this latter aspect we include the lagged S&P500 index returns R_{t-1}^{SP} as a control variable in the mean equation:

$$\mu_{it} = a_{0i} + a_{1i}R_{t-1} + a_{2i}R_{t-1}^{\mathrm{SP}} \qquad \text{for } i = 1, 2. \tag{3}$$

In contrast to the mean equation (3) the specification of an adequate GARCH-process for the regime-specific variance h_{it} is more problematic. Technically, this complication is phrased as *"path dependence"* and stems from the GARCH lag structure which causes the regime-specific conditional variance to depend on the entire history $\{S_t, S_{t-1}, \ldots, S_0\}$ of the regime-indicator S_t. We will circumvent this problem by applying the same collapsing procedure as Gray (1996b). For this we have posited in Eq. (2) that the data generating process that determines which regime observation t comes from in fact depends on the probability p_{it} as calculated from Eq. (9) below. From Eq. (2) the variance of the stock return at date t can be expressed as:

$$\begin{aligned} h_t &= E\left[R_t^2|\phi_{t-1}\right] - \left\{E\left[R_t|\phi_{t-1}\right]\right\}^2 \\ &= p_{1t}\left(\mu_{1t}^2 + h_{1t}\right) + (1-p_{1t}) \cdot \left(\mu_{2t}^2 + h_{2t}\right) - \left[p_{1t}\mu_{1t} + (1-p_{1t})\,\mu_{2t}\right]^2. \end{aligned} \tag{4}$$

The quantity h_t can be thought of as an aggregate of conditional variances from both regimes and now provides the basis for the specification of the regime-specific conditional variances $h_{it+1}, i = 1, 2$ in the form of parsimonious GARCH(1,1) models. However, instead of using a conventional GARCH(1,1) structure, we follow the econometric motivation by Dueker (1997) and adopt a slightly modified GARCH equation. For this, it is convenient to parameterize the degrees of freedom from the t-distribution (1) by $q = 1/\nu$, so that $(1 - 2q) = (\nu - 2)/\nu$, and to specify the alternative GARCH equation as:

$$h_{it} = b_{0i} + b_{1i}(1 - 2q_i)\epsilon_{t-1}^2 + b_{2i}h_{t-1} \tag{5}$$

with h_{t-1} as given according to Eq. (4), while ϵ_{t-1} is obtained from:

$$\begin{aligned} \epsilon_{t-1} &= R_{t-1} - E\left[R_{t-1}|\phi_{t-2}\right] \\ &= R_{t-1} - \left[p_{1t-1}\mu_{1t-1} + (1 - p_{1t-1})\,\mu_{2t-1}\right]. \end{aligned} \tag{6}$$

To close the model, it remains to specify the transition probabilities of the regime indicator S_t. For simplicity we consider a first order Markov process with constant transition probabilities, i.e. for $\pi_1, \pi_2 \in [0, 1]$ we define:

$$\begin{aligned} &\Pr\{S_t = 1|S_{t-1} = 1\} = \pi_1, \\ &\Pr\{S_t = 2|S_{t-1} = 1\} = 1 - \pi_1, \\ &\Pr\{S_t = 2|S_{t-1} = 2\} = \pi_2, \\ &\Pr\{S_t = 1|S_{t-1} = 2\} = 1 - \pi_2. \end{aligned} \tag{7}$$

Now, following Wilfling (2008), we obtain the log-likelihood function Λ of our Markov-switching-GARCH(1,1) model:

$$\begin{aligned} \Lambda &= \sum_{t=1}^{T} \log\left\{ p_{1t}\frac{\Gamma[(\nu_1+1)/2]}{\Gamma[\nu_1/2]\sqrt{\pi\nu_1 h_{1t}}}\left[1 + \frac{(R_t - \mu_{1t})^2}{h_{1t}\nu_1}\right]^{-(\nu_1+1)/2} \right. \\ &\left. +(1 - p_{1t})\frac{\Gamma[(\nu_2+1)/2]}{\Gamma[\nu_2/2]\sqrt{\pi\nu_2 \cdot h_{2t}}}\left[1 + \frac{(R_t - \mu_{2t})^2}{h_{2t}\nu_2}\right]^{-(\nu_2+1)/2} \right\}. \end{aligned} \tag{8}$$

The log-likelihood function (8) contains the *ex-ante* probabilities $p_{1i} = \Pr\{S_t = 1|\phi_{t-1}\}$. The whole series of *ex-ante* probabilities can be estimated recursively by

$$p_{1t} = \pi_1 \cdot \frac{f_{1t-1}p_{1t-1}}{f_{1t-1}p_{1t-1} + f_{2t-1}\,(1 - p_{1t-1})} + (1 - \pi_2) \cdot \frac{f_{2t-1}\,(1 - p_{1t-1})}{f_{1t-1}p_{1t-1} + f_{2t-1}\,(1 - p_{1t-1})}, \tag{9}$$

where f_{1t} and f_{2t} denote the $t_{\nu_1,\mu_{1t},h_{1t}}$- and $t_{\nu_2,\mu_{2t},h_{2t}}$-density functions from Eq. (1), each evaluated at $x = R_t$.

4 Empirical Results

Table 1 presents the maximum-likelihood estimates of the Markov-switching-GARCH model from the equations (1) to (9) for the WIG20, mWIG40 and TechWIG index returns. Maximization of the log-likelihood function was performed by the 'MAXIMIZE'-routine within the software package RATS 7.1 using the BFGS-algorithm, heteroscedasticity-consistent estimates of standard errors and suitably chosen starting values for all parameters involved. Overall, the majority of the coefficients of the mean and GARCH equations (3) and (5) are statistically significant at the 1% level in all estimations.

[Insert Table 1 about here]

Most autoregressive coefficients a_{11} and a_{12} are statistically significant and positive. A positive autoregressive structure of order one in stock index returns is an empirical finding often reported in the literature which can be explained by non-synchronous trading (Lo and MacKinlay (1990)), time-varying expected returns (Conrad and Kaul (1988)) and transaction costs (Mech (1993)). In contrast, the TechWIG exhibits a significant negative autoregressive coefficient in regime 1 while the autoregressive coefficient a_{11} is statistically insignificant for the WIG20. For all three stock market indices the coefficients of the lagged S&P500 index returns R_{t-1}^{SP} are statistically significant (at least at the 10% level) and positive in both regimes showing strong interdependence between US and Polish stock markets.

When looking at the estimated parameters describing the conditional volatility process we find the well-established result of volatility persistence in both regimes for all three datasets. The coefficient sums ($\hat{b}_{1i}(1 - 2\hat{q}_i) + \hat{b}_{2i}$ for $i = 1, 2$) do not exceed one confirming the stationary of conditional volatility processes in all regimes.

The constant transition probabilities π_1 and π_2 are close to one in all estimations. Since both quantities represent the probability of the data generating process to remain in the same volatility regime during the transition from date $t - 1$ to t, both volatility regimes reveal a high degree of persistence.

The lower part of Table I contains a diagnostic check of the model fit by providing Ljung-Box Q-statistics for serial correlation of the squared standardized residuals for the lags 1, 2, 3, 5, and 10. For the mWIG40 and the TechWIG the null hypothesis of no autocorrelation cannot be rejected up to lag 10 at any conventional significance level providing further econometric evidence in favour of our two-regime Markov-switching-GARCH specification. In contrast, for the WIG20 the Ljung-Box Q-statistics indicates serial correlation of the squared standardized residuals for the lags 1, 2, and 3. This result is caused by two extreme daily returns of the WIG20 index on the 28 and 29 October 1997 associated with the Asian crisis. Including an outlier dummy variable in the mean equation (3) to control for those two extreme daily returns does not affect the empirical results while no significant serial correlation of the squared standardized residuals is found up to lag 10. Summing up, the two-regime Markov-switching-GARCH specification properly captures the non-linear variance dynamics of the WIG20, mWIG40, and TechWIG index returns.

Next, we address two conditional probabilities which are relevant for detecting how often and at which dates the Polish stock market switched between the high and the low volatility regimes. First, the *ex-ante* probabilities $p_{1t} \equiv \Pr\{S_t = 1|\phi_{t-1}\}, t = 2, \ldots, T$, which can be estimated recursively via Eq. (9), and second, the so-called *smoothed* probabilities $\Pr\{S_t = 1|\phi_T\}, t = 1, \ldots, T$, which can be computed after model estimation by the use of filter techniques.[5] The *ex-ante* probabilities are useful in forecasting one-step-ahead regimes based on an information set which evolves over time. In our context, the *ex-ante* probabilities reflect current market perceptions of the one-step-ahead volatility regime, thus representing an adequate measure of stock market volatility sentiments. In contrast to this, the *smoothed* probabilities are based on the full sample-information set ϕ_T and thus provide a basis for inferring *ex post* if and when volatility regime switches have occurred in the sample.

Figures 3, 4, and 5 display the regime-1 probabilities (upper panels) along with the conditional variance processes (lower panels) estimated for the Markov-switching-GARCH model for the WIG20, the mWIG40, and the TechWIG. The *ex-ante* probabilities are represented by the thin lines while the bold lines depict the *smoothed* regime-1 probabilities. Since the *ex-ante* probabilities are determined by an evolving (and thus smaller) information set, they exhibit a more erratic dynamic behaviour than the *smoothed* regime-1 probabilities. In all panels the background is coloured in gray for the time period after the introduction of index futures trading. In all figures high probabilities (upper panels) are associated with high volatility periods (lower panels) this indicates the regime-1 is the high volatility regime.

According to the destabilization hypothesis uninformed investors induce noise in the price discovery process and lower the information content of prices a with the consequence

[5]The *smoothed* probabilities for the WIG20 index returns were computed on the basis of a filter algorithm provided by Gray (1996a).

of an increase in the spot market volatility. Presumably uninformed individual investors are the dominant trades in the Polish index futures markets. If the destabilization hypothesis is valid and individual investors are uninformed traders, we should see clear cut evidence in favour of a permanent increase of stock market volatility after the introduction of the index futures market segment. In terms of our Markov-switching-GARCH approach after the futures market introduction the processes of the spot market returns should transfer to regime-1 associated with a high level of conditional variances.

Looking at figure 3 for the WIG20 index, the regime-1 probability and the conditional variance indicate two periods during which the process is in the high volatility regime. The first of these two periods begins around February 1997 when all blue chip stocks contained in the WIG20 became continuously traded. High conditional variances at the end of 1997, mid 1998 and at the beginning of 1999 are associated with the Asian (October/November 1997), the Russian (August/September 1998) and the Brazilian (January 1999) crises, respectively. The high volatility period ends in May 1999. The second period (March 2000 to April 2001) in the regime-1 is clearly driven by bear markets following the burst of the 'dot-com bubble'. More importantly, Figures 3 demonstrate the impact of WIG20 index futures trading on the conditional volatility of the underlying stock market index. The introduction of WIG20 index futures on the 16 January 1998 falls into the high volatility regime and no immediate regime switch follows. From April 2001 to the end of the sample period, the process remains in the low volatility regime even though index futures trading has intensified substantially during this time period (see Figure 1). This empirical evidence is in contradiction to the hypothesis that uninformed individual investors on the index futures market introduce noisy price signals into the spot market and thus destabilise it.

[Insert Figure 3 about here]

Next, we investigate the regime-1 probability and the conditional variance for mWIG40 index (Figure 4). Three periods of high volatility emerge. As for the WIG20, the first volatile period, between end of 1997 and beginning of 1999, is related to the Asian, Russian, and Brazilian crises. The second period in regime-1 arises in form of two spikes of medium duration and stems from the collapsing 'dot-com bubble'. The mWIG40 index switches back to the high volatility regime at the end of 2005. No regime switch takes place around the introduction of index futures trading on 18 February 2002. We can conclude that the introduction of the mWIG40 index future did not destabilise the underlying stock market index.

[Insert Figure 4 about here]

Figure 5 shows that the TechWIG index starts in the high volatility regime at the beginning of 2000. Eight month after introduction of futures trading, around April 2001, a gradual transition towards the low volatility regime occurs. This movement coincides with the end of the financial turmoil following the burst of the 'dot-com bubble'. In contrast to the destabilization hypothesis this finding demonstrates potential stabilising effects of index futures trading on TechWIG stock price dynamics.

[Insert Figure 5 about here]

In this paper, we investigate the impact of the introduction of index futures trading on the conditional volatility of the underlying spot market. Since the Polish index futures market is dominated by individual investors, we hypothesise that derivatives traders are mostly uninformed. Consistent with this view, index futures trading should have increased spot market volatility. Surprisingly, the introduction of index futures does not destabilise the underlying cash markets. The observed switches between volatility regimes seem to be driven by other exogenous events. The empirical findings imply that the individuals transacting on the Polish derivatives market are less uninformed than the literature on individual investors' trading behaviour suggests.

5 Summary and Conclusions

The aim of this paper is to investigate the impact of the introduction of index futures trading in Poland on the conditional return volatility of the underlying stock index markets. Since the Polish index futures market is dominated by presumably uninformed individual traders we are able to investigate the destabilization hypothesis more directly than available studies do. However, index futures trading does not have a destabilising effect on the Polish stock market, implying that Polish individuals transacting in index futures are better informed and more rational traders than previous literature suggests for individuals. The introduction of index futures therefore might lead to better information flows into the cash market and hence to a more efficient spot market. This evidence does not justify the regulation of derivatives trading in order to stabilise the cash market.

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Figure 1: Index Futures Contracts Trading Volume and Open Interest

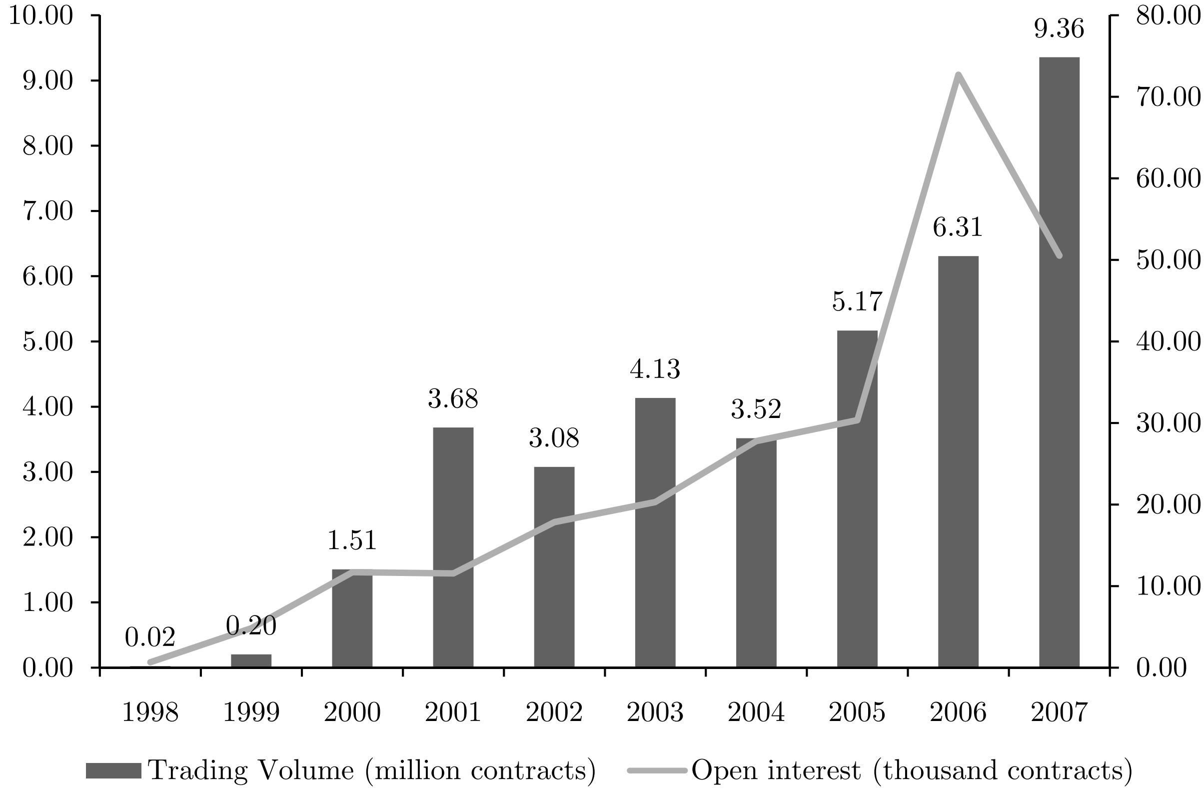
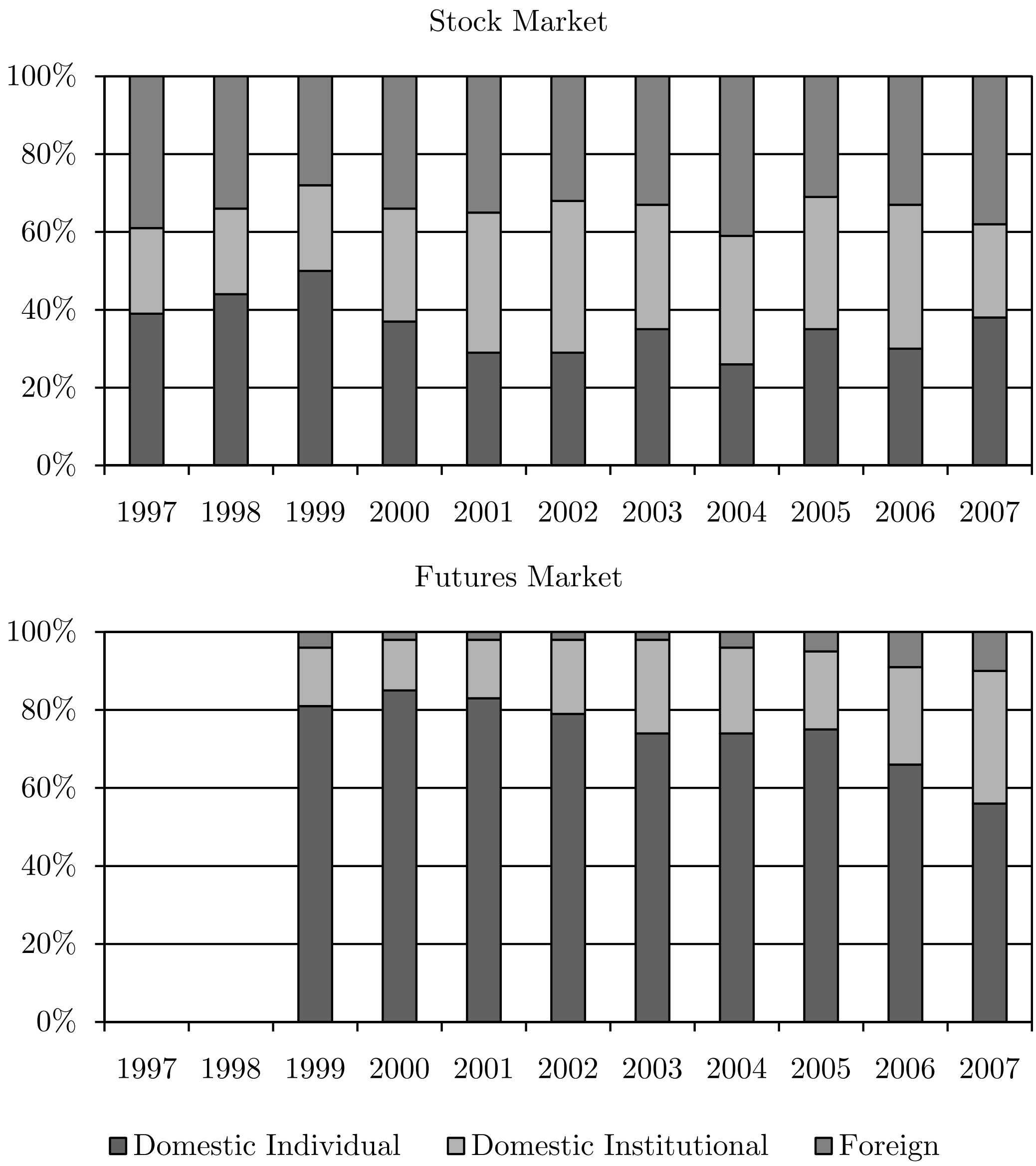


Figure 2: Shares in Trading of Different Investor Groups for WSE Spot and Futures Markets



Note: The figure graphs annual percentage shares in trading turnover of different investor groups for trading in the WSE spot and futures markets, respectively.

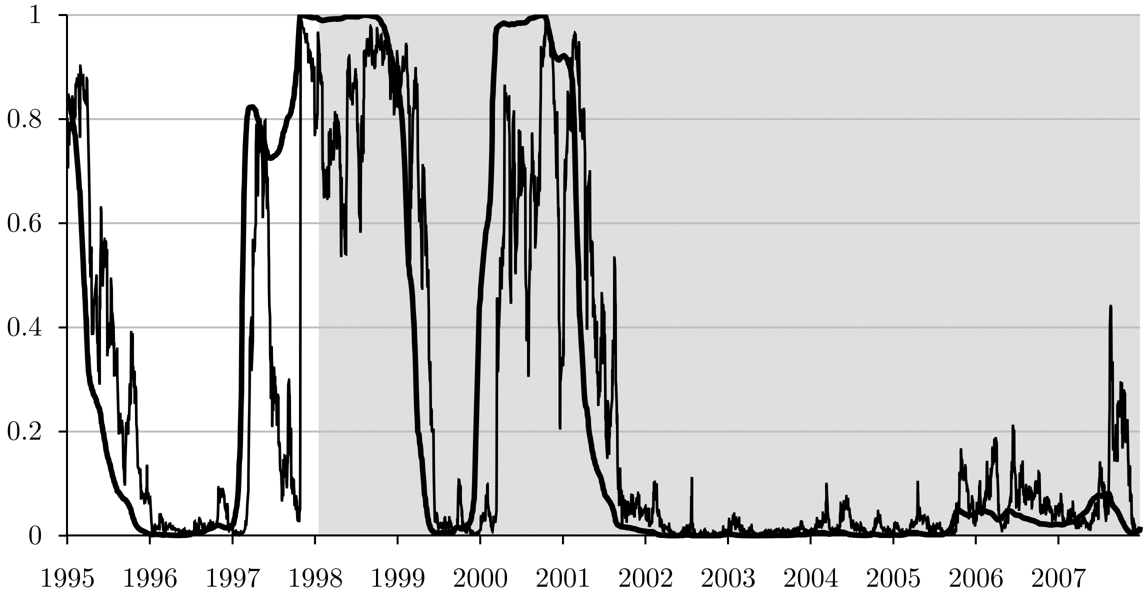
Table 1: Estimates and related statistics for Markov-switching-GARCH model

	WIG20		mWIG40		TechWIG	
Parameters	Estimate	p-value	Estimate	p-value	Estimate	p-value
Regime 1						
α_{01}	-0.0713	(0.382)	0.1744	(0.000)***	-0.1133	(0.078)*
α_{11}	-0.0548	(0.108)	0.0221	(0.068)*	-0.1121	(0.001)***
α_{21}	0.5502	(0.000)***	0.5671	(0.000)***	0.6794	(0.000)***
β_{01}	0.1885	(0.030)**	0.0262	(0.193)	0.0413	(0.725)
β_{11}	0.1811	(0.000)***	0.0878	(0.031)**	0.1459	(0.009)***
β_{21}	0.7789	(0.000)***	0.8883	(0.000)***	0.8489	(0.000)***
q_1	0.0960	(0.000)***	0.1428	(0.000)***	0.0500	(0.129)
$[\hat{\beta}_{11}(1-2\hat{q}_1)+\hat{\beta}_{21}]$	[0.9253]		[0.9510]		[0.9804]	
Regime 2						
α_{02}	0.0658	(0.018)**	0.0353	(0.068)*	0.0689	(0.024)**
α_{12}	0.0419	(0.036)**	0.1231	(0.000)***	0.0442	(0.065)*
α_{22}	0.2438	(0.000)***	0.1211	(0.000)***	0.2642	(0.000)***
β_{02}	0.0237	(0.149)	0.0268	(0.004)***	0.0152	(0.014)**
β_{12}	0.0446	(0.001)***	0.1412	(0.000)***	0.0583	(0.000)***
β_{22}	0.9386	(0.000)***	0.8000	(0.000)***	0.9295	(0.000)***
q_2	0.1016	(0.000)***	0.1040	(0.000)***	0.1583	(0.000)***
$[\hat{\beta}_{12}(1-2\hat{q}_2)+\hat{\beta}_{22}]$	[0.9741]		[0.9118]		[0.9695]	
Transition Probabilities						
π_1	0.9958	(0.000)***	0.9950	(0.000)***	0.9953	(0.000)***
π_2	0.9996	(0.000)***	0.9980	(0.000)***	0.9998	(0.000)***
Log-likelihood						
	-6249.4160		-3534.9395		-6254.9014	
Residual analysis	Estimate	p-value	Estimate	p-value	Estimate	p-value
LB_1^2	4.6379	(0.031)**	0.1885	(0.664)	0.0007	(0.933)
LB_2^2	4.6394	(0.098)*	0.8927	(0.640)	0.5152	(0.773)
LB_3^2	7.2663	(0.064)*	0.9149	(0.822)	0.8096	(0.847)
LB_5^2	7.5626	(0.182)	2.5063	(0.776)	3.3599	(0.645)
LB_{10}^2	12.8956	(0.230)	4.7305	(0.908)	6.3027	(0.789)

Notes: Estimates for parameters from equations (1)-(9). For the WIG20 the sample covers the 1 November 1994 to 31 December 2007 period. For the mWIG40 (TechWIG) the sample period starts from 21 September 1998 (31 December 1999) to 31 December 2007. LB_i^2 denotes the Ljung-Box Q-statistics for serial correlation of the squared standardized residuals up to i lags. $^{*},^{**},^{***}$ denote statistical significance at the 10%, 5% and 1% level, respectively.

Figure 3: Regime-1 probabilities and conditional variances (WIG20)

Regime-1 probabilities



Conditional variances

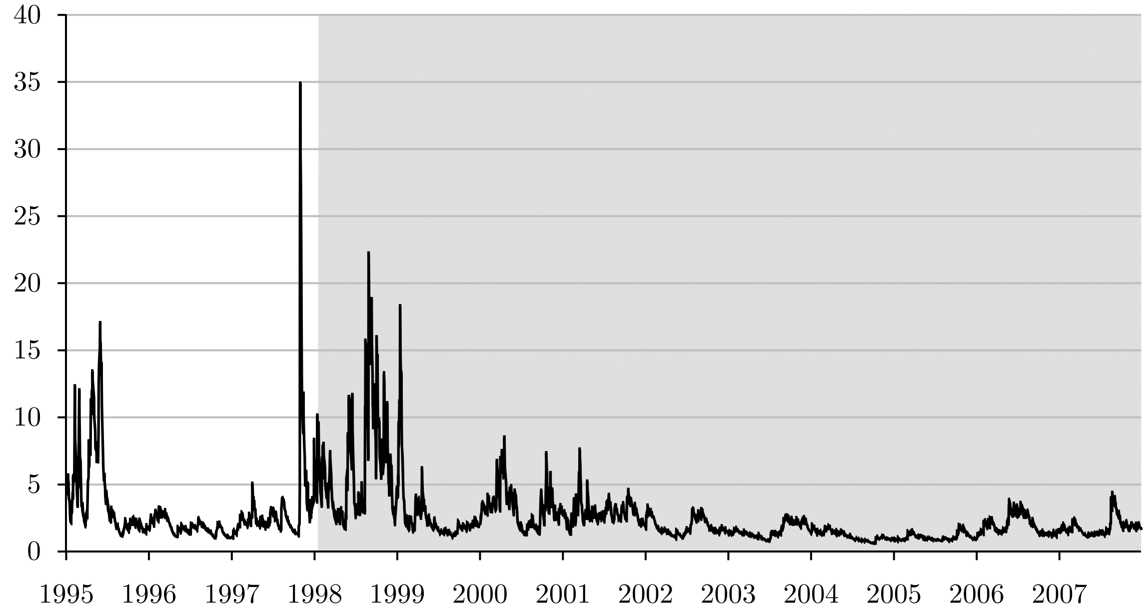


Figure 4: Regime-1 probabilities and conditional variances (mWIG40)

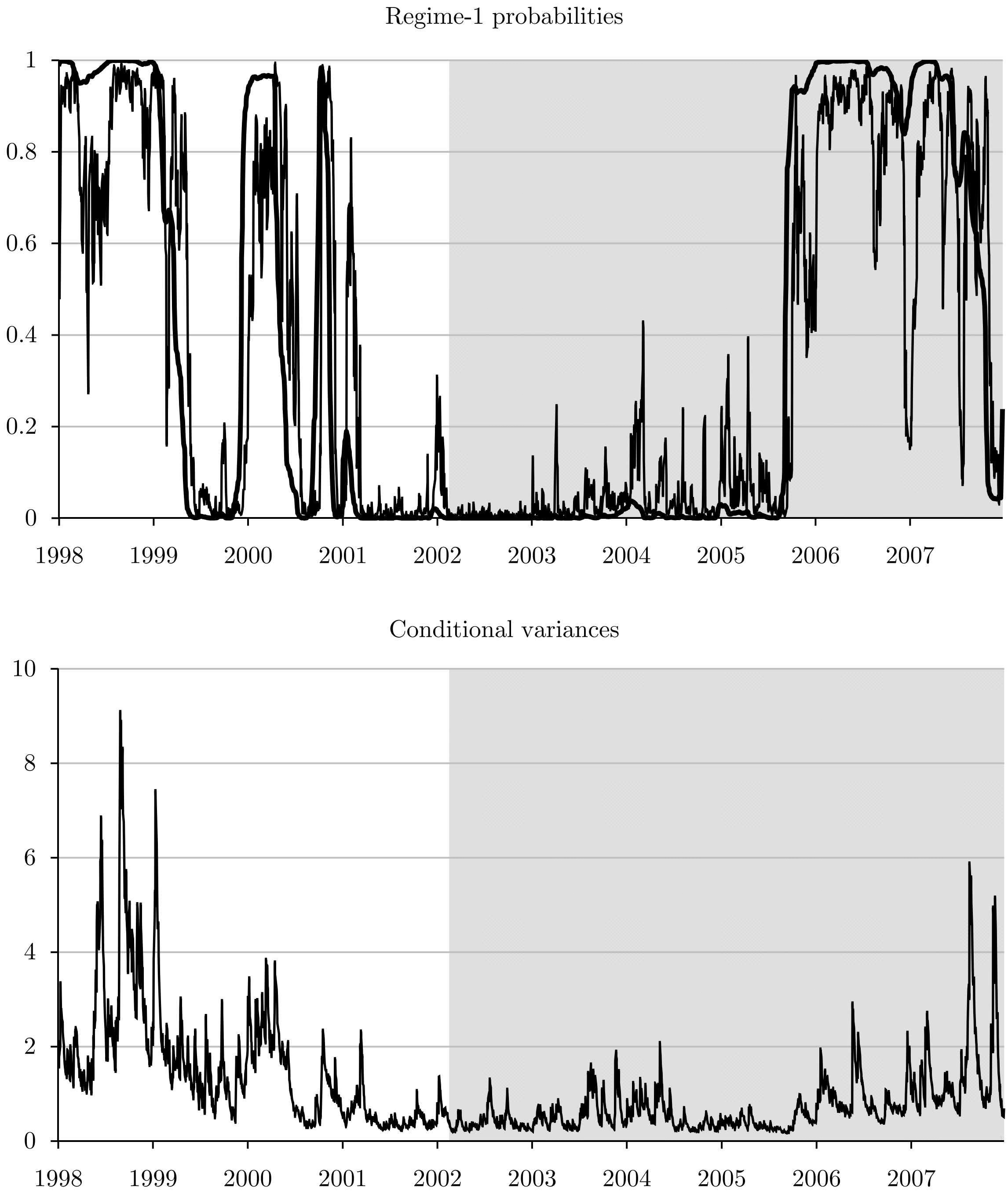


Figure 5: Regime-1 probabilities and conditional variances (TechWIG)