Price Discovery and Investor Structure in Stock Index Futures*

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Abstract

Previous literature on price discovery in stock index futures and spot markets neglects the role of different investor groups. This paper relates time-varying spot-futures linkages studied within a VECM-DCC-GARCH framework to changes in the investor structure of the futures market over time. Empirical results suggest that during the dominance of presumably uninformed private investors, the futures market does not contribute to price discovery. By contrast, there is evidence of information flows from futures to spot markets and a significant increase in conditional correlation between both markets as institutional investors' share in trading volume increases. We derive implications for the design of emerging futures markets.

JEL Classification: G10, G14, G15, G18

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1 Introduction

Since the introduction of stock index futures trading, extensive research has been devoted to the question whether index futures trading contributes to the efficiency of the underlying stock markets in terms of price discovery. Under frictionless markets, new information should be impounded simultaneously in futures and spot prices. In reality, however, futures markets are likely to incorporate market-wide information more efficiently than spot markets due to their inherent leverage, low transaction costs, and a lack of short-sale restrictions. A large body of literature looking at mature futures markets has confirmed that stock index futures typically lead the cash market, implying that a large part of price discovery takes place in the futures market, e.g. the US (Stoll and Whaley, 1990; Wahab and Lashgari, 1993; Koutmos and Tucker, 1996; Pizzi et al., 1998; Hasbrouck, 2003; Chou and Chung, 2006), the UK (Tse, 1999; Brooks et al., 2001), Japan (Iihara et al., 1996; Covrig et al., 2004) or Germany (Booth et al., 1999; Gaul and Theissen, 2008).

It is important to notice that the futures markets investigated in the literature so far are rather homogeneous in terms of their investor structure. Historically, the inception of futures trading in developed financial markets coincided with the rise of institutional ownership in the early 1980s. Thus, futures markets under investigation in previous literature are typically dominated by institutional investors. In the finance literature, institutions are usually presumed to be well-informed, rational investors whereas individuals are viewed as uninformed or driven by sentiment and behavioral biases (Lee et al., 1999; Cohen et al., 2002; Barber and Odean, 2008; Kaniel et al., 2008). In this paper, we ask whether the long-standing evidence of price discovery in futures markets is robust with respect to the investor structure. More specifically, we investigate whether the dominance of presumably unsophisticated individual investors in the futures market impairs the informational contribution of futures trading. If unsophisticated investors trade in a market, this may reduce the quality of the price signal and lower the information content of prices. In this setting, it is possible that the futures market may not perform its price discovery function.

The importance of different groups of investors in explaining patterns in asset prices and returns has long been recognized in various areas of financial research. For instance, Boehmer and Kelley (2009) show that the informational efficiency of transaction prices of individual NYSE-listed stocks increases in the share of institutional ownership. Moreover, previous literature relates seasonalities in daily stock returns to the trading behavior of individual versus institutional investors (Lakonishok and Maberly, 1990; Chan et al., 2004). Contributions by Gompers and Metrick (2001) and Phalippou (2008) highlight the role of institutional ownership in explaining cross-sectional return anomalies like the size and value effects. Even though recent literature on futures trading has noted peculiarities in emerging markets related to investor structure and behavioral biases (McMillan and Üklü, 2009), the price discovery dimension of this problem has been neglected so far.

We investigate this issue in the case of the Polish WIG20 index futures market, which offers a unique investor structure. While foreign and domestic institutional investors constantly account for about two thirds of spot trading volume, the futures market is dominated by domestic private individuals. These supposedly unsophisticated investors accounted for 75-80% of annual futures trading volume between 1998 and 2004. However, a change in mutual fund regulation in the fall of 2004 triggered a considerable increase in institutional stock index futures trading, causing the share of individual investors to decline to 53% by 2008. This shift in the investor structure of futures markets allows us to relate empirical evidence on price discovery to the changing market shares of different investor groups. As a consequence for our empirical investigation, we split the whole sample into a subperiod of individual investor dominance (1998 - 2004) as well as a subperiod of increased institutional trading (2005 - 2009).

Our empirical investigation uses a vector error correction model with a multivariate DCC-GARCH extension (VECM-DCC-GARCH). This setup allows us to jointly analyze lead-lag relationships in returns, volatility transmission and conditional correlation between the two markets. We find that under the dominance of presumably unsophisticated individual investors in the futures market, price discovery occurs mainly in the spot market, which is dominated by foreign and domestic institutional investors. By contrast, the growing influence of institutional investors on the futures market in the later period coincides with intensified information flows from futures to cash prices in terms of error correction and volatility spillovers. Moreover, the level of conditional correlation between futures and spot returns is significantly higher after the rise in institutional trading.

These findings are of interest for regulators in emerging markets, who face the decision of whether or not to introduce stock index futures. Our results indicate that efficiency gains expected from futures trading critically hinge on the sophistication of market participants. Allowing informed institutional investors to trade futures contracts may improve the quality of the price signal emerging from this market. This will allow futures trading to perform the desired price discovery function.

The plan of the paper is as follows. Sections 2 and 3 present institutional details of the WIG20 index futures markets and our empirical approach, respectively. Our dataset is described in Section 4, while estimation results are discussed in Section 5. Section 6 summarizes our findings and concludes.

2 Institutional Details

The Warsaw Stock Exchange (WSE) dates back to 1817 when a first stock exchange was founded in Poland.[1] Warsaw grew to become the most important stock market in Poland by the end of the 1930s, when it closed due to the beginning of WWII. Only after the demise of communism in 1989, economic transition lead to a reopening of the exchange on April 16, 1991. Since then, the WSE has emerged as the largest stock market in all Central and Eastern European (CEE) countries. As of 2008, total market capitalization of the WSE has reached USD 91 billion, which is far more than other Eastern European markets like Budapest (USD 18 billion) and Ljubljana (USD 12 billion). In fact, its size rivals that of smaller Western European exchanges such as Vienna (76 billion USD) or Luxembourg (67 billion USD).

Initially, there was a spot market only. On 16 January 1998, about seven years after the reopening of the WSE, the first futures contract on the blue-chip index WIG20 was introduced. This contract is not only the oldest in the Polish futures market, it is also the most popular by far, accounting for 93% of total trading volume in all derivatives in 2008. With a volume of 11.77 million contracts traded and a notional value of USD 124,526 million, total index futures trading at the WSE in 2008 is comparable to Western European markets such as the Borsa Italia (7.82 million contracts; USD 1,136,457 million)

[1] Information on institutional detail of the WSE, like trading volume, contract specifications and investor structure can be found in various issues of the WSE Annual Fact Book. The comparisons of the WSE with other exchanges are based on World Federation of Exchanges Annual Report 2008.

or the MEFF in Spain (10.58 million contracts; USD 1,318,295 million). At the same time, the Polish stock index futures market is considerably larger compared to those of other CEE countries, as for instance that of Budapest (3.60 million contracts; USD 4,268 million) and even those of some Western European exchanges such as Athens (2.83 million contracts; USD 37,604 million) or Vienna (0.24 million contracts; USD 19,356 million). The WIG20 index futures contract considered in this study, as well as the constituents of the underlying WIG20 index, are traded in the continuous trading system of the WSE.

Most relevant to our research question is the unique investor structure on the Polish spot and futures market. As can be seen from Figure 1, there is an astonishing difference in investor structure between spot and futures markets at the WSE. During the last 10 years, the spot turnover shares of domestic individual, domestic institutional and foreign investors have been equal and relatively stable.

[Insert Figure 1 about here]

By contrast, the futures market is clearly dominated by domestic individual traders who accounted for 75 - 80 % of trading volume during the 1998 - 2004 period. During recent years, individual investors' proportion of futures trading volume has considerably declined, reaching a low of 53 % in 2008. By the same token, foreign and domestic insititutional trading has nearly doubled. Part of the observed change in the investor structure at the WSE futures market can be rationalized by a reform in the regulatory framework of the Polish financial market. Originally, domestic mutual funds were not permitted to invest in derivatives. A change in legislation, i.e. a new Act of Investment Funds coming into effect in the Fall of 2004, opened the index futures market to this class of institutional investors. At the same time, a generally more liberal open-end mutual fund regulation policy triggered a rapid growth of the number of mutual funds in the Polish market, beginning only in the following year of 2005 (Bank of Poland, 2005). Moreover, the new Act allowed foreign investment funds to operate directly in the Polish market. These institutional changes led to a massive growth of mutual funds' total assets under management, which quadrupled between 2005 and 2007 (Analizy Online, 2008). Consequently, motivated by the observed change in the investor structure and the reform in the regulatory framework, we split the sample into a subperiod of individual investor dominance (1998 - 2004) as well as a subperiod of increased institutional trading (2005 - 2009).

Private investors are often perceived as less sophisticated, whereas informed trading is usually attributed to professional institutional investors. Therefore, the observed change in the investor structure of the futures market, i.e. the initial dominance of presumably uninformed individual investors and the rise in institutional trading after 2005, provides a promising setting for an investigation of price discovery and investor structure.

3 Model and Empirical Approach

Simple arbitrage arguments such as the cost-of-carry model imply that the logarithms of futures, f_t, and spot prices, s_t, are cointegrated with a common stochastic trend.[2] The cointegrating relationship between the two price series can be represented as

$$f_t = \alpha_0 + \alpha_1 s_t \ . \tag{1}$$

Intuitively, market frictions will cause temporal deviations from this equilibrium relationship. The dynamics of the cointegrated logarithms of futures, f_t, and spot prices, s_t, can be studied using a bivariate vector error correction model (VECM) of the following form

$$\Delta s_t = \beta_{s,0} + \gamma_s ec_{t-1} + \sum_{j=1}^{p} \beta_{ss,j} \Delta s_{t-j} + \sum_{j=1}^{q} \beta_{sf,j} \Delta f_{t-j} + \varepsilon_{s,t} \tag{2}$$

$$\Delta f_t = \beta_{f,0} + \gamma_f ec_{t-1} + \sum_{j=1}^{p} \beta_{fs,j} \Delta s_{t-j} + \sum_{j=1}^{q} \beta_{ff,j} \Delta f_{t-j} + \varepsilon_{f,t} \tag{3}$$

where $ec_t = f_t - \alpha_0 - \alpha_1 s_t$ is the estimated error correction term.

This framework accounts for different aspects of the short-term and long-term relationship between the two variables. The error correction coefficients, γ_s and γ_f, measure the speed of adjustment in response to deviations from the long-run equilibrium. If for instance the futures price is high relative to its equilibrium value, a negative price change

[2] In what follows, we abstract from time-varying interest rates and dividends as well as the time to maturity of the futures contract.

on the futures and/or a positive price change on the spot market will correct the mispricing. Therefore, we generally expect $\gamma_f < 0$ and $\gamma_s > 0$. However, is is also possible that the error correction coefficient of the spot market has the same sign as the error correction coefficient of the futures market.[3] As pointed out by Zhong et al. (2004), two opposing effects determine the sign of the error correction coefficient in the spot equation (γ_s). Suppose the spot index is below its equilibrium value ($ec_t > 0$). This may induce arbitrageurs to buy the constituting stocks of the index, leading to a price increase. Equilibrium would be restored due to the arbitrage effect, implying a positive sign on the error correction term. However, unlike the futures contract, the index is not a traded asset itself but rather a weighted average of individual assets. In fact, some constituting stocks may depreciate even more due to short-term momentum. In this case, the index may deviate even further from equilibrium. Thus the momentum effect implies that the sign on the error correction term in the spot equation can be negative.[4] Moreover, the short-term predictive power of one variable for the other is captured by coefficients $\beta_{sf,i}$ and $\beta_{fs,i}$. Finally, the coefficients measuring the reaction of spot and futures returns to their own lagged values ($\beta_{ss,i}, \beta_{ff,i}$) indicate the degree of mean-reverting behavior of both time series.

We employ two distinct approaches to characterize interaction in futures and spot returns. First, the concept of Granger causality can be used to formally test for unidirectional causality or feedback relationships in returns. For spot prices to Granger-cause futures prices for instance, past spot prices must have some information content for predicting futures prices. Such causality can run either through long-term (error correction, γ_f) or short-term dependencies (lagged returns, $\beta_{fs,i}$). More specifically, we test the following null hypotheses: $H_0 : \beta_{fs,1} = \beta_{fs,2} = ... = \beta_{fs,p} = 0$ and $H_0 : \gamma_f = 0$. In case we can reject at least one of them, this is interpreted as evidence of Granger causality running from spot to futures prices. Similarly, futures prices are said to Granger-cause spot prices if either $H_0 : \beta_{sf,1} = \beta_{sf,2} = ... = \beta_{sf,q} = 0$ or $H_0 : \gamma_s = 0$ can be rejected. Note that causality must exist in at least one direction for spot and futures prices to be cointegrated.

[3] Given that γ_f is negative, the error correction coefficient of the spot market does not need to be strictly positive. To restore the long-run equilibrium it is only required that the $\gamma_s < \gamma_f$ in absolute values.

[4] Tswei and Lai (2009) argue that equal signs on error correction coefficients can be explained by overreaction to news in the less informative futures market and under-reaction in the more informative spot market.

If both variables Granger-cause each other, there is bidirectional Granger causality, which implies that the price discovery process is interdependent. Granger causality tests can be conducted by performing t-tests on error correction coefficients and F-tests on the joint significance of the sum of the lags of each variable.

Second, to quantify the relative price discovery contribution of both markets, we use a simple measure based on the error correction coefficients γ_s and γ_f proposed by Schwarz and Szakmary (1994). Specifically, the common factor weights of futures and spot markets can be written as[5]

$$\theta_f = \frac{\gamma_s}{\gamma_s + |\gamma_f|} \quad \text{and} \quad \theta_s = 1 - \theta_f = \frac{|\gamma_f|}{\gamma_s + |\gamma_f|} \ . \tag{4}$$

If the price discovery process takes place exclusively in the futures market, $\theta_f = 1$. Conversely, if price discovery occurs in the spot market only, $\theta_f = 0$. The intuition is that the market with the lower speed of adjustment coefficient does not follow but rather initiates the mispricing, implying that the price discovery process takes place mainly in this market. It is important to notice that this approach is closely related to other popular measures used to assess the relative contributions of different markets to price discovery such as the common factor weights of Gonzalo and Granger (1995) and information shares developed by Hasbrouck (1995).[6] Comparing the three approaches, Theissen (2002) shows that the Schwarz and Szakmary (1994) measure can be derived from the Gonzalo-Granger framework and conclusions are qualitatively similar to those based on information shares.

Chan et al. (1991) first highlighted the role of spillover effects between conditional volatilities as an indicator of price discovery. We account for such potential linkages in second moments using a multivariate generalized autoregressive conditional heteroskedasticity (GARCH) modeling framework. Define the variance-covariance matrix of spot and

[5]Equation 4 applies to the case where $\gamma_f < 0$ and $\gamma_s > 0$. Note that the absolute value operator is not required if the speed of adjustment coefficients have the same sign.

[6]Hasbrouck (1995) defines price discovery in terms of the variance of all innovations in a VECM to the common factor. Gonzalo and Granger (1995) consider only permanent shocks where each markets contribution to the common factor is defined to be a function of only the error correction coefficient in a VECM.

futures residuals conditional on the information set at point in time t (denoted Ψ_t) as

$$var(\epsilon_{s,t}, \epsilon_{f,t}|\Psi_{t-1}) \equiv H_t = \begin{bmatrix} h_{ss,t} \; h_{sf,t} \\ h_{sf,t} \; h_{ff,t} \end{bmatrix} . \tag{5}$$

We assume that conditional variances follow autoregressive processes of the following form

$$h_{ss,t} = \omega_s + \delta_{s,1} h_{ss,t-1} + \delta_{s,2} \epsilon^2_{s,t-1} + \delta_{s,3} \; \epsilon^2_{s,t-1} \; I_{s,t} + \delta_{s,f} \; \epsilon^2_{f,t-1} + \delta_{s,ec} \; ec^2_{t-1} \tag{6}$$

$$h_{ff,t} = \omega_f + \delta_{f,1} h_{ff,t-1} + \delta_{f,2} \epsilon^2_{f,t-1} + \delta_{f,3} \; \epsilon^2_{f,t-1} \; I_{f,t} + \delta_{f,s} \; \epsilon^2_{s,t-1} + \delta_{f,ec} \; ec^2_{t-1} \; . \tag{7}$$

This specification accounts for various effects on conditional volatility. First, we allow for asymmetric reactions to positive and negative return innovations by including a multiplicative dummy term where the indicator variable $I_{i,t}$ takes on the value of 1 if $\epsilon_{i,t-1} < 0$. Second, the error correction term is included. This is because we are interested in whether deviations from the equilibrium relationship of futures and cash prices have an effect on the variance of either of the two series. Intuitively, in the event of disequilibria, trading in both the spot and futures markets will intensify as arbitrageurs seek to exploit the temporary mispricing. Intensified trading activities will cause higher volatility in both markets. Thus, the sign of the error correction terms in both variance equations is expected to be positive. Third, cross-spillover effects between the two markets are captured by coefficients $\delta_{s,f}$ and $\delta_{f,s}$, respectively. These can be interpreted in terms of information flows between markets (Chan et al., 1991; Tse, 1999; Kavussanos and Visvikis, 2004). If, for instance, $\delta_{s,f}$ is significant, we can reject the null hypothesis of no volatility spillovers from futures to spot prices. This indicates information flows from the futures to the spot market and implies a dominant role of the futures market in price discovery.

Various approaches to modeling conditional covariances have been proposed in the literature.[7] We follow Engle (2002) in allowing conditional correlation to vary over time.

[7]See Bauwens et al. (2006) for a survey. An example of modeling spot and futures prices in a DCC framework is Lien and Yang (2006). Comparable empirical approaches, simultaneously accounting for error correction, volatility spillovers and conditional correlation have been employed to study prices discovery (Tse, 1999; Zhong et al., 2004; So and Tse, 2004; Kavussanos and Visvikis, 2004) as well as in the context of hedging (Lien and Tse, 1998; Brooks et al., 2002; Lien and Yang, 2006) or predictability (Brooks et al., 2001).

His Dynamic Conditional Correlation (DCC) model ensures the positive definiteness of H_t under simple conditions on the parameters, while not restricting correlations to be time-invariant as in the Constant Conditional Correlation approach of Bollerslev (1990). Thus it provides us with a direct measure of time-varying correlation between futures and cash markets. Let $\rho_{ij,t}$ denote the conditional correlation coefficient between spot index and futures prices. The variance-covariance matrix of residuals can then be rewritten as

$$H_t = \begin{bmatrix} h_{ss,t} & \rho_{sf,t}\sqrt{h_{ss,t}h_{ff,t}} \\ \rho_{sf,t}\sqrt{h_{ss,t}h_{jff,t}} & h_{ff,t} \end{bmatrix} = D_t R_t D_t \tag{8}$$

where $D_t = diag(h_{11,t}^{1/2}...h_{22,t}^{1/2})$ is the diagonal matrix of conditional standard deviations. The matrix of conditional correlations (R_t) is given as[8]

$$R_t = diag(q_{11,t}^{1/2}...q_{22,t}^{1/2})\ Q_t\ diag(q_{11,t}^{1/2}...q_{22,t}^{1/2})\ . \tag{9}$$

The matrix Q_t depends on squared standardized residuals ($u_{i,t} = \varepsilon_{i,t}/\sqrt{h_{ii,t}}$), their unconditional variance-covariance matrix ($\bar{Q}$) and its own lagged value[9]

$$Q_t = (1 - \kappa_1 - \kappa_2)\bar{Q} + \kappa_1 u_{t-1}u'_{t-1} + \kappa_2 Q_{t-1} \tag{10}$$

with $\kappa_1, \kappa_2 > 0$ and $\kappa_1 + \kappa_2 < 1$.

Previous empirical finance literature views institutional investors as rational and well-informed whereas individual investors are characterized as irrational, less informed, and driven by behavioral biases. Taking into account these arguments and the specific institutional setting of the Polish spot and futures markets, we posit the following two hypotheses about return and volatility interactions and price discovery. First, before 2005, due to the dominance of uninformed private investors, the price signal emerging from futures markets is presumably noisy. This will limit the contribution of futures trading to price discovery. Therefore, we expect futures prices to not unidirectionally Granger-cause spot prices, a low common factor weight for the futures market and insignificant futures-spot volatility

[8]The diagonal matrices in Equation 9 ensure that Q_t is a well-defined correlation matrix.

[9]Note that in empirical applications, the $\bar{Q}$ is replaced by its sample analogue.

transmission during the 1998 - 2004 period. Second, after 2005, legislation allows and encourages well-informed (domestic and foreign) institutional investors to trade in index futures. Due to low transaction costs, they will use this instrument when trading on private information. Therefore, we expect the information content of futures prices to increase and thus price discovery to move to futures markets, at least to some extent, during the 2005 - 2009 sample period. This will be reflected in an increase in the common factor weight of the futures market as well as significant volatility spillovers from futures to spot returns.

4 Data

Our dataset consists of daily close prices for the WIG20 index and daily settlement prices for the WIG20 futures contract. Both time series are taken from Thomson Financial Datastream. The delivery months for the WIG20 futures contract are 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. As is common for index futures, trading volume is highest for the contract that is closest to maturity. This is true until the beginning of the delivery month, when trading activity shifts to the next most immediate maturity. Moreover, volatility of futures prices for the closest-to-maturity contract typically increases sharply during this time period. Therefore, when constructing our time-series of index futures quotes, we only consider observations for the most immediate contract except for the expiry month.

The WIG20 index and the corresponding futures price are expressed in natural logarithms by $s_t = log(S_t)$ and $f_t = log(F_t)$, respectively. Contiunously compounded daily returns are calculated as

$$\Delta s_t = (s_t - s_{t-1}) \tag{11}$$

$$\Delta f_t = (f_t - f_{t-1}) \ . \tag{12}$$

Table 1 presents descriptive statistics of prices and returns series for the WIG20 index and the corresponding index futures time series. Panel A of Table 1 contains details regarding the distributional characteristics whereas Panel B concerns intertemporal dependencies.

[Insert Table 1 about here]

Prices and returns series in cash and futures markets exhibit similarity in terms of means, standard deviations as well as minimums and maximums. The sample distributions of index and index futures returns are skewed left and leptokurtic. The Jarque-Bera test statistics clearly rejects the null hypothesis of normality for spot and futures returns. We also apply the Ljung-Box Q test for serial correlation to returns and squared returns. The Ljung-Box test statistics for 5 and 20 lags are significant for returns and squared returns, suggesting the presence of significant serial correlation and heteroskedasticity. Moreover, judging from the size of the test statistic, nonlinear dependencies seem to be far more prevalent. Similarly, the Engle (1982) test of autoregressive conditional heteroskedasticity (ARCH) clearly rejects the null of no such effects in both index and futures returns.

We also examine the stationarity of the time series using the Augmented Dickey-Fuller (ADF), the Phillips-Perron (PP) and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test. The results are reported in Panel A of Table 2. The index spot and futures prices appear to be first difference stationary.

[Insert Table 2 about here]

The results of Johansen cointegration tests are presented in Panel B of Table 2. Both the trace and the eigenvalues tests strongly reject the absence of cointegration but do not reject the existence of one cointegrating relation between spot and futures prices. Panel B also reports estimated coefficients of the cointegration equation. The coefficient of the spot price is very close to -1. Likelihood ratio test statistics for $a_0 = 0$, $a_1 = -1$, as well as the joint hypothesis $a_0 = 0$ and $a_1 = -1$, are not significant. These results suggest that the basis spread could serve as a representative of the cointegrating relation between spot and futures prices. Our results are robust whether we use the basis spread or the estimated error correction term in our model.

5 Empirical Results

5.1 Baseline Results on Price Discovery

We choose to limit the lag-length for past changes of futures and spot index prices in the VECM to 2. Comparisons with higher order models based on Schwarz's Bayesian Information Criterion favor this parsimonious specification. Tests for autocorrelation of standardized residuals confirm model adequacy. The model is estimated via Quasi Maximum Likelihood with robust standard errors based on the procedure suggested by Bollerslev and Wooldridge (1992).

Our baseline estimations take into account the full sample period from January 16, 1998 to June 30, 2009. Results are summarized in Table 3. As shown in Panel A, the coefficient on lagged spot index changes in the spot equation ($\beta_{ss,1}$) is significant at the 1% level and has a negative sign. The corresponding coefficient in the futures equation ($\beta_{ff,1}$) is also negative but not statistically different from 0 at conventional levels of significance. We conclude that, in contrast to futures prices, the index exhibits mean reversion. By contrast, coefficients on second order lags $\beta_{ss,2}$ and $\beta_{ff,2}$ are insignificant.

[Insert Table 3 about here]

In order to investigate Granger causality, we first test joint hypotheses on the return interaction terms. While we cannot reject the hypothesis that the impact of spot on futures returns is zero ($H_0 : \beta_{fs,1} = \beta_{fs,2} = 0$), a similar test for the spot equation easily rejects the null hypothesis of no cross-effects at the one percent level of significance ($H_0 : \beta_{sf,1} = \beta_{sf,2} = 0$). Moreover, coefficients on error correction terms in both equations are statistically different from 0 at conventional levels. We find γ_f to be slightly larger than γ_s in absolute terms, which implies that futures returns exhibit a stronger reaction to departures from long-run equilibrium than does the spot index. In sum, we find bidirectional Granger causality between futures and spot prices.[10]

[10] As a robustness check we conduct Granger causality tests using the procedure proposed by Toda and Yamamoto (1995). Consistent with the results from the Granger causality tests based on the VECM, we find a bidirectional Granger causality between the spot and the futures market.

The negative sign of both speed of adjustment coefficients is consistent with standard market mechanisms determining the reaction of futures and spot prices to equilibrium perturbations. Intuitively, if the futures price is higher than implied by the cointegration relationship ($ec_t > 0$), rational arbitrageurs will go short in the future, causing (γ_f) to be negative. The significantly parameter estimate of γ_s implies that momentum dominates the arbitrage effect, which is in line with evidence reported by Zhong et al. (2004) for Mexico.

As a complement to Granger causality tests and inference on error correction terms, common factor weights provide us with a quantitative measure of the relative price discovery contributions of both markets. The low value of θ_f reported in Table 3 suggests that the spot market dominates in the process of price discovery. In sum, based on our results regarding common factor weights and Granger causality, we can conclude that price discovery to a large part occurs in the spot market. This finding is in contrast to long-standing evidence for mature markets (e.g. Wahab and Lashgari, 1993; Koutmos and Tucker, 1996; Pizzi et al., 1998).

As can be seen from significant point estimates of $\delta_{s,1}$ and $\delta_{f,1}$, there is volatility clustering in both spot index and future returns. Squared past return innovations also have a significant effect, which is stronger if the surprise in returns was negative. The coefficient measuring such asymmetries in spot (futures) volatility, $\delta_{s,3}$ ($\delta_{f,3}$), is statistically different from 0 at the 1% level of significance. Such asymmetric reactions of the conditional variance of stock returns are usually interpreted as evidence for the leverage effect of Black (1976). Note that these asymmetries are less significant for the futures contract. Intuitively, as Koutmos and Tucker (1996) note, a positive response of conditional return volatility to market drops is more difficult to rationalize in the case of futures compared to cash stock markets. We also investigate the volatility effects of the cointegrating relationship. Our results indicate that the conditional variance of spot returns is higher, the more futures and spot prices deviate from their long-run equilibrium relationship. Zhong et al. (2004) report similar results for Mexico. Even though this result may imply that futures trading has an impact on the volatility of the underlying (via the cointegrating relationship), the effect is small in magnitude.

Most importantly, as discussed in Section 3, volatility transmission can be interpreted in terms of information flow between futures and cash markets (Chan et al., 1991; Tse, 1999;

Zhong et al., 2004; So and Tse, 2004). Thus analyzing interactions between the conditional variances of both return series provides us with another measure of price discovery. Indeed we find some evidence of a volatility spillover from the futures to the cash market. As evinced by a significant estimate of δ_{sf} in Table 3, squared innovations in futures returns appear to affect spot market volatility.

Finally, our empirical setup allows for dynamic conditional correlation between both markets within the modeling framework of Engle (2002). Panel C of Table 3 reports the corresponding coefficients, which are different from zero at the five and one percent level of significance, respectively.

Summing up, estimation results for the total 1998 - 2009 sample period suggest that the futures market does not fully perform the expected price discovery function. This result is in contrast to previous literature for developed financial markets. Mixed evidence regarding return and volatility interaction between spot and futures returns may in fact be due to confounded effects of informed trading moving between both markets. Most notably, our baseline estimations neglect massive changes in the investor structure of the futures market during the 2005 - 2009 period (see Section 2). In the next step, we therefore try to shed light on the relationship between price discovery and investor structure over time.

5.2 Price Discovery, Information Flows and Changes in the Investor Structure of the Futures Market

In order to analyze the potential relationship between price discovery and investor structure, we reestimate our model for two subsamples, running from the inception of futures trading until December 31, 2004 and from January 3, 2005 until the end of the full sample period. This choice is motivated by regulatory regime shifts in 2004, triggering a change in the investor structure of the derivatives market from 2005 onwards. As discussed in Section 2, new legislation (1) allowed mutual to invest in index futures and (2) spurred a rapid growth of their assets under management during recent years. This has led to a pronounced increase in the share of institutional investors' futures trading volume. Comparing parameter estimates across sample periods therefore allows us to relate findings on price discovery to changes in the investor structure.

Tables 4 and 5 summarize our estimation results. The evidence regarding price discovery and information flows remarkably differs across the two subsamples. Several detailed findings bear mention. For the first subsample, there is only weak evidence of short-term interactions as the coefficients in the futures (spot) equation are jointly insignificant (only marginally significant). Moreover, the speed of adjustment coefficient in the spot equation (γ_s) is indistinguishable from 0 while the corresponding parameter in the futures equation (γ_f) is highly significant. This implies that during the 1998 - 2004 sample period, only the futures price appears to react to temporary deviations from the long-term relationship of both markets. Correspondingly, the common factor weight of the futures market, θ_f, is low relative to the estimate for the full sample period and new information is impounded in spot prices first. Thus in line with our initial hypothesis, the futures market, dominated by uninformed domestic private investors, contributes little to price discovery during the 1998 - 2004 period.

After 2005, by contrast, the point estimate of γ_s is statistically significant at the 5% level. This implies that the spot market tends to react more to deviations from long-term equilibrium as presumably well-informed institutional investors begin to use futures contracts to trade on private information. As discussed above, a momentum effect may dominate these adjustments and result in a negative sign on γ_s. At the same time, the common factor weight for the futures market more than doubles compared to the earlier subsample. Taken together, our estimates of error correction parameters suggest that the price discovery contribution of the futures market rises sharply during the period of increased institutional trading.

A comparison of volatility spillovers between futures and spot markets across both sample periods supports the notion of institutional trading increasing the informational contribution of futures markets. For the earlier subsample, the coefficients $\delta_{s,f}$ and $\delta_{f,s}$ are not statistically different from zero at conventional levels of significance. By contrast, we find a significant point estimate of $\delta_{s,f}$ for the 2005 - 2009 period, indicating considerable information flows from futures to spot markets. This is again consistent with our hypothesis that the increasing share of well-informed institutional investors in the futures market after 2005 has improved the information content of futures prices.

Summing up, the contribution of the WIG20 index futures market to price discovery in the Polish stock market has increased considerably in recent years. One interpretation is that these stronger linkages reflect market participants' learning about the arbitrage mechanism linking both markets. More likely, this result can be related to the simultaneous change in the investor structure.

[Insert Tables 4 and 5 about here]

Finally, we explore potential effects of institutional regime shifts on the comovement of both markets. Figure 2 plots estimated conditional correlations coefficients for our full sample period. Visual inspection reveals that the correlation between futures prices and the underlying appears to gradually increase over time. Additional estimations of our VECM-GARCH-DCC model for mature stock index futures such as the French CAC40 or the Dow Jones Industrial Average show that the overall level of conditional correlation for the WIG20 is comparably low, especially during earlier years.[11]

[Insert Figure 2 about here]

In order to quantify the effect of institutional regime shifts, we regress the conditional correlation coefficient of innovations in log futures and spot index changes on two dummies capturing the changing investor structure after 2004 and, for robustness, the introduction of WARSET in 2000. Table 6 reports coefficient estimates. Evidently, the introduction of the computerized trading system as well as the change in the investor structure have led to an increase in conditional correlation between futures and cash markets. Ten years after the inception of WIG20 futures trading, both the investor structure and the level of conditional correlation with the underlying are now more similar to mature markets.

[Insert Table 6 about here]

[11]Estimation results for mature stock market index futures are available from the authors on request.

6 Conclusions

Previous literature has extensively analyzed price discovery and lead-lag relationships between stock index futures and spot markets (for example Stoll and Whaley, 1990; Wahab and Lashgari, 1993; Booth et al., 1999; Chou and Chung, 2006; Covrig et al., 2004; Gaul and Theissen, 2008). There is widespread evidence that futures trading contributes to price discovery and thus to the efficiency of stock markets. It is important to notice that most of the markets under investigation in existing studies are similar in terms of investor structure in that the share of institutional trading is high. In fact, previous literature neglects potential effects of investor structure, especially dominant individual investors, on the relative price discovery contributions of stock index futures and spot markets.

This paper aims at filling this void. We study information flows between the Polish blue-chip index WIG20 and its futures contract traded at the Warsaw Stock Exchange and relate our results to changes in the investor structure of the futures market. During the 1998 - 2004 period, presumably uninformed private investors account for about 80 % of futures trading volume. For this sample period, we find that price discovery occurs mainly in the spot market whereas the contribution of the futures market remains rather low. After 2005, however, the futures market experiences a steep increase in the trading volume share of foreign and domestic institutional investors due to changes in regulation. Our empirical results for the 2005 - 2009 subsample indicate substantial information flows from futures to spot prices, mirrored by an increased common factor weight of the futures market as well as significant unidirectional volatility transmission. This suggests that a growing fraction of institutional trading increases the information content of futures prices. Moreover, we document that the shift in the investor structure has also led to a significant increase in conditional correlation between both markets.

These results highlight an important link between the role of derivatives markets in the price formation process and the investor structure in these markets. In line with previous literature, we find some evidence that futures markets perform a price discovery function and thus contribute to market efficiency. However, our findings also suggest that this beneficial effect critically hinges on the participation of well-informed institutional investors. This has important implications for emerging market regulators who face decisions regarding the design of domestic futures markets. While restricting market participation of institutions like pension funds may be popular with the public, it may also compromise the efficiency of domestic derivatives markets. Therefore, to reap the fruits of index futures trading in terms of increased stock market efficiency, regulators should not prohibit but rather encourage institutional futures trading.

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Table 1: Descriptive Statistics of WIG20 Index and Futures Prices and Returns.

Panel A: Distribution

	Mean	Median	Maximum	Minimum	Std. dev.	Skewness	Kurtosis	Jarque-Bera
Index prices	7.5235	7.4658	8.2733	6.8786	0.3582	–	–	–
Futures prices	7.5209	7.4668	8.2822	6.7935	0.3619	–	–	–
Index returns	0.0094	0.0112	9.0006	−10.4491	1.8673	−0.1835	5.4985	762.5967***
Futures returns	0.0097	0.0000	10.6646	−10.7631	1.9522	−0.05	5.7200	885.8718***

Panel B: Autocorrelations

	Lag 1 - 5 autocorrelations					$Q(5)$	$Q(20)$	$ARCH(20)$
	1	2	3	4	5			
Index returns	0.0142	0.0222	−0.0179	0.0260	−0.0166	5.629	24.537***	24.576***
Squared index returns	0.1568	0.1705	0.2367	0.2272	0.2088	589.097***	991.158***	
Future returns	−0.0034	0.0121	0.0158	0.0096	−0.0086	1.659	25.714***	22.506***
Squared future returns	0.1326	0.1301	0.1822	0.1795	0.1948	396.311***	811.358***	

Note: All prices are in natural logarithms and returns are the first differences of log prices denoted in percentage points. $Q(5)$ ($Q(20)$) is the Ljung-Box Q test of serial correlation of up to 5 (20) lags. $ARCH(20)$ is the F-test of autoregressive conditional heteroskedasticity effect with 20 lags proposed by Engle(1982). The sample period is January 16, 1998 to June 30, 2009 (2870 observations). *,** ,*** denote statistical significance at the 10%, 5% and 1% level, respectively.

Table 2: Unit Root Tests and Cointegration Analysis of WIG20 Index and Futures Returns.

Panel A: Unit root tests

	ADF test		PP test		KPSS test	
	Intercept	Intercept and trend	Intercept	Intercept and trend	Intercept	Intercept and trend
Index prices	−1.495	−1.554	−1.514	−1.587	3.420***	0.636***
Futures prices	−1.547	−1.618	−1.566	−1.654	3.445***	0.597***
	None	Intercept	None	Intercept	Intercept	
Index returns	−52.798***	−52.790***	−52.797***	−52.789***	0.113	
Futures returns	−53.735***	−53.727***	−53.736***	−53.728***	0.108	

Panel B: Cointegration analysis

Johansen cointegration tests

Null hypothesis	Maximum eigenvalues statistics	95% CV for the maximum eigenvalues test	Trace statistics	95% CV for the trace test
None ($r = 0$)	78.7336***	15.67	81.2033***	19.96
At most one ($r = 1$)	2.4697	9.24	2.4697	9.42

Estimated cointegration vector

$f_t = \alpha_0 + \alpha_1 s_t + ec_t$		Restriction tests		
α_0	α_1	$\alpha_0 = 0$ $[\chi^2(1)]$	$\alpha_1 = -1$ $[\chi^2(1)]$	$\alpha_0 = 0$ and $\alpha_1 = 1$ $[\chi^2(2)]$
0.0750	−1.0096***	2.429	2.282	3.408

Note: All prices are in natural logarithms and returns are the first differences of log prices. The ADF test refers to the Augmented-Dickey-Fuller test, the PP test to the Phillips-Perron test, and the KPSS test to the Kwiatkowski-Phillips-Schmidt-Shin test. Lag length selection in all tests is based on Schwarz's Information Criterion (SBIC). We use a second-order vector autoregressive model for the Johansen (1991) cointegration test and restriction tests. Critical values (CV) of the Johansen cointegration tests come from Osterwald-Lenum (1992). The sample period is January 16, 1998 to June 30, 2009. *,** ,*** denote statistical significance at the 10%, 5% and 1% level, respectively.

Table 3: Estimation Results for WIG20 Index and Index Futures, 1998-2009.

Dependent variable	Parameter estimates (p-values)					
	Panel A: $\Delta s_t = \beta_{s,0} + \gamma_s ec_{t-1} + \sum_{j=1}^{p} \beta_{ss,j}\Delta s_{t-j} + \sum_{j=1}^{q} \beta_{sf,j}\Delta f_{t-j} + \varepsilon_{s,t}$ $\Delta f_t = \beta_{f,0} + \gamma_f ec_{t-1} + \sum_{j=1}^{p} \beta_{fs,j}\Delta s_{t-j} + \sum_{j=1}^{q} \beta_{ff,j}\Delta f_{t-j} + \varepsilon_{f,t}$					
Δs_t	$\beta_{s,0}$	γ_s	$\beta_{ss,1}$	$\beta_{ss,2}$	$\beta_{sf,1}$	$\beta_{sf,2}$
	0.315 (0.193)	−0.033 (0.035)**	−0.190 (0.000)***	−0.034 (0.384)	0.219 (0.000)***	0.045 (0.237)
Δf_t	$\beta_{f,0}$	γ_f	$\beta_{fs,1}$	$\beta_{fs,2}$	$\beta_{ff,1}$	$\beta_{ff,2}$
	0.414 (0.106)	−0.090 (0.000)***	0.037 (0.368)	0.074 (0.076)*	−0.032 (0.426)	−0.059 (0.140)
	Restriction tests					
	$\beta_{sf,1} = \beta_{sf,2} = 0$ $[\chi^2(2)]$		$\beta_{fs,1} = \beta_{fs,2} = 0$ $[\chi^2(2)]$		θ_f	
	22.114***		2.344		0.27	
	Panel B: $h_{ss,t} = \omega_s + \delta_{s,1} h_{ss,t-1} + \delta_{s,2}\epsilon^2_{s,t-1} + \delta_{s,3}\,\epsilon^2_{s,t-1}\,I_{s,t} + \delta_{s,f}\,\epsilon^2_{f,t-1} + \delta_{s,ec}\,ec^2_{t-1}$ $h_{ff,t} = \omega_f + \delta_{f,1} h_{ff,t-1} + \delta_{f,2}\epsilon^2_{f,t-1} + \delta_{f,3}\,\epsilon^2_{f,t-1}\,I_{f,t} + \delta_{f,s}\,\epsilon^2_{s,t-1} + \delta_{f,ec}\,ec^2_{t-1}$					
$h_{ss,t}$	ω_s	$\delta_{s,1}$	$\delta_{s,2}$	$\delta_{s,3}$	$\delta_{s,f}$	$\delta_{s,ec}$
	0.004 (0.000)***	0.900 (0.000)***	0.039 (0.000)***	0.040 (0.000)***	0.014 (0.047)**	0.005 (0.002)***
$h_{ff,t}$	ω_f	$\delta_{f,1}$	$\delta_{f,2}$	$\delta_{f,3}$	$\delta_{f,s}$	$\delta_{f,ec}$
	0.004 (0.000)***	0.918 (0.000)***	0.051 (0.000)***	0.025 (0.002)***	0.008 (0.225)	0.002 (0.125)
	Panel C: $Q_t = (1-\kappa_1-\kappa_2)\bar{Q} + \kappa_1 u_{t-1}u'_{t-1} + \kappa_2 Q_{t-1}$					
Q_t	κ_1	κ_2				
	0.019 (0.013)**	0.979 (0.000)***				

Note: Intercepts in the mean ($\beta_{s,0}, \beta_{f,0}$) and conditional volatility equations (ω_s, ω_f) are multiplied by 10^4. The sample period is January 16, 1998 - June 30, 2009. Robust Bollerslev and Wooldridge (1992) p-values are in parentheses.

Table 4: Estimation Results for WIG20 Index and Index Futures, 1998-2004.

Dependent variable	Parameter estimates (p-values)					
	Panel A:	$\Delta s_t = \beta_{s,0} + \gamma_s ec_{t-1} + \sum_{j=1}^{p} \beta_{ss,j}\Delta s_{t-j} + \sum_{j=1}^{q} \beta_{sf,j}\Delta f_{t-j} + \varepsilon_{s,t}$ $\Delta f_t = \beta_{f,0} + \gamma_f ec_{t-1} + \sum_{j=1}^{p} \beta_{fs,j}\Delta s_{t-j} + \sum_{j=1}^{q} \beta_{ff,j}\Delta f_{t-j} + \varepsilon_{f,t}$				
Δs_t	$\beta_{s,0}$	γ_s	$\beta_{ss,1}$	$\beta_{ss,2}$	$\beta_{sf,1}$	$\beta_{sf,2}$
	0.295 (0.358)	−0.012 (0.502)	−0.090 (0.007)***	−0.007 (0.873)	0.124 (0.000)***	0.032 (0.425)
Δf_t	$\beta_{f,0}$	γ_f	$\beta_{fs,1}$	$\beta_{fs,2}$	$\beta_{ff,1}$	$\beta_{ff,2}$
	0.571 (0.089)*	−0.063 (0.000)***	0.067 (0.050)**	0.062 (0.163)	−0.060 (0.071)*	−0.027 (0.514)
	Restriction tests					
	$\beta_{sf,1} = \beta_{sf,2} = 0$ [$\chi^2(2)$]		$\beta_{fs,1} = \beta_{fs,2} = 0$ [$\chi^2(2)$]		θ_f	
	5.142*		2.098		0.16	
	Panel B:	$h_{ss,t} = \omega_s + \delta_{s,1} h_{ss,t-1} + \delta_{s,2}\epsilon^2_{s,t-1} + \delta_{s,3}\ \epsilon^2_{s,t-1}\ I_{s,t} + \delta_{s,f}\ \epsilon^2_{f,t-1} + \delta_{s,ec}\ ec^2_{t-1}$ $h_{ff,t} = \omega_f + \delta_{f,1} h_{ff,t-1} + \delta_{f,2}\epsilon^2_{f,t-1} + \delta_{f,3}\ \epsilon^2_{f,t-1}\ I_{f,t} + \delta_{f,s}\ \epsilon^2_{s,t-1} + \delta_{f,ec}\ ec^2_{t-1}$				
$h_{ss,t}$	ω_s	$\delta_{s,1}$	$\delta_{s,2}$	$\delta_{s,3}$	$\delta_{s,f}$	$\delta_{s,ec}$
	0.005 (0.099)*	0.901 (0.000)***	0.038 (0.000)***	0.044 (0.077)*	0.013 (0.356)	0.006 (0.049)**
$h_{ff,t}$	ω_f	$\delta_{f,1}$	$\delta_{f,2}$	$\delta_{f,3}$	$\delta_{f,s}$	$\delta_{f,ec}$
	0.005 (0.021)**	0.906 (0.000)***	0.058 (0.000)***	0.031 (0.077)*	0.005 (0.665)	0.002 (0.222)
	Panel C:	$Q_t = (1-\kappa_1-\kappa_2)\bar{Q} + \kappa_1 u_{t-1}u'_{t-1} + \kappa_2 Q_{t-1}$				
Q_t	κ_1	κ_2				
	0.015 (0.000)***	0.984 (0.000)***				

Note: Intercepts in the mean ($\beta_{s,0}, \beta_{f,0}$) and conditional volatility equations (ω_s, ω_f) are multiplied by 10^4. The sample period is January 16, 1998 - December 31, 2004. Robust Bollerslev and Wooldridge (1992) p-values are in parentheses.

Table 5: Estimation Results for WIG20 Index and Index Futures, 2005-2009.

Dependent variable	Parameter estimates (p-values)					
	Panel A: $\Delta s_t = \beta_{s,0} + \gamma_s ec_{t-1} + \sum_{j=1}^{p} \beta_{ss,j} \Delta s_{t-j} + \sum_{j=1}^{q} \beta_{sf,j} \Delta f_{t-j} + \varepsilon_{s,t}$ $\Delta f_t = \beta_{f,0} + \gamma_f ec_{t-1} + \sum_{j=1}^{p} \beta_{fs,j} \Delta s_{t-j} + \sum_{j=1}^{q} \beta_{ff,j} \Delta f_{t-j} + \varepsilon_{f,t}$					
Δs_t	$\beta_{s,0}$	γ_s	$\beta_{ss,1}$	$\beta_{ss,2}$	$\beta_{sf,1}$	$\beta_{sf,2}$
	0.208 (0.575)	−0.120 (0.020)**	−0.385 (0.000)***	−0.069 (0.432)	0.400 (0.000)***	0.061 (0.471)
Δf_t	$\beta_{f,0}$	γ_f	$\beta_{fs,1}$	$\beta_{fs,2}$	$\beta_{ff,1}$	$\beta_{ff,2}$
	0.102 (0.801)	−0.226 (0.000)***	−0.0874 (0.313)	0.0928 (0.318)	0.078 (0.337)	−0.104 (0.251)
	Restriction tests					
	$\beta_{sf,1} = \beta_{sf,2} = 0$ $[\chi^2(2)]$		$\beta_{fs,1} = \beta_{fs,2} = 0$ $[\chi^2(2)]$		θ_f	
	16.998***		3.074		0.35	
	Panel B: $h_{ss,t} = \omega_s + \delta_{s,1} h_{ss,t-1} + \delta_{s,2} \epsilon^2_{s,t-1} + \delta_{s,3}\ \epsilon^2_{s,t-1}\ I_{s,t} + \delta_{s,f}\ \epsilon^2_{f,t-1} + \delta_{s,ec}\ ec^2_{t-1}$ $h_{ff,t} = \omega_f + \delta_{f,1} h_{ff,t-1} + \delta_{f,2} \epsilon^2_{f,t-1} + \delta_{f,3}\ \epsilon^2_{f,t-1}\ I_{f,t} + \delta_{f,s}\ \epsilon^2_{s,t-1} + \delta_{f,ec}\ ec^2_{t-1}$					
$h_{ss,t}$	ω_s	$\delta_{s,1}$	$\delta_{s,2}$	$\delta_{s,3}$	$\delta_{s,f}$	$\delta_{s,ec}$
	0.009 (0.002)***	0.870 (0.000)***	0.040 (0.000)***	0.031 (0.051)*	0.029 (0.056)*	0.029 (0.081)*
$h_{ff,t}$	ω_f	$\delta_{f,1}$	$\delta_{f,2}$	$\delta_{f,3}$	$\delta_{f,s}$	$\delta_{f,ec}$
	0.005 (0.019)**	0.914 (0.000)***	0.053 (0.001)***	0.017 (0.203)	0.001 (0.881)	0.035 (0.003)***
	Panel C: $Q_t = (1 - \kappa_1 - \kappa_2)\bar{Q} + \kappa_1 u_{t-1} u'_{t-1} + \kappa_2 Q_{t-1}$					
Q_t	κ_1	κ_2				
	0.048 (0.005)***	0.690 (0.000)***				

Note: Intercepts in the mean $(\beta_{s,0}, \beta_{f,0})$ and conditional volatility equations (ω_s, ω_f) are multiplied by 10^4. The sample period is January 3, 2005 - June 30, 2009. Robust Bollerslev and Wooldridge (1992) p-values are in parentheses.

Table 6: Dynamic Conditional Correlation and Changes in the Institutional Setting.

ζ_0	ζ_1	ζ_2
0.900	0.029	0.019
(0.000)***	(0.000)***	(0.000)***

Note: The estimated regression is

$$\rho_{sf,t} = \zeta_0 + \zeta_1\ D_t^1 + \zeta_2\ D_t^2 + e_t\ ,$$

where the dummy variable D_t^1 (D_t^2) is zero until November 17, 2000 (January 3, 2005) and takes on the value of one afterwards. $\rho_{sf,t}$ is the time series of dynamic conditional correlation coefficients obtained from Maximum Likelihood estimation of the model outlined in Equations 2 through 10. The sample period is January 16, 1998 - June 30, 2009 (2870 observations).

Figure 1: Shares in Trading Turnover of Different Investor Groups for WSE Stock and Futures Markets.

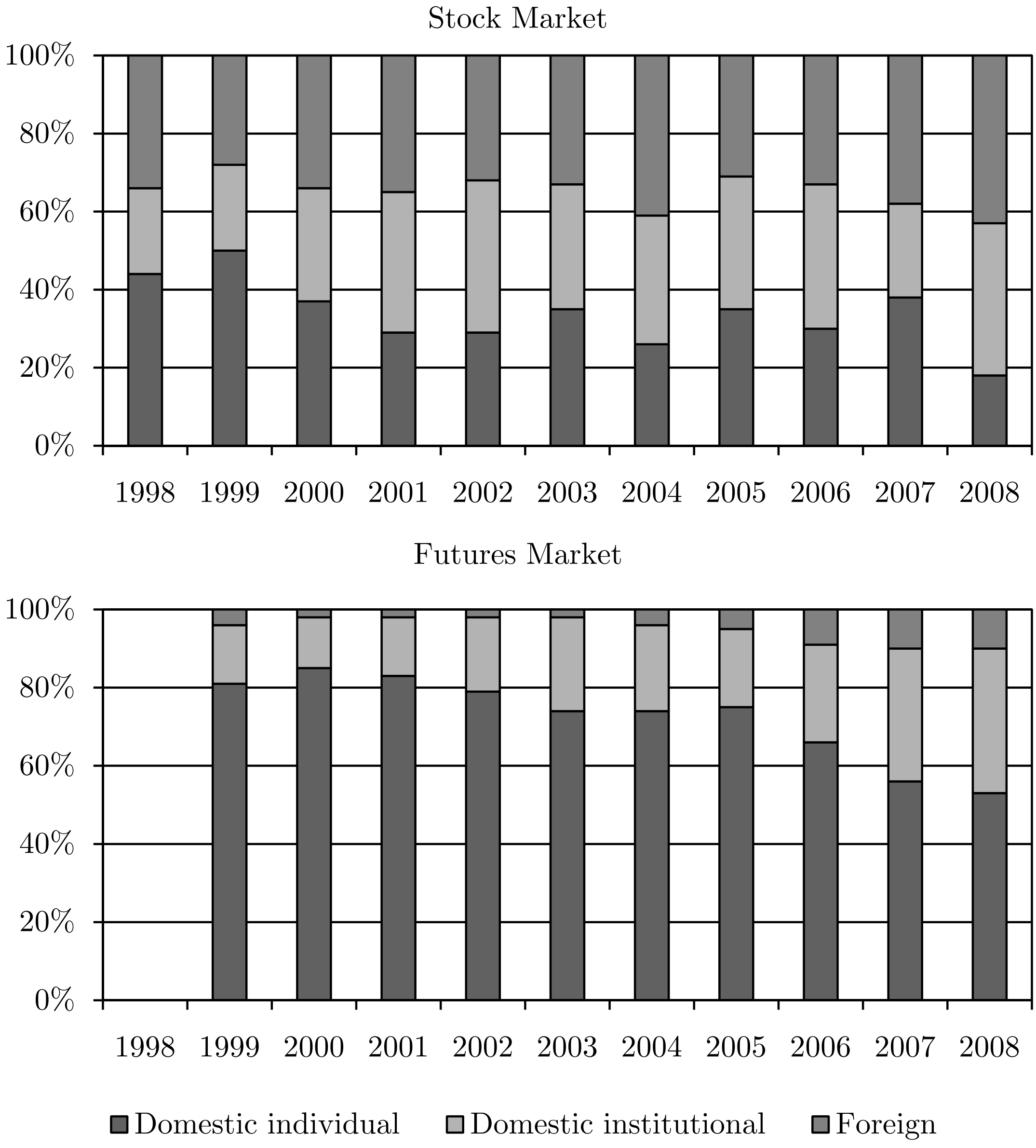

Figure 2: Conditional Correlation Between Stock Index Futures and Spot Market.

Note: The figure shows conditional correlation coefficients estimated from the model described in Section 3, Equations 2 through 10. The sample period is January 16, 1998 - June 30, 2009.