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Euro, dollar or Swiss franc: which currency had the greatest impact on the Hungarian, Polish and Czech economies during the global financial crisis?

Agata Kliber,[a] Piotr Płuciennik[b]

Abstract. The article presents an analysis of the impact of foreign currency dynamics on the fundamentals (basic indices of the economic performance) of the Czech Republic, Hungary and Poland during the financial crisis of 2007/2008 and its aftermath until 2017. The subject of the analysis are three currencies: the US dollar, the euro and the Swiss franc. The assessment of their impact on the fundamentals of the three above-mentioned economies is based on the joint volatilities of bond spreads and currencies. A series of copula-GARCH models was estimated. The research demonstrates that the impact of foreign currencies was the strongest in the case of Poland and Hungary, as these two countries were more dependent on loans in foreign currencies than the Czech Republic. Another finding shows that the impact decreased significantly in Hungary after its government introduced loan conversion.

Keywords: bond spread, copula-GARCH model, debt crisis, Central Europe

JEL: C32, C51, G01, G15

1. Introduction

The importance of exchange rates to the functioning of the whole economy is unquestionable. Exchange rates link the domestic economy of a given country with the international economy. Nominal exchange rates determine real exchange rates. The real exchange rate is an important factor influencing macroeconomic stability (see: Williamson, 2009). This article aims to analyse the impact of foreign currency dynamics on the fundamentals (most basic indices of economic performance, including GDP, inflation, interest rates, government credibility, etc.) of selected Central European economies: the Czech Republic, Hungary and Poland, during the financial crisis of 2007/2008 and its aftermath until 2017. All the countries are European Union members which retained their national currencies until the end of the studied period. All of them had floating exchange rate regimes at the beginning of the studied period. In November 2013, the Czech Republic decided to change the regime of its currency into an 'other managed arrangements' group (International Monetary Fund [IMF], 2014).

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The condition of the fundamentals was measured through the spreads of the bond yields to the safest economy in the region, i.e. Germany. Data was taken daily and covered the period of 2008–2014. Its source was the Thomson Reuters Datastream and the Stooq.pl portal.

It is a well-known fact that the dynamics of Central European currencies is strongly affected by the EUR-USD dynamics (see e.g. Doman, M., 2009). When analysing effective interest rates, it becomes clear that the dynamics is indeed composed mainly of euro and dollar FX rates. However, as Polish and Hungarian households have been heavily indebted in the Swiss franc (CHF), we assumed that this currency could also have affected these economies. Therefore, the impact of the three currencies (USD, EUR and CHF) on the fundamentals of the three V-4 economies,[1] which retained their own currencies during the studied period, was taken into account in this study. Its aim is then to check whether the impact of the Swiss franc might have been stronger than or at least as strong as the impact of the euro and the US dollar, despite having been a marginal part of the effective rates.

The following research hypotheses were formulated:

- The influence of the euro and the US dollar should be greater than the influence of the Swiss franc, since the effective exchange rates are composed mainly of euros and dollars;
- The impact of the Swiss franc on the fundamentals of the Polish and Hungarian economies should be greater than on the Czech economy;
- The replacement of a free-floating exchange rate regime in the Czech Republic with a managed-floating regime should permanently weaken the relationship between bond spreads and foreign exchange rates;
- The reforms implemented in Hungary, especially the obligatory conversion of loans from the foreign-currency-denominated to the forint-denominated, were likely to contribute to the weakening of the relationship between the Hungarian fundamentals and the exchange rates.

The article has the following structure: the first part presents the dynamics of the sovereign spreads of Polish, Hungarian and Czech bonds to those of German bonds, and the dynamics of the respective exchange rates. It also shows descriptive statistics and the changes of the variables in the context of the economic situation of the countries. The second part contains a description of the model used to demonstrate the influence of the exchange rates on the fundamentals of the economies under study. The last section is devoted to a discussion on the findings in relation to the domestic policies of the countries.

[1] V4 denotes Visegrad economies: the Czech Republic, Hungary, Poland, Slovakia.

2. Literature review

The issue of dependence between exchange rates has been extensively described in the literature. Rebitzky (2010) analysed articles on this subject written from 1990, so over 30 years, and according to his findings, most researchers agree that momentous news has a significant influence on exchange rates. Yet, some other researchers show that there are exceptions – Engel & West (2005), for example, demonstrate that fundamentals (such as relative money supplies, outputs, inflation and interest rates) do not necessarily improve floating exchange rate forecasts, while the opposite relationships hold. Nevertheless, researchers agree that relationships between exchange rates and fundamentals indeed exist and are statistically significant. Indicators describing the fundamentals are published monthly or even less frequently, which makes it problematic to include them in the model together with the daily quoted exchange rates without any loss of information on the latter.

For this reason, researchers approximate the condition of fundamentals with either sovereign credit default swaps (sCDS) series or bond spreads, i.e. the difference between the yield of domestic government bonds to the yield of the bond considered the safest in the region. In this paper, the condition of fundamentals has been assessed through the dynamics of the spreads of 10-years' sovereign bonds yields compared to the yields of 10-years' German bonds. Numerous researchers indicate that sovereign bonds are significantly more sensitive to the domestic condition of the economy than the alternative measure of sovereign risk – the spreads of sCDS. For instance, Matei & Cheptea (2012), who analysed spreads of European bonds against the German ones, demonstrated that large fiscal deficits and public debt, as well as political risks and, to a lesser extent, liquidity, are likely to put substantial upward pressures on sovereign bond yields in many advanced European economies. Kocsis (2014) proved that in the case of sCDS, global and regional factors can be clearly derived, yet no such factors exist in the case of domestic bond markets. In his opinion, bond spreads reflect different monetary policies or an overall domestic policy, which differs across countries. The idiosyncratic factor can explain even up to 80% of the variance of bond yields (e.g. in Hungary, where in the case of sCDS this figure is only 33%). Claeys & Vašiček (2014) indicated that the movements of bond spreads anticipate changes in ratings prepared by credit rating agencies.

Considering the above, the condition of fundamentals was measured through the spreads of the bond yields against the safest economy in the region, i.e. Germany.

Our aim was to assess how the exchange rate market affected the fundamentals of the selected Central and East Europe (CEE) economies over the years 2008–2017, during which very important international events occurred. First of all, the beginning of the period saw the outbreak of an international financial crisis which spread across Europe. Simultaneously, following an international growth risk, speculators attacked Central European currencies (2008/2009). Kliber & Kliber (2010) showed that in 2008, when the speculative attacks took place, the CEE currencies were strongly affected by the common fluctuations of the EUR-PLN and the EUR-HUF exchange rates. In 2009, investors who up to that point had used to paint the CEE currencies with a broad brush, seemed to have started noticing the differences among them (see for instance: Kliber, 2009). During the crisis, the internal situation of Hungary was declining, while the economies of Poland and the Czech Republic got affected by the crisis only to a limited extent (see e.g. Kliber & Płuciennik, 2015, 2017; Komárková et al., 2013). However, both Polish and Hungarian households were massively indebted in Swiss franc, as prior to the crisis a lot of people had taken mortgage loans in this currency. This phenomenon did not occur on such a large scale in the Czech Republic, though.[2]

3. The data

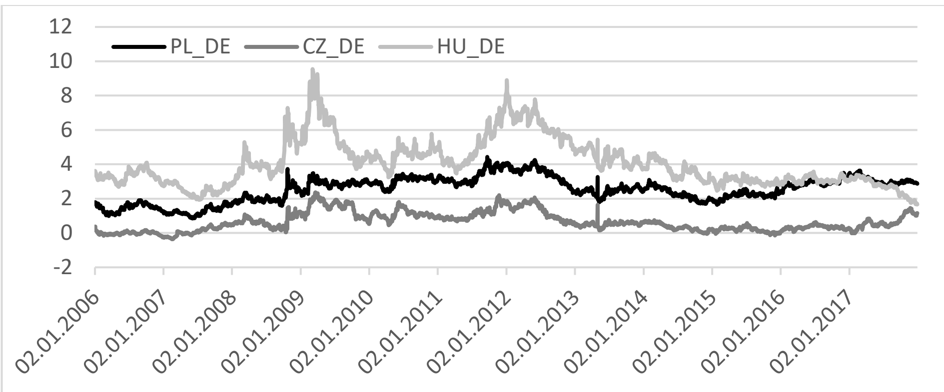
Figure 1 presents the dynamics of the CEE yields against the German yields. All the bonds were of a 10-years' maturity. Spreads are interpreted as measures of risk of a given country against the safest one in the region. The highest value of the spread was observed in Hungary. In fact, it was this country that was affected by the crisis to the largest extent of all the three analysed countries. The value of the Polish spread was lower than the Hungarian one, but higher than the Czech one. The dynamics of the Polish and Czech spreads were similar, and different from the dynamics of the Hungarian spread.

Two peaks were observed in the Hungarian data: the first one occurred in March 2009, and the second in January 2012. The first peak can be attributed to foreign currency attacks on the forint and to the new legislation introduced in Hungary which limited the role and independence of the central bank. The Hungarian currency depreciated then by 26% against the euro (see e.g. Valentinyi, 2012). By November 2011, the forint had depreciated by 56% against the Swiss franc. The country faced serious problems with foreign currency loans. As a consequence, in September 2011 the Hungarian government passed legislation that unilaterally changed the terms and conditions of all foreign currency loan contracts. The cost of

[2] According to Brown et al. (2009), in 2007 59% of total bank lending to households was in CHF, while the corresponding ratio for non-financial enterprises amounted to 16%. In Poland, 90% of CHF lending was taken out by households and was secured by mortgages. Loans to non-banking clients in Hungary in 2007 amounted to CHF 36.2 bn, while in Poland to CHF 30.9 bn. The amount of claims by Polish banks was thus almost as high as in Hungary, but while the share of CHF claims among the total foreign currency claims was higher in Poland (69%), their share of the total loans was lower (17%).

the transaction had to be borne entirely by banks. In mid-December 2011 the government and banks agreed to share the costs of any further arrangements. Figure 1 shows a constant and steep growth of the Hungarian spread up to the beginning of January 2011. Afterwards, the spread started to decrease.

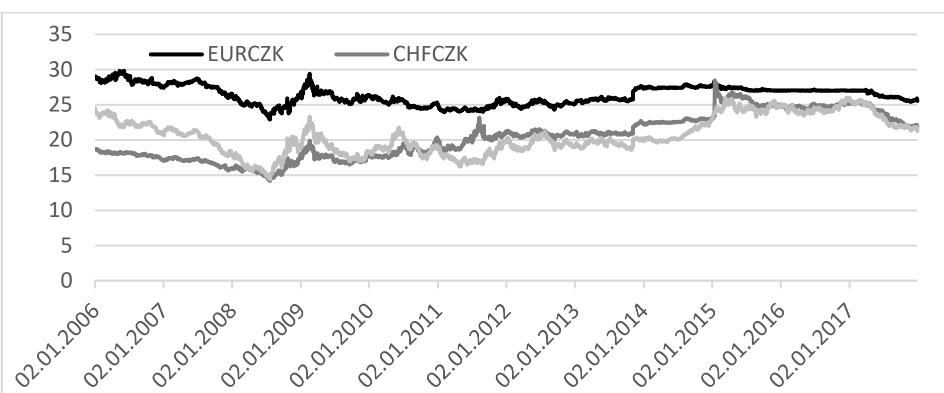
Figure 1. Dynamics of the spreads of Polish, Hungarian and Czech bond yields against the yields of German bonds



Note. PL_DE illustrates the difference between the Polish and German yields, HU_DE between the Hungarian and German yields, and CZK_DE between the Czech and German yields.
Source: authors' calculations.

The two peaks were not distinctly marked in the dynamics of the spreads of the Polish and Czech bonds. The peaks in the corresponding periods are observable, but not very high. The most spectacular increases were observed in November 2008 during the speculative attacks on local currencies.

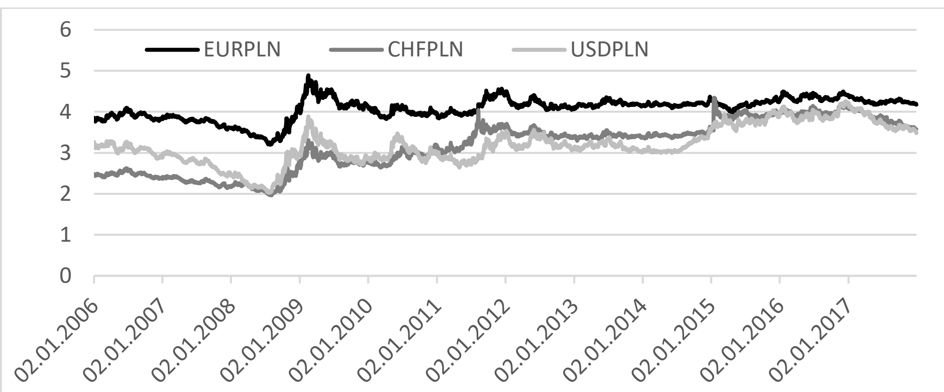
Figure 2. Dynamics of the exchange rate of the Czech koruna to Swiss franc, US dollar and the euro



Source: authors' calculations.

In Figures 3 to 5 the dynamics of the three currencies are presented: the Czech koruna, Hungarian forint and the Polish zloty, each expressed in three foreign currencies: the Swiss franc, the euro and the US dollar. At the beginning of the period, all the three countries had floating exchange rates, and additionally, Poland and the Czech Republic had free-floating rates. In November 2013, the Czech Republic decided to switch to 'other managed arrangement' (IMF, 2014). This IMF category includes currencies which are allowed to float independently but with discretionary management, as well as other practices which may apply to one currency only. The change took place on 7th November 2013 and was justified by problems with inflation targeting and expected continuous overshooting (IMF, 2014). The Czech National Bank announced that it would intervene in the foreign exchange market to weaken the koruna so that the exchange rate against the euro remained close to CZK 27 (but it would not intervene to strengthen the currency towards this level). This change is clearly visible in the dynamics of the Czech koruna as shown in Figure 2.

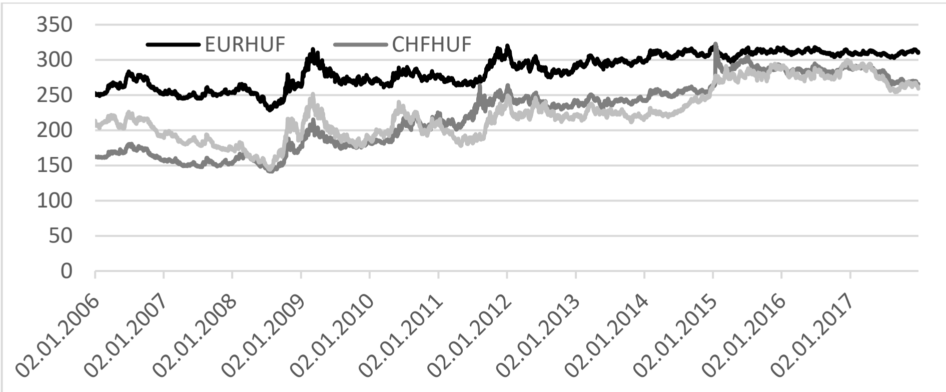
Figure 3. Dynamics of the exchange rate of the Polish zloty against the Swiss franc, US dollar and euro



Source: authors' calculations.

Figure 3 presents the dynamics of the exchange rates of the Polish zloty expressed in Swiss francs, US dollars and euros. A period of appreciation was observed until July 2008, followed by a sharp depreciation that continued until February 2009. Over the period 2010–2017, the price of the euro fluctuated around PLN 4.2, with the exception of late 2011–early 2012, when it grew to approximately 4.5 PLN. The joint dynamics of the CHF-PLN and the USD-PLN exchange rates is very interesting. Although at the beginning of 2006 the price of the Swiss franc was much lower than that of the US dollar, already in July 2008 the prices of both currencies almost equalled, and from then on, until the end of 2017, they 'intertwined', e.g. from January 2011 to January 2012 the US dollar was cheaper than the Swiss franc.

Figure 4. Dynamics of the exchange rate of the Hungarian forint to the Swiss franc, US dollar and the euro



Source: authors' calculations.

Figure 4 presents the dynamics of the prices of different currencies (Swiss francs, euros and US dollars) expressed in Hungarian forints, which was constantly depreciating over the analysed period. The end of February and beginning of March 2009 saw a sharp peak of the exchange rates of all the three analysed CEE currencies. This situation can be attributed to the deteriorating situation in Hungary and to the signing of the Supplemental Memorandum of Understanding between the European Union and Hungary on 11th March 2009.[3] Since all the three currencies and the bond spreads reacted at the same time, it may be presumed that investors might have expected the Hungarian problems to spread across CEE.

Table 1 contains descriptive statistics of the changes of the spreads and exchange rate series (the changes were modelled, as the levels of the spreads are non-stationary). The stationarity of the change series is confirmed by the KPSS test (Kwiatkowski et al., 1992) and the HML test (Harris et al., 2006). The exchange rates of the forint proved the most volatile in terms of standard deviation. The exchange rates of the Polish zloty and the Czech koruna demonstrate similar, yet significantly smaller standard deviations (than those of the forint). Additionally, the Hungarian bond spread was more volatile than the Czech and Polish ones, but the difference between their standard deviations was not as great as in the case of exchange rates. The ARCH effect applies to all the studied cases. All considered series are leptokurtic, but the different values of kurtosis of the considered time series is also worth

[3] The Memorandum of Understanding between the European Union and Hungary was signed in November 2008. In December 2008, Hungary received a disbursement of EUR 2 bn, while the second instalment was planned for March 2009. In February 2009, in the light of a deteriorating growth outlook for 2009, the Commission services together with the IMF staff revised the deficit target for 2009. This revision, together with a number of additional policy conditions, was laid down in the first Supplemental Memorandum of Understanding in March 2009.

noticing, for instance: 5.41 for changes of the USD-HUF exchange rate and 29.5 for the changes of the Polish bond spread.

Table 1. Descriptive statistics of the changes of bond spreads and foreign exchange rates

Variable	Obs No.	Mean	Std. Dev.	Skewness	Kurtosis	Min.	Max.
dCZ	3109	0.000251	0.065114	0.504075	85.593820	−1.2140	1.21600
dHU	3109	−0.000629	0.131711	−0.079470	15.631520	−1.2980	1.21600
dPL	3109	0.000350	0.072206	0.342191	60.561380	−1.2410	1.21600
dEUdCZK	3109	−0.001099	0.112960	0.065964	12.906930	−0.8740	1.17200
dCHFCZK	3109	0.001024	0.152034	11.880500	428.107900	−1.7340	5.14100
dUSDCZK	3109	−0.001034	0.160680	0.047913	4.241496	−1.2590	1.01100
dEUdHUF	3109	0.018729	1.723809	0.275266	4.353579	−11.8700	12.44000
dCHFHUF	3109	0.033204	2.073979	6.406754	189.420000	−20.3100	57.09000
dUSDHUF	3109	0.014577	2.102789	0.139020	3.059948	−11.2300	13.04100
dEUdPLN	3109	0.000102	0.024049	0.164908	6.841460	−0.1644	0.20150
dCHFPLN	3109	0.000340	0.028253	6.230465	183.971200	−0.2906	0.77090
dUSDPLN	3109	6.904E−05	0.029696	0.166432	4.370532	−0.1974	0.21304

Note. dCZ – changes in the Czech bonds spread, dPL – changes in the Polish bond spread, and dHU – changes in the Hungarian bonds spread.
Source: authors' calculations.

The dynamics of exchange rates reflect not only the domestic monetary policy, but also the monetary policy of the countries whose currencies are being referenced. Thus, in all the three charts a peak of exchange rates expressed in the Swiss franc can be observed in September 2011. Following the peak, the exchange rates relatively stabilised and their volatility decreased. This situation might have resulted from the Swiss National Bank's (SNB) decision announced on 6 September 2011, to no longer tolerate the EUR-CHF exchange rates below a minimum of CHF 1.20. The institution committed itself to enforcing this minimum rate and was prepared to buy foreign currencies in unlimited quantities. In addition, the SNB emphasised that, even the rate of CHF 1.20 per euro meant that the Swiss franc was too strong and was expected to continue to weaken over time (see: Chronicle of monetary events 1948-2016 on the official webpage of SNB and IMF, 2011). As a consequence, the previously free-floating exchange rate regime was reclassified to 'other managed arrangement' (IMF, 2012). From January 2012 to January 2013 the Swiss franc remained within a narrow 2% margin of the announced minimum exchange rate. Therefore, starting from January 2012, the de facto exchange rate arrangement was retrospectively reclassified from 'other managed arrangement' to 'stabilised arrangement'. For the sub-period of January – May 2013 it was again re-classified to 'managed arrangement' due to the departure of the currency from the stabilised band against the euro. However, starting from May 2013 the Swiss franc followed an appreciating trend within a 2% band against the euro. Therefore, the de facto exchange rate was reclassified to a 'crawl-like arrangement' from 'other managed arrangement' (IMF, 2013).

When considering the dynamics of the USD exchange rate, a peak which was not observed in the dynamics of the other exchange rates could be noticed in June 2010. This peak was most likely to have resulted from the EUR-USD exchange rate. On 8th June 2010, the price of the euro in the US dollars reached its minimum (1.1942). Interestingly, the EUR-CZK, EUR-HUF and EUR-PLN exchange rates did not react strongly to this. In conclusion, the EUR-USD dynamics affects the dynamics of the domestic CEE currencies against the US dollar stronger than against the euro.

4. The model

The goal of this study was to model interdependencies among bond spreads and exchange rates within a given country. Therefore, prior to selecting the appropriate model, the data was pre-tested against various hypotheses. One of them referred to constant or time-varying correlations. In the case of all series, the test of Engle & Sheppard (2001) and Tse (2000) strongly rejected the hypothesis of a constant conditional correlation. Therefore, only models with time-varying conditional correlation were taken into account.

The initial idea was to model the data using a DCC model (e.g. the one of Engle & Sheppard, 2001 or Tse & Tsui, 2002). One of the pre-requisites for this model is that all univariate conditional error distributions should be the same. However, even a short analysis of the descriptive statistics (Table 1) suggests that the empirical distributions of the modelled data vary across samples, as, for instance, in the case of kurtosis (from 5.41 for the variable dUSD-HUF, to 29.5 when considering the Polish bond spread). In consequence, when fitting univariate GARCH models with GED distributions to the data, different values of the degrees of freedom were recorded, depending on the analysed variable, which supports the thesis that conditional distributions also vary across samples. In such cases it is recommendable to apply the conditional copula model instead of using a DCC model. In dynamic copula models the structure of the dependence and dynamics of each univariate series is modelled separately. It allowed the application of GARCH-type models with GED innovations to model conditional variance of each univariate data, and then the t-copula to model the structure of dependence.

Copula models allow the use of measures of dependence other than the Pearson coefficient. When the time series distribution is not normal, using Pearson's correlation coefficient to identify the dependencies between random variables may yield misleading conclusions (Lindskog, 2000), since this coefficient is very sensitive to outliers. Moreover, a correlation equal to zero implies independence only if the variables are normally distributed. The heavier the tails, the larger the error of the estimator. Since the applied data are strongly leptokurtic (see Table 1), the option to use Pearson's correlation was rejected and the focus shifted onto the Kendall τ.

Another advantage of using copula is the possibility to investigate the dependencies between extreme values using tail-dependence measures. Assuming that the links between exchange rates and bond spreads tend to grow in response to internal or external shocks the studied economies experience, this approach seems to be the most relevant.

To sum up, in order to assess the strength of the aforementioned links among the analysed countries, the conditional copula model was applied. This model offers no restrictions on marginal distributions, and it allows for determining measures of dependencies other than the correlation coefficient does.

We further present a dynamic estimation of the rank correlation coefficient, the Kendall τ, as well as the tail dependence coefficient (λ). The latter measure is of particular importance to the analysis. It provides information on the probability of the transmission of extreme events from the risk countries to other countries. Schmidt (2002) explains that asymptotic dependencies should not be identified with a linear correlation coefficient. It is a well-known fact that in some cases the correlation between the considered series is strong, yet no dependence exists in the tails. It should be noted that a bivariate normal distribution is asymptotically tail-independent if its correlation coefficient ρ is less than 1.

Conditional copulas were introduced by Patton (2002, 2006). The author derived the properties of conditional joint distributions and the conditional copula from the properties of unconditional distributions and the copula. Conditional copulas have been applied and developed by numerous scientists, including Cifter & Ozun (2007), Doman, R. (2009, 2010), Hafner & Manner (2012), and Jondeau & Rockinger (2002, 2006).

Let the multivariate time series be denoted by $x_t = x_{1,t}, \ldots, x_{d,t}$. The general copula model can be described by the following formulas:

$$x_{i,t}|\Omega_{t-1} \sim F_{i,t}(\cdot\,|\Omega_{t-1}) \text{ for } i = 1, \ldots, d,$$
$$x_t|\Omega_{t-1} \sim F_t(\cdot\,|\Omega_{t-1}),$$
$$F_t(x_t|\Omega_{t-1}) = C_t(F_{1,t}(x_{1,t}|\Omega_{t-1}), \ldots, F_{d,t}(x_{d,t}|\Omega_{t-1})|\Omega_{t-1}),$$

where $\Omega_{t-1}$is the information set up to the moment $t-1$ inclusively. The existence and uniqueness of the C_t copula is guaranteed by the Sklar theorem for conditional copulas, introduced by Patton (2002). Let us consider the following model:

$$F_t(x_t, \alpha_1, \ldots, \alpha_d, \theta|\Omega_{t-1}) = C_t(F_{1,t}(x_{1,t}|\Omega_{t-1}, \alpha_1), \ldots, F_{d,t}(x_{d,t}|\Omega_{t-1}, \alpha_d)|\Omega_{t-1}, \theta),$$

where α_i is the parameter vector of the marginal conditional distribution $F_{i,t}$, and θ is the parameter vector of the conditional copula C_t. This model is estimated through the maximisation of the likelihood function in the following form:

$$L(\alpha_1, ..., \alpha_d, \theta) = \sum_{t=1}^{T} \ln c_t\big(x_{1,t}\big|\Omega_{t-1}, \alpha_1\big), ..., F_{d,t}\big(x_{d,t}\big|\Omega_{t-1}, \alpha_d\big)|\Omega_{t-1}, \theta) + \\ + \sum_{t=1}^{T}\sum_{j=1}^{d} \ln f_{j,t}(x_{j,t}|\Omega_{t-1}, \alpha_j),$$

where $f_{j,t}$ denotes the conditional marginal density function and c_t – the density function of the copula C_t.

The research herein is based on the DCC-t-copula model. The model was applied in two steps using the maximum likelihood method. In the first step, each univariate series $x_{i,t}$ is fitted; and the $u_t = u_{1,t}, ..., u_{d,t}$ is the multivariate time series, with each $u_{i,t}$ having been determined as the value of the cumulative distribution function for $\varepsilon_{i,t}$, to one of the univariate GARCH-type models with the t Student or GED innovation distribution.

$$\begin{gathered} x_{i,t} = \mu_{i,t} + y_{i,t}, \\ y_{i,t} = \sigma_{i,t}\varepsilon_{i,t}, \\ \varepsilon_{i,t} \sim \text{iid}(0,1), \\ u_{i,t} = F_i(\varepsilon_{i,t}), \end{gathered} \tag{1}$$

where $\varepsilon_{i,t}$ stands for the standardised residual series and F_i is the cumulative distribution function of the innovation distribution from the model fitted to $x_{i,t}$. The conditional mean $\mu_{i,t}$ was modelled as an ARMA-type model of the following form:

$$x_{i,t} = a_0 + \sum_{i=1}^{p} a_i x_{t-i} + \sum_{j=1}^{q} b_j y_{t-j.}$$

The authors apply standard GARCH models (Bollerslev, 1986), GJR-GARCH (Glosten et al., 1993), the IGARCH (Engle & Bollerslev, 1986) with t Student or GED innovation distribution with κ degrees of freedom to describe the dynamics of $\sigma_{i,t}^2$. In specific models, the conditional variance equations show the following specifications:

- GARCH(p,q) – $\sigma_t^2 = \omega + \sum_{i=1}^{p} \alpha_i\, y_{t-i}^2 + \sum_{j=1}^{q} \beta_j \sigma_{t-j}^2$, where y_t is the residual series;

- GJR-GARCH(p,q) $-\ \sigma_t^2 = \omega + \sum_{i=1}^{p} \alpha_i\, y_{t-i}^2 + \gamma_i S_{t-i}^{-} y_{t-i}^2 + \sum_{j=1}^{q} \beta_j \sigma_{t-j}^2$, where S_t^- is a dummy variable that takes the value of 1 when γ_i is negative and 0 when it is positive;
- IGARCH(1,1) – $\sigma_t^2 = \alpha y_{t-1}^2 + \beta\sigma_{t-1}^2$, where $\alpha + \beta = 1$.

In the second step, the conditional t copula is fitted to the u_t series, where the rank correlation matrix R_t is driven by the DCC model of Engle (2002).

$$C_{\upsilon,R_t}^{t}(u_t) = \int_{-\infty}^{t_\upsilon^{-1}(u_{1,t})} \cdots \int_{-\infty}^{t_\upsilon^{-1}(u_{d,t})} \frac{\Gamma\left(\frac{\upsilon+d}{2}\right)}{\Gamma\left(\frac{\upsilon}{2}\right)\sqrt{(\pi\upsilon)^d |R_t|}} \left(1 + \frac{x' R_t^{-1} x}{\upsilon}\right)^{-\frac{\upsilon+d}{2}} d^d x, \quad (2)$$

where $u_t = \left(u_{1,t}, \dots, u_{d,t}\right)'$, $x = (x_1, \dots, x_d)'$, $\Gamma(\cdot)$ is the gamma function,

$$R_t = \mathrm{diag}(Q_t)^{-1/2}\ Q_t \mathrm{diag}(Q_t)^{-1/2},$$

where the positive-definite matrix Q_t is described by the following formula:

$$Q_t = (1 - \textstyle\sum_{m=1}^{M} \alpha_m - \sum_{n=1}^{N} \beta_n)\bar{Q} + \sum_{m=1}^{M} \alpha_m\, \tilde{u}_{t-m} \tilde{u}_{t-m}' + \sum_{n=1}^{N} \beta_n\, Q_{t-n}, \quad (3)$$

where $\tilde{u}_t = \left[t_\upsilon^{-1}\left(u_{1,t}\right), \dots, t_\upsilon^{-1}\left(u_{d,t}\right)\right]$, where t_υ^{-1} is the inverse of the univariate, standardised t Student distribution, with υ denoting degrees of freedom. The log-likelihood function is provided by the following formula:

$$\begin{aligned} L_{\mathrm{St}}(R_t, \upsilon, \theta, \tilde{u}_t) = -T \ln \frac{\Gamma\left(\frac{d+\upsilon}{2}\right)}{\Gamma\left(\frac{\upsilon}{2}\right)} - pT \ln \frac{\Gamma\left(\frac{\upsilon+1}{2}\right)}{\Gamma\left(\frac{\upsilon}{2}\right)} - \\ - \frac{d+\upsilon}{2} \sum_{t=1}^{T} \ln\left(1 + \frac{\tilde{u}_t' R_t^{-1} \tilde{u}_t}{\upsilon}\right) \\ - \sum_{t=1}^{T} \ln|R_t(\theta)| + \frac{\upsilon+1}{2} \sum_{t=1}^{T} \sum_{i=1}^{p} \left(1 + \frac{\tilde{u}_{i,t}^2}{\upsilon}\right), \end{aligned} \quad (4)$$

where θ is the DCC parameter vector. More details about conditional copulas can be found in Doman & Doman (2013), Patton (2002) and Patton (2006).

The Kendall τ is applied as a measure of dependence. This is a measure of the ‘concordance’. Let (x_1, y_1), (x_2, y_2), (x_n, y_n) be a set of observation pairs generated from random variables X and Y. Observation pairs (x_i, y_i) and (x_j, y_j) are concordant if their ranks are consistent (i.e. if $x_i > x_j$ and $y_i > y_j$ or $x_i < x_j$ and $y_i < y_j$).

Similarly, observation pairs (x_i, y_i) and (x_j, y_j) are disconcordant if their ranges are not consistent (i.e. if $x_i < x_j$ and $y_i > y_j$ or $x_i > x_j$ and $y_i < y_j$). If $x_i = x_j$ or $y_i = y_j$, then observation pairs are neither concordant nor disconcordant. The Kendall τ coefficient is the difference between the probability of concordance of observation pairs (x_i, y_i) and (x_j, y_j), and the probability of their disconcordance. Thus,

$$\tau(X, Y) = P[(x_i - x_j)(y_i - y_j) > 0] - P[(x_i - x_j)(y_i - y_j) < 0]. \tag{5}$$

In the case of the conditional t copula, the Kendall τ coefficient is given by the formula:

$$\tau(X, Y) = \frac{2}{\pi}\arcsin(\rho),$$

where $\rho = R_{12,t}$ is the correlation coefficient between X and Y.

In this research, it is particularly important to check how the occurrence of extreme values of one series influences the probability of the occurrence of extreme values of the other series. The tail dependence coefficients λ^L and λ^U provide asymptotic measures of the dependence in the left and right tail, respectively. If F_1 and F_2 are cumulative distributions of vector (X, Y), then the tail dependence coefficients are given by the following formulas:

$$\lambda^L(X, Y) = \lim_{\alpha \to 0+} P(Y \le F_2^{-1}(\alpha) | X \le F_1^{-1}(\alpha)), \tag{6}$$

$$\lambda^U(X, Y) = \lim_{\alpha \to 1-} P(Y > F_2^{-1}(\alpha) | X > F_1^{-1}(\alpha)), \tag{7}$$

if the limits exist. In the case of the t copula, they are given by the formula:

$$\lambda(X, Y) = \lambda^U(X, Y) = \lambda^L(X, Y) = 2t_{\upsilon+1}\left(-\sqrt{\frac{(\upsilon+1)(1-\rho)}{1+\rho}}\right).$$

5. The results

The first step involves fitting a GARCH-type model to each considered series. Since a long sample is used, in some cases it was necessary to use a complicated GARCH-type model specification to explain the autocorrelation in squared residuals. The results of the estimations of the models are presented in Table 1. The details of the Box-Pierce test for standardised residuals and squared standardised residuals show that the models indeed explain the linear and non-linear dependencies. Following the estimation of the univariate models, standardised residuals were collected, and 4-dimensional t-copulas with a conditional covariance matrix explained by the DCC(1,1) model were fitted to the $u_{i,t}$ series. Taking into account the purpose of the study, the copulas were fitted to three exchange rates and the bond spread for each country separately. The estimation results are presented in Table 3.

Table 2. Results of the estimations of univariate GARCH models

dCHFCZK		dEURCZK		dUSDCZK		dCZ	
par.	estimate	par.	estimate	par.	estimate	par.	estimate
a_1	−0.055700***	a_1	−0.105700***	$ln(\xi)$	0.046570***	a_1	−0.0817***
v	4.241659	a_2	−0.051900***	v	7.055000***	a_2	−0.0321***
α_1	0.198700***	v	4.726103	ω	0.000100	a_3	0.0083***
β_1	0.801200***	α_1	0.098200***	α_1	0.049667***	v	1.0732***
		β_1	0.901800***	β_1	0.948515***	ω	0.4240
						α_1	0.0557***
						β_1	0.9325***
dCHFHUF		dEURHUF		dUSDHUF		dHU	
par.	estimate	par.	estimate	par.	estimate	par.	estimate
v	5.33840	v	1.3445	v	0.890000	v	1.0266
$ln(\xi)$	0.06856***	$ln(\xi)$	0.1075***	$b1$	−0.165100***	ω	0.0001
ω	0.05900	ω	0.0140	ω	0.027660	α_1	0.1320***
α_1	0.13470***	α_1	0.1126***	α_1	0.075700***	γ_1	−0.0923***
β_1	0.85520***	γ_1	−0.0881***	γ_1	−0.060000***	β_1	0.9085***
		β_1	0.9263***	β_1	0.947287***		
dCHFPLN		dEURPLN		dUSDPLN		dPL	
par.	estimate	par.	estimate	par.	estimate	par.	estimate
v	1.423200	v	1.422732	$ln(\xi)$	0.083279***	a_1	0.02700***
ω	0.077207	ω	0.05460 (–)	v	8.976840***	a_2	−0.04470***
α_1	0.177400***	α_1	0.125400***	ω	0.037000	v	8.67900
γ_1	−0.097600***	γ_1	−0.084700***	α_1	0.089780***	ω	1.15000
β_1	0.875100***	β_1	0.905500***	γ_1	−0.050640***	α_1	0.09206***
				β_1	0.910220***	β_1	0.89259***

Note. The following models were applied for the respective series:
dCHFCZK – GARCH(1,1) (t Student), dEURCZK – AR(2)-IGARCH(1,1) (t Student), dUSDCZK – GARCH(1,1) (skewed t Student), dCZ – AR(4)-GARCH(1,1) (GED) with restriction $a_3 = 0$.
dCHFHUF – GARCH(1,1) (skewed t Student), dEURHUF – GJR-GARCH(1,1) (skewed t Student), dUSDHUF – GJR-GARCH(1,1) (GED), dHU – GJR-GARCH(1,1) (GED).
dCHFPLN – GJR-GARCH(1,1) (GED), dEURPLN – GJR-GARCH(1,1) (GED), dUSDPLN – GJR-GARCH(1,1) (t Student), dPL – AR(2)-GARCH(1,1) (t Student).
Source: authors' calculations.

Table 3. Estimation results of 4-dimensional DCC-t-copulas with a conditional matrix R_t explained by the DCC(1,1) model – the Czech Republic, Hungary, Poland

Country	Parameter	Estimate	Std. error	*t*-stats	*p*-value
The Czech Republic	υ	10.8744	1.577	6.8935	<0.0001
	α_1	0.0299	0.002	12.3636	<0.0001
	β_1	0.9672	0.003	361.0533	<0.0001
Hungary	υ	10.0158	0.779	12.8511	<0.0001
	α_1	0.0251	0.003	9.4857	<0.0001
	β_1	0.9747	0.003	350.9480	<0.0001
Poland	υ	13.9115	1.234	11.2704	<0.0001
	α_1	0.0390	0.003	15.0528	<0.0001
	β_1	0.9600	0.003	345.9441	<0.0001

Source: authors' calculations.

The results of the estimations are presented in Figures 5 to 13. As the data suggest, the dynamics of the interrelations are partially similar for the studied countries, which is assumed to be related to selected international events of that period. Thus, the international events which were likely to have impacted the interrelationships (and the risk of volatility transmission) are described below, followed by an account of any domestic events which might have affected the dynamics as well.

First of all, when analysing the relationship of bond spreads with the euro, an 'echo' of the Greek crisis is visible in all the investigated cases (May 2010 saw the beginning of the Greek crisis, while the summer of 2011 a cut of the nominal value of Greek bonds). As regards the relationship of the exchange rates of domestic currencies to the Swiss franc, a global peak was observed in August 2011, followed by a period of a diminishing relationship between the spreads and exchange rates. This situation may have resulted from the previously-mentioned decision of the Swiss National Bank to control the EUR-CHF exchange rate.

A very sudden and steep decline of relationships in all the analysed cases was recorded in the second half of 2012, which may be explained by the situation on international markets. A fall in risk aversion was observed on the global financial market, mainly due to central banks' policies. The European Central Bank (ECB), for instance, took measures to improve the liquidity of the banking sector in the Eurozone and to reduce the tensions on the governmental bond markets of the selected Eurozone countries (Narodowy Bank Polski [NBP], 2012). The result of the election in Greece in June 2012 added to the fall in risk aversion. In July of the same year the ECB lowered interest rates (the main refinancing operations (MRO) rate to 0.75% and the interest rate on deposit facility to 0%) and in September it launched the outright monetary transactions (OMT) purchases programme for sovereign bonds in the Eurozone secondary markets. As a result, the risk perceived by the market participants decreased (as illustrated by the dynamics of sovereign CDS spreads), and the euro appreciated against the US dollar. The fall in global risk has evidently translated into a decline of the volatility transmission measured by parameter λ.

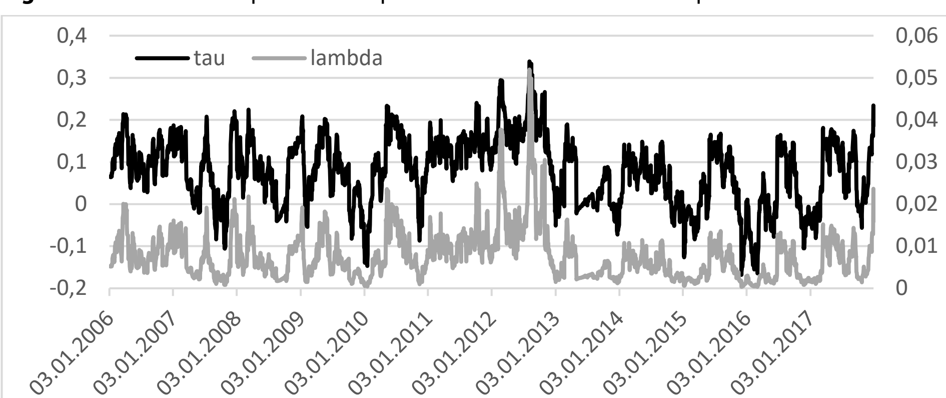
With respect to the dynamics of dependencies between the bond spreads and foreign exchange rates against the US dollar, the relationships in all the cases declined dramatically and reached negative levels at the end of 2013. The beginning of 2013 marked a sudden growth of government bond yields of mature markets due to the increase in global risk caused by the expected reduction of the scale of quantitative easing programmes in the US and the possible increase of interest rates by the US Federal Reserve (FED) at the end of 2013 (NBP, 2013). As a result, global markets' volatility increased, as well as the yields of government bonds in developed markets, and a sharp decrease in prices in global stock markets was observed. The second period, witnessing a significant rise in the risk aversion, occurred at the turn of August and September 2013, when the prices on global financial markets were negatively affected by the uncertainty regarding the time and scope of the shift in the FED's monetary policy and the possible military intervention in Syria. In September 2013, the markets were taken by surprise by the US macroeconomic data and the FED's decision to keep the level of the asset purchase programme unchanged, at the same time postponing the expected tightening of the monetary policy. This led to a decline in the yields on the Treasury debt securities on developed markets. At the same time, together with the ongoing political crisis in the US, the US dollar weakened against the euro. As the USD-PLN exchange rate was mainly affected by the EUR-USD exchange rate, the exchange rate of the zloty against the US dollar declined to its lowest level in over two years (NBP, 2013). The situation changed at the end of 2013, when the FED started tapering its asset purchase programme, and the Federal Open Market Committee (FOMC) reduced the quantitative easing programme following the committee's meeting in December 2013. At the same time, the ECB undertook measures to stimulate economic growth in the Eurozone and counteract the inflation rate persisting below the inflation target (NBP, 2014). These actions might have contributed to the drop of the τ coefficient, describing the relationships between the domestic bond spreads and exchange rates against the US dollar, below zero.

5.1. The Czech Republic

Figures 5–7 present the results of the estimations of the interdependencies between the Czech bond spread and the exchange rate of the Czech koruna, measured by the Kendall τ and λ (to make the picture clear, the value of the Spearman ρ was omitted, but its dynamics was similar to the dynamics of the τ coefficients, albeit its values were higher). As we can see, Kendall's τ was the highest in the case of the relationships between the bond spread and the EUR-CZK exchange rate. It fluctuated around 0.3, while the remaining exchange rates around 0.2.

The value of coefficient λ was very low throughout the whole period, within which three instances of growth occurred as a consequence of international events. The first one appeared in late 2008 and early 2009, and resulted from the transmission of the crisis to Europe. The second peak was observed in spring 2010. It was most probably caused by the beginning of the Greek problems. The third one took place in the second half of 2012 and its source is likely to have been the early elections in Greece. However, the value of λ decreased shortly afterwards and as early as in November 2012 it returned to nearly 0, which might have been the effect of the policy adopted by the Czech National Bank (CNB). In late 2012, the interest rates in the Czech Republic reached 0% and the CNB announced that it would intervene in the foreign exchange market to weaken the koruna if necessary (such verbal interventions took place in 2013). Moreover, from November 2012, the CNB suspended the sales of foreign exchange reserve revenues (which were aimed at preventing a continuous rise in the reserve level), and began to publish monthly foreign exchange transaction data on its website, with a two-month lag. Moreover, May 2013 saw a sudden drop in the Kendall τ. This possibly was the consequence of the aforementioned policies and of the FED's decision to give green light to the gradual withdrawal of monetary stimulus (which influenced the EUR-USD exchange rate – see below).

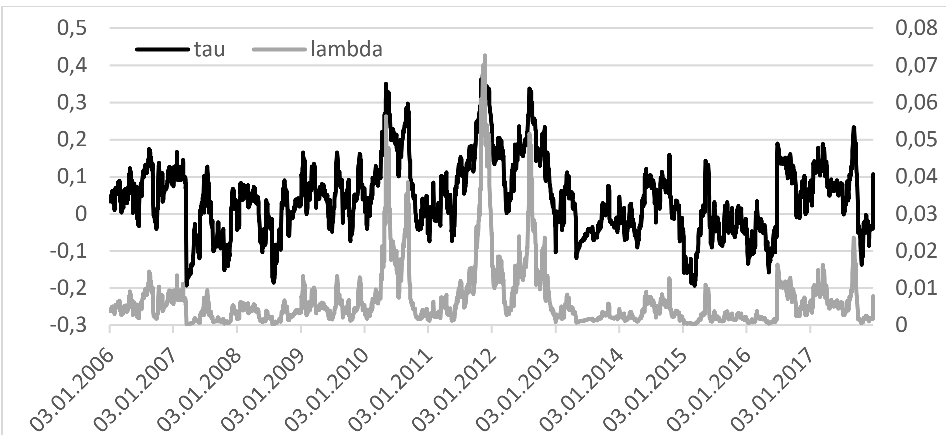
Figure 5. Estimates of copula-GARCH parameters τ and λ: the Czech spread and CHF-CZK



Note. Black line – τ coefficient, grey line – λ.
Source: authors' calculations.

Figure 6 presents the dynamics of the relationships between the Czech bonds spread and the EUR-CZK exchange rate. As mentioned before, it reflects to a great extent the dynamically changing situation on the pan-European market (for instance, the peak in May 2010 can be attributed to the Greek crisis). Starting from 2013, a change of relationships between the two magnitudes may be observed. It was the time when the CNB decided to stabilise the Czech koruna around the euro. This change is also visible in the remaining exchange rates, but in their case it is not as drastic.

Figure 6. Estimates of copula-GARCH parameters τ and λ: the Czech spread and EUR-CZK

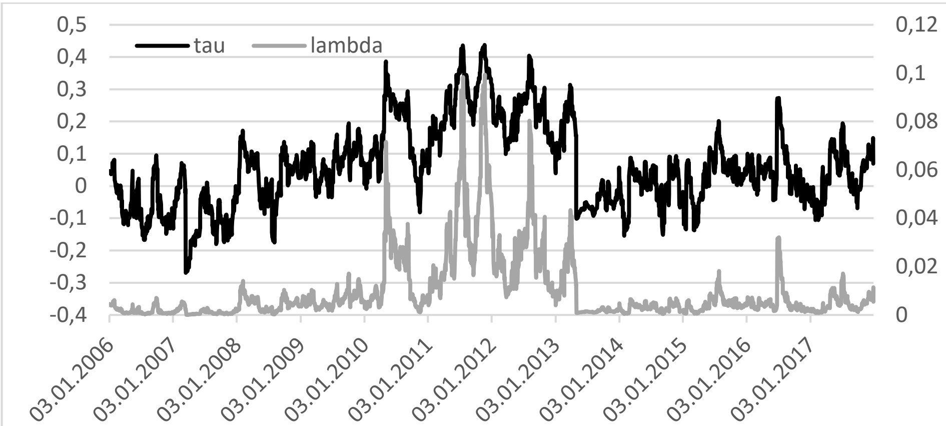


Note. Black line – τ coefficient, grey line – λ.
Source: authors' calculations.

A very interesting pattern emerges relating to the interactions between the Czech bond spread and the USD-CZK exchange rate (Figure 7). A spike of λ occurred also in November 2011, which might have resulted from the worsening economic situation in the US. In October 2011, the US President announced a USD 447 bn plan to stimulate the economy. The plan did not have the expected effect on market uncertainty, so the FED representatives announced further monetary policy easing through the purchase of securities (NBP, 2011).

The Kendall τ took negative values between November 2013 and January 2014. As explained before, the situation was most likely the result of the FED's announcement of a possible gradual withdrawal of monetary stimulus and the simultaneous introduction of a quantitative easing policy by the ECB. The fall in the relationship was most likely reinforced by the decision of the Czech National Bank, which, starting November 2013, began using the exchange rate as an additional monetary policy instrument.

Figure 7. Estimates of copula-GARCH parameters τ and λ: Czech spread and USD-CZK

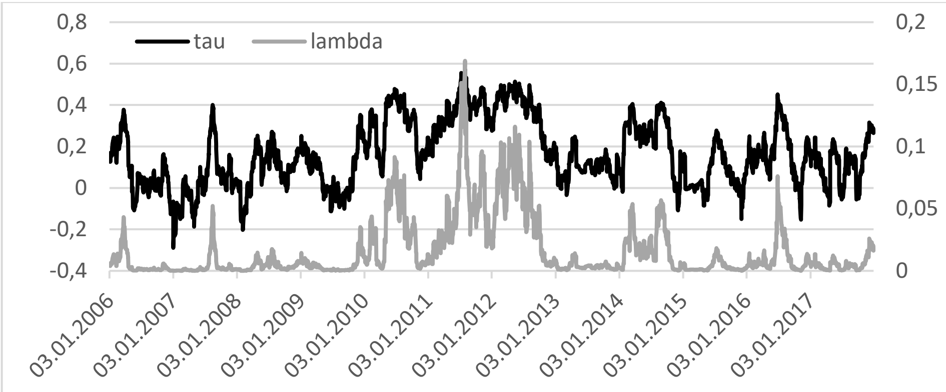


Note. Black line – τ coefficient, grey line – λ.
Source: authors' calculations.

5.2. Poland

Among all the analysed currencies, the strongest relationships were obtained for bond spreads and the CHF-PLN exchange rate. The years 2011–2012 saw a growth in these relationships, with the highest peak coinciding with the beginning of the stabilisation of the Swiss franc (Figure 8). Over this period, the correlation was quite stable and the probability of extreme events transmission diminished. This change was probably caused by foreign exchange (FX) interventions. On 21st April 2011, the Polish Ministry of Finance announced it would regularly sell some foreign currency received by Poland in the framework of EU funds directly on the domestic market. On 6th July 2011, the president of the National Bank of Poland (NBP) announced the possibility of its intervention in the FX market in order to prevent the excess volatility of the price of the Polish zloty. Subsequent interventions took place in 2011 in September (a common intervention of the NBP and the State Development Bank of Poland – BGK), October, November and December 2011 (Blox, n.d.). The last one, which was performed on 29th December 2011, was also a joint intervention of the NBP and the BGK, aimed at preventing the further depreciation of the Polish zloty and recalculating the value of Poland's foreign debt. During the period of interventions, the risk of volatility spillover from the foreign exchange market onto the domestic market was gradually decreasing.

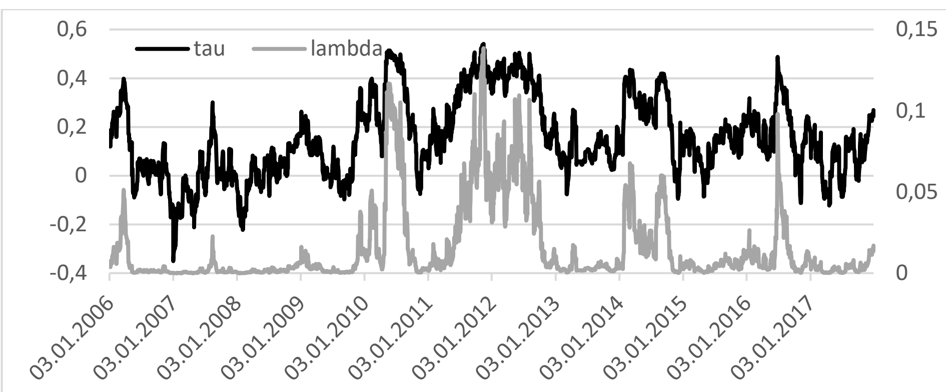
Figure 8. Estimates of copula-GARCH parameters τ and λ: Polish spread and CHF-PLN



Note. Black line – τ coefficient, grey line – λ.
Source: authors' calculations.

The relationships with the euro were rather stable over the whole period, although some peaks did occur during the most hectic moments of the European debt crisis. The highest one was observed in August 2012 (Figure 9).

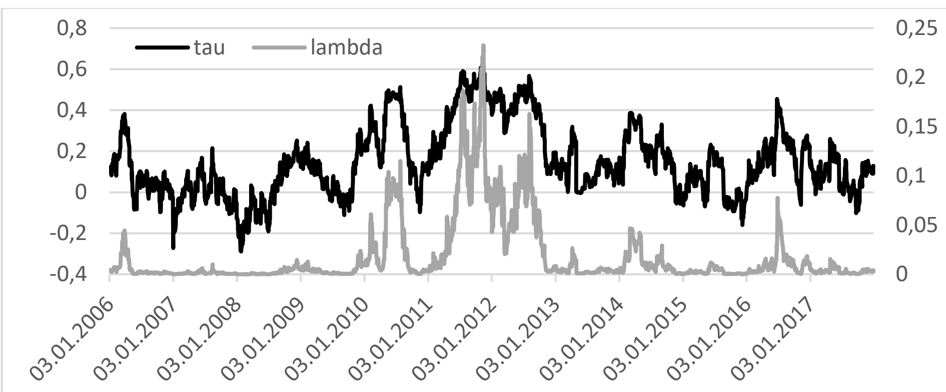
Figure 9. Estimates of the copula-GARCH parameters τ and λ: Polish spread and EUR-PLN



Note. Black line – τ coefficient, grey line – λ.
Source: authors' calculations.

Interestingly, the USD-PLN relationships (see Figure 10) were quite stable until the end of the first half of 2012. After that period, the τ (correlation) and λ (probability of extreme events transmission from the FX market) spiked, followed by a deep and long-lasting fall in these two kinds of relationships. It should be noted that apart from the above-mentioned international events that could have caused this fall, some legislative changes were also introduced concerning the method of determining the ratio of public debt to GDP. This decision contributed to the decrease of the exchange rates' volatility (see: Bank Gospodarstwa Krajowego [BGK], 2013), and additionally were likely to have influenced the risk of the volatility transmission between foreign exchange and bond markets. Moreover, in 2012, Poland introduced a regulatory framework for foreign currency lending, requiring banks to offer mortgages only in the currency of the mortgagor's income, and additionally to impose stricter creditworthiness standards to foreign exchange credit exposure (IMF, 2014). This is likely to have contributed to the reduction of households' exposure to the currency risk and to the decline of relationships between the bond spreads and exchange rates. Three interventions performed in 2012 (one in February and two in May) did not seem to affect the relationships between the exchange rates and bond spreads.

Figure 10. Estimates of copula-GARCH parameters τ and λ: Polish spread and USD-PLN



Note. Black line – τ coefficient, grey line – λ.
Source: authors' calculations.

5.3. Hungary

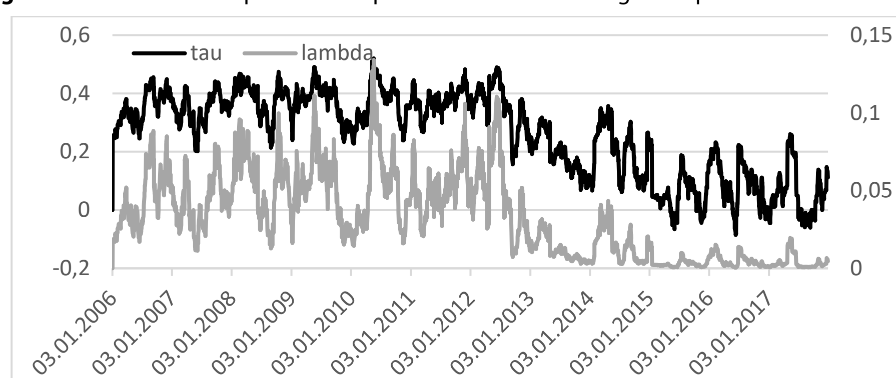
This subchapter focuses on the estimates of interrelations between the Hungarian bond spread and the exchange rates (Figures 11–13). The difference between their dynamics in Hungary and the dynamics of the relationships on the Polish and Czech markets is striking. The Hungarian market seems to have been the most affected by its domestic situation, presumably resulting from the Hungarian government's policy.

In Poland, prior to the crisis, hundreds of thousands of households took out mortgage loans denominated in foreign currency, most often in the Swiss franc or the euro. The advantage of foreign currency loans over those in the domestic currency was that in that period, the former offered substantially lower instalments than mortgages in the Polish zloty (Gereben et al., 2011). However, the financial crisis caused the real estate values to plunge and domestic currencies to weaken. This made foreign – currency-denominated loans more difficult to pay off. In May 2011, the Hungarian government adopted repayment schemes, allowing foreign currency mortgage loans to be repaid in lump sum at artificially weak exchange rates before maturity. Some customers benefited from the scheme and in result the total amount of outstanding foreign currency-denominated mortgages dropped by over 19%. However, the cost which the Hungarian banks had to face was high – it was assessed at approximately USD 1.2 bn (Agabekian, 2013).

On 15th December 2011, the Orban administration and the Hungarian Bank Association concluded an agreement that enabled mortgagors to repay the entire sum of their Swiss franc- and euro-denominated loans at the above-mentioned fixed rates by March 1, 2012. Banks, in turn, were permitted to deduct one-third of the exchange rate losses resulting from this arrangement from their payment of the government-imposed financial-sector tax in 2012 (Lambert, n.d.).

On February 27, 2012 the Hungarian central bank (Magyar Nemzeti Bank, further the MNB) terminated the foreign currency sales tender programme. Prior to that, the MNB operated a programme of tenders of foreign exchange sales to provide banks with foreign currency to close their open positions arising from the early repayment of foreign currency-denominated mortgages (IMF, 2013).

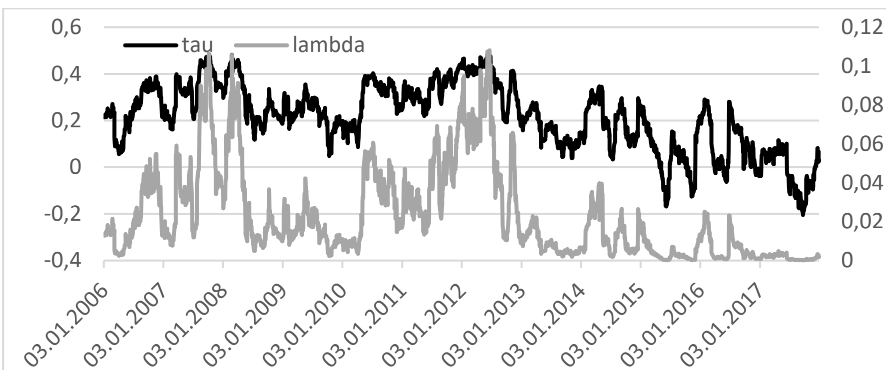
Figure 11. Estimates of copula-GARCH parameters τ and λ: Hungarian spread and CHF-HUF



Note. Black line – τ coefficient, grey line – λ.
Source: authors' calculations.

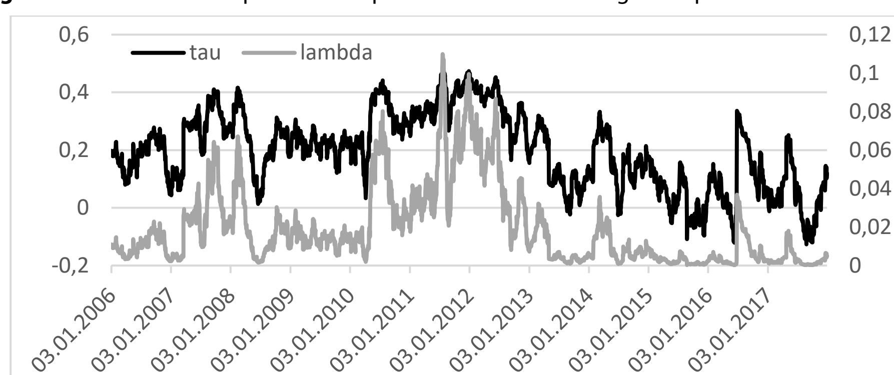
Figures 11–13 indicate the moment in which in all the studied cases the relationships between the bond spreads and the exchange rates dropped sharply. This fall in the case of the CHF-HUF exchange rate is especially striking. The drop occurred in slightly different moments for each currency. Nevertheless, the fact that it was most significant in the case of Hungary seems to have evidently resulted from the Hungarian government's policy. Thus, out of the three analysed economies, Hungary is the one where domestic factors influenced the relationships between the bond spread and exchange rates to the largest extent. This conclusion should not be surprising, considering the fact that according to numerous authors (e.g. Baldacci & Kumar, 2010; Jaramillo & Weber, 2013), domestic factors play a prominent role when an economy is in crisis, influencing other variables. In this context we can say that Hungary – unlike Poland and the Czech Republic – underwent a crisis of its own.

Figure 12. Estimates of copula-GARCH parameters τ and λ: Hungarian spread and EUR-HUF



Note. Black line – τ coefficient, grey line – λ.
Source: authors' calculations.

Figure 13. Estimates of copula-GARCH parameters τ and λ: Hungarian spread and USDHUF



Note. Black line – τ coefficient, grey line – λ.
Source: authors' calculations.

6. Conclusions

The aim of the article was to verify the impact of foreign currencies on the fundamentals of three CEE economies: the Czech Republic, Hungary and Poland. The condition of the fundamentals was measured through the dynamics of the spreads of 10-years' sovereign bonds against a 10-years' yield of German bonds. Two of the countries – Poland and the Czech Republic – performed relatively well during the financial crisis, while Hungary struggled with internal problems.

We estimated a 4-dimensional copula-GARCH model, on the basis of which we obtained a time-varying estimate of the Kendall τ and tail dependence coefficient λ, illustrating the probability of volatility spillovers between the sovereign market and the FX one. Three most important reference currencies were considered: the US dollar, the euro and the Swiss franc. The impact of the first two on the investigated economies is indisputable (see e.g. Doman M., 2009), while the importance of the Swiss franc stems from a huge exposure of households to mortgages denominated in this currency. This exposure was especially great in Poland and Hungary.

The question formulated in the title of this paper concerns the degree to which the currencies affect the fundamentals of the CEE economies. The analysis of the values of the Kendall τ – the measure of concordance, and the values of coefficient λ – the probability of extreme events transmission, shows that at the beginning of the period covered by the study the highest values were obtained for interrelations with the Swiss franc in Poland and Hungary. These results can be explained by the fact that Polish and Hungarian households had been heavily indebted in the Swiss franc, while this phenomenon was not that strong in the Czech Republic. Still, even in the case of the Czech Republic, the highest values of τ were also obtained in relation to the Swiss franc, which might be the result of various interactions among the CEE currencies (see for instance Orlowski, 2016) or the important role of the Swiss franc in the whole of Europe. Moreover, the Swiss central bank actually stabilised the franc around the euro, so the high dependence between the Czech koruna and the Swiss franc could also be the consequence of the strong relationship between the franc and the euro.

On the basis of the analysis presented in this paper, it can be additionally concluded that in the case of the Czech Republic and Poland, the dependencies between the FX and the sovereign market grew in the periods of international turmoil. Domestic events, such as currency interventions, seemed to have a weak influence on the interrelations and the probability of volatility spillover between the markets. The change of the floating regime of the koruna had merely a short-term effect on the relationships between the exchange rates and the fundamentals – after a short period of the relationships' weakening, their strengthening occurred again.

Hungary was in a different situation, as the domestic credit policy of the Hungarian government seemed to have a significant influence on the dependencies between the sovereign market and FX market. The probability of spillovers diminished visibly following the implementation of the obligatory conversion of foreign currency-denominated loans to forint-denominated ones. From this point of view, the controversial policy of the Hungarian government should be considered effective. The obtained results also support a thesis, quite frequently cited, that when the market is distressed, domestic factors influence the sovereign market to the largest extent, more so than international circumstances.

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The effect of financial, macroeconomic and sentimental factors on stock market volatility

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Abstract. The aim of this paper is to find economic factors that could be helpful in explaining the market's shifts between periods of prosperity and crisis. The study took into account the main stock indices from developed markets of the USA, Germany and Great Britain, and from two emerging markets, i.e. Poland and Turkey. The analysis confirms the existence of two different states of volatility in these markets, namely the state with a positive returns' mean and low volatility, and the state with a negative or insignificant mean and high volatility. The Markov-switching model with a dynamic probability matrix was applied in the study. The subject of the analysis was the impact of domestic and global factors, such as VIX and TED spread, oil prices, sentiment indices (ZEW), and macroeconomic indices (unemployment, long-term interest rate, CPI), on the probability of switching between the states. The authors concluded that in all the examined countries, changes in long-term interest rates have an influence on market returns. However, the direction of this impact is different for developed and emerging markets. As regards developed markets, high prices of oil, 10-year bonds, and the ZEW index can suggest a high probability of the countries remaining in the first state, whereas an increase in the VIX index and the TED spread significantly reduces the probability of staying in this state. The other studied factors proved to be rather local in nature.

Keywords: regime shift, equity volatility, macroeconomic factors, sentimental factors, financial markets, TVPMS model

JEL: C52, G11, G15, G32

1. Introduction

Understanding the stock market's mood is crucial for investors and policymakers. The diversification strategies created to reduce investment risks are closely tied to a given stock market's nature. After the global financial crisis, theorists and practitioners began to take notice of the volatility of international stock markets. A 'volatility shift' means that volatility transitions from a low to a high level, usually corresponding to crisis periods (Aloy et al., 2014). From the practical point of view, it is worth knowing what impact various indicators have on the markets.

A lot of research has been conducted on the factors which interact with financial markets, such as political events, the economic situation and investors' expectations (Huang et al., 2005). As the stock market is a part of the economy and stock prices are often determined on a cash-flow basis, fundamental macroeconomic indicators can

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influence stock market prices and they tend to be included in the portfolio investment decision-making process (Chen, 2009; Haq & Larson, 2016; Pilinkus, 2010). Rapach et al. (2005) presented evidence that stock returns can be predicted on the basis of macroeconomic variables. Chen (2009) investigated whether macroeconomic variables can predict a recession in the stock market. The author evaluated series such as interest rate spreads, inflation rates, money stocks, aggregated output, unemployment rates, federal funds rates, federal government debt, and nominal exchange rates, and concluded that bear markets can be easily predicted on the basis of macroeconomic variables. The relationships between stock prices and chosen economic variables were discussed by a variety of scientists, including Mahmood & Dinniah (2009). Chang (2009) and Humpe & Macmillan (2007) approach this issue using the Markov-switching mechanism. Nasseh & Strauss (2000) proved the existence of a long-run relationship between stock prices and the macroeconomic activity in six major European countries. They concluded that stock markets were driven by economic fundamentals and a number of interrelated factors, such as production, business expectations, interest rates and the CPI. The existence of long-run equilibrium relationships among stock prices, industrial production, real exchange rates, interest rates and inflation in the United States was investigated by Kim (2003). Celebi & Hönig (2019) demonstrated that the impact of external factors on stock prices in Germany is stronger in times of crisis than in the pre- or post-crisis periods. Research on the developing Vietnamese stock market (Nasir et al., 2020) also showed a link between macroeconomic variables and stock prices. Real economic growth and easy crediting had a positive impact on the stock market, whereas inflation caused long-term negative effects. The impact of sentiment indicators, based on the expectations of analysts and investors, was discussed for example by Kvietkauskienė & Plakys (2017) and many others (see Algaba et al., 2020). The German Zentrum für Europäische Wirtschaftsforschung (ZEW) Economic Sentiment Index proved to have predictive power for technology-oriented stock companies in Germany (Homolka & Pavelková, 2018). Also, the tone of the economic news in the media (García, 2013; Lischka, 2015) or the overall mood of Facebook users (Siganos et al., 2014) were found to be good indicators for stock markets about the general economic situation, especially during recession.

This paper examines the effect of various indicators on stock market returns. A selection of developed markets was analysed: the USA, Germany and Great Britain, as well as two emerging markets, i.e. Poland[1] and Turkey. Research was

[1] A leading global index provider, FTSE Russell promoted Poland from the status of an Emerging Market to the status of a Developed Market on 24 September 2018.

performed on monthly returns of the main stock indices of the considered countries (SPX, DAX, WIG, XU, FTM) and monthly data of the exogenous variables from the period of January 2001 to January 2019. The research revealed the existence of volatility shifts from a low to a high level, usually corresponding to prosperity and crisis periods, respectively. Subsequently, an attempt was made to determine which of the indicators – global or domestic – could be of use in explaining or predicting volatility shifts.

The applied methodology is based on the Markov-switching model (Hamilton, 1990). The regime switching models with a Markov switching mechanism for modelling financial time series were discussed by Chollete et al. (2009), Jondeau & Rockinger (2006), Rodriguez (2007) and others. Switching models were also analysed by Czapkiewicz (2018), Doman (2011) and Doman & Doman (2014). In order to verify the impact of financial, macroeconomic and sentimental factors on the stock market volatility, we adopted the Copula-GARCH model with Markov switching with a time-varying transition probability matrix. A time-varying transition probability Markov-switching (TVPMS) framework was originally proposed by Filardo (Filardo, 1994) and further developed by Kim et al. (2008). This approach has already been applied by researchers to verify the influence of selected indicators on the behaviour of some financial time series. For example, Boudt et al. (2012) used the TVPMS mechanism to study the impact of the VIX or TED spread on the dependencies between weekly returns on the US headquartered bank holding companies. Aloy et al. (2014) showed volatility shifts between tranquil and crisis periods in the East Asian equity markets. Dufrénot et al. (2014) applied this method to study the impact of the anticipated macroeconomic fundamentals on the Eurozone sovereign spreads, while Toparlı et al. (2019) used it to study the impact of oil prices on the stock returns in Turkey. The TVPMS model is discussed in detail also in a monograph by Czapkiewicz (2018).

This article considers only three developed markets and only two developing markets; nevertheless, some observations could be made. The relationship between financial and macroeconomic factors and market volatility is the subject of numerous articles. However, the specific contribution of this paper is the verification of the thesis that both global factors such as the VIX, TED spread, oil prices, the ZEW index, and chosen macroeconomic variables, including the consumer price index, long-term interest rates and unemployment rates, may be crucial for the state of the volatility of markets (emerging or developed). Particular attention is devoted to the impact of the ZEW sentiment factor on the markets. To the authors' best knowledge, this factor has not been widely studied yet. The paper attempts to investigate what variables may affect regime shifts and seeks an answer to the questions whether sentimental factors matter and weather it is only macroeconomic and financial data that impact regime shifts.

The TVPMS model was defined by Filardo (1994), but the financial literature fails to provide any further information on its usage. There are no ready-to-use procedure libraries, thus using this model in practice requires the implementation of one's own algorithms. In addition, the applied methodology makes it possible to study the impact of these factors on market volatility in each of the states (prosperity or crisis) considered separately. This model applied in practice shows whether the examined factors are of greater importance in the period of prosperity or in the period of crisis. To the best of the authors' knowledge, such a study has not been conducted for the Polish nor Turkish market.

This paper further contains the following parts: Section 2 describes the model's specifications and its estimation procedure, Section 3 presents the results of the empirical study, while the conclusions are provided in the last, fourth, section of the paper.

2. Econometric framework

2.1. The TVPMS model

Let us consider a process $(S_t, R_t)_{t\in\mathbb{N}}$, where $(R_t)_{t\in\mathbb{N}}$ is a returns time series, $(S_t)_{t\in\mathbb{N}}$ is a hidden Markov process with transition matrix P_t and with two states, i.e. $s_t \epsilon \{1,2\}$. The matrix of the transition probabilities is defined as follows:

$$P_t = \begin{bmatrix} p_t^{11} = \dfrac{exp(x_{t-1}^T \beta_1)}{1 + exp(x_{t-1}^T \beta_1)} & p_t^{12} = 1 - \dfrac{exp(x_{t-1}^T \beta_1)}{1 + exp(x_{t-1}^T \beta_1)} \\ p_t^{21} = 1 - \dfrac{exp(x_{t-1}^T \beta_2)}{1 + exp(x_{t-1}^T \beta_2)} & p_t^{22} = \dfrac{exp(x_{t-1}^T \beta_2)}{1 + exp(x_{t-1}^T \beta_2)} \end{bmatrix},$$

where $p_t^{ij} = P(S_t = j | S_{t-1} = i)$ is a time-varying transition probability (model TVPMS), evolving as a logistic function of $x_{t-1}^T \beta_i$, and matrix x_{t-1}^T (or x_t^T) contains variables that affect transition probabilities. We assume that:

$$R_t = \mu_{s_t} + b_{s_t} R_{t-1} + \varepsilon_t \quad \text{and} \quad \varepsilon_i \sim N(0, \sigma_{s_t}). \tag{1}$$

If there is no statistically meaningful impact of the exogenous variables x_t^T on returns, then the TVPMS model converges to the Markov-switching model with fixed transition probabilities (MS model). In this case, the LM test statistic could be applied to test the null hypothesis, which assumes that the considered models are equivalent against the alternative hypothesis which assumes that the dynamic model is better (Vuong, 1989). This statistic takes the following form:

$$LM = 2\big(l(\theta) - l_F(\theta_1)\big), \tag{2}$$

where $l(\theta)$ and $l_F(\theta_1)$ are the log-likelihood functions of the models with time-varying and fixed transition probabilities, respectively. Despite the fact that in this test the classical regularity conditions are not fulfilled, the asymptotic distribution of LM is the central chi-square distribution (Czapkiewicz, 2018; Vuong, 1989; White & Domiwitz, 1984).

2.2. The procedure of estimating the Markov-switching model parameters

The estimation of the unknown model parameters is performed on the basis of the Hamilton filters (Hamilton, 1990). Let θ denote the collected parameters $(\mu_1, \mu_2, b_1, b_2, \sigma_1, \sigma_2)$ from (1) and parameters of the transition probabilities: $\beta_1 = (\beta_{01}, \beta_{11})$, $\beta_2 = (\beta_{02}, \beta_{12})$. The log-likelihood function takes the following form:

$$l(\theta) = \textstyle\sum_{t=1}^{T} \log\Big(\sum_{j=1}^{2} f_{s_t}(r_t|\Omega_{t-1};\theta)\, P(S_t = s_t|\Omega_{t-1};\theta)\Big), \tag{3}$$

where $f_{s_t}(\cdot)$ is the distribution of the random variable R_t conditional on the information set Ω_{t-1} in the $\mathrm{s_t}$ state, $s_t \epsilon \{1, 2\}$, and r_t is the observable of return at time t. Let η_t denote a vector of two densities governed by the Markov process at date t:

$$\eta_t = [f_1(r_t|\Omega_{t-1};\theta), f_2(r_t|\Omega_{t-1};\theta)]^T, \tag{4}$$

and let $\hat{\xi}_{t|t-1}$ denote the collected conditional probabilities $P(S_t = j|\Omega_{t-1};\theta)$:

$$\hat{\xi}_{t|t-1} = [P(S_t = 1|\Omega_{t-1};\theta), P(S_t = 2|\Omega_{t-1};\theta)]^T. \tag{5}$$

The optimal inference and forecast for each t in the sample can be found by iteration, using the following pair of equations:

$$\hat{\xi}_{t|t} = \frac{\hat{\xi}_{t|t-1} \odot \eta_t}{1^T(\hat{\xi}_{t|t-1} \odot \eta_t)},$$

$$\hat{\xi}_{t|t+1} = P_t \hat{\xi}_{t|t}.$$

Hence, the symbol $\odot$ denotes an element-by-element multiplication. The log-likelihood function takes now the form presented below:

$$l(\theta) = \sum_{t=1}^{T} \log\left(1^T\left(\hat{\xi}_{t|t-1} \odot \eta_t\right)\right).$$

The parameter estimates of the standard Markov-switching model (MS) are performed in the same way, but instead of the time-varying transition matrix, we take a matrix with fixed transition probabilities. To evaluate the model's goodness-of-fit, we use the diagnostic test proposed by Diebold et al. (1998). Let F be the conditional cumulative distribution functions of R_t. If a distribution is correctly specified, $u_t = F(r_t|R_{t-1}; \theta_1)$ should be i.i.d. uniform [0, 1] distributed.

3. Empirical study

3.1. Data

The research concerns five countries: the USA, Germany, the United Kingdom (as developed markets), and Poland and Turkey (as East European, emerging markets). The US stock exchange is the one with the largest capitalisation in the world, Germany has the strongest economy in Europe, whereas the London Stock Exchange is the largest stock market in Europe. The Warsaw Stock Exchange represents the stock markets of Eastern Europe and has long been part of the group of developing markets. The Turkish stock exchange represents behaviour typical for developing markets. We consider monthly returns of main stock indices, computed as $r_{it} = \ln\frac{P_{i,t}}{P_{i,t-1}}$, where $P_{i,t}$ is the closing price of i-th index in t, and $P_{i,t-1}$ is the closing price in the previous month (i.e. the closing price of the last session in a given month). The following indices were considered: WIG (Poland), DAX (Germany), FTM (UK), XU (Turkey), and SPX (S&P500, the USA). Monthly data values of the indices came from the period of January 2001 to January 2019. We decided to use monthly data, as the selected exogenous variables are noted on a monthly basis. Table 1 presents the basic descriptive statistics for all indices' returns: mean, median and standard deviation. The mean of the returns ranges from 0.2 to 1.1 percent. The largest mean is for Turkey, whereas the lowest – for Germany. The mean of the Polish index's returns is 0.06 percent which situates it in second place (after Turkey). In all the cases, the median is higher than the mean, which is also reflected in negative skewness. Kurtosis is high, compared to the value of 3 for normal distribution. All things considered, the return rates of these indices come from distributions typical for financial data: with most of the values very small, but positives, and some rare, but severe losses. The standard deviation is the largest for Turkey. All the series were also tested with the Dickey-Fuller test, which confirmed their stationarity (p-value < 0.01).

Table 1. Descriptive statistics of indices' returns

Country (index)	Mean	Median	Standard deviation	Skewness	Kurtosis
The US (SPX)	0.003	0.009	0.042	−0.457	16.432
Germany (DAX)	0.002	0.008	0.060	−0.166	9.788
The UK (FTM)	0.005	0.008	0.048	−0.585	10.734
Poland (WIG)	0.006	0.007	0.061	−0.692	10.337
Turkey (XU)	0.011	0.015	0.095	−0.227	8.064

Note. The table reports descriptive statistics of monthly indices' returns from January 2001 to January 2019. All means are insignificant.

Source: authors' calculation.

We consider financial, sentiment and macroeconomic factors to investigate their co-movement with stock indices. As financial factors, we take into account the VIX and TED spread indices. The VIX is a volatility index, often referred to as 'fear index'. It was first introduced by the Chicago Board Options Exchange (CBOE) in 1993 to measure expectations of the volatility of the S&P500 index's options. The price of an option represents the expectations of a 30-day forward-looking volatility. The TED spread is computed as the difference between the three-month U.S. government Treasury bill and the three-month LIBOR and is considered to be an indicator of credit risk. High values mean that investors are prone to allocating money into secure government treasury bills, rather than lending them to banks.

As a sentiment factor, we consider the ZEW Index. The German Zentrum für Europäische Wirtschaftsforschung (ZEW) Economic Sentiment Index is based on a survey of German institutional investors and analysts. Positive values indicate optimism, whereas negative ones are a sign of pessimism.

The considered macroeconomic factors (such as the consumer price index, long-term interest rate, and the unemployment rate) are defined for each country separately. As the long-term interest rate, we consider the price of 10-year bonds. We also included the price of oil, as it is well known for its importance for financial markets.

3.2. Results of the study

First of all, we have considered the MS model. Table 2 presents the estimation results, from which we can assume the existence of two states. The first state, which can be identified as prosperity, is characterised by positive mean and a relatively small standard deviation. The second one, identified as a crisis, has usually a twice as big standard deviation and an insignificant or negative mean. Only for the UK, we observe a significant b_2 in the state of high volatility, so a negative return in a $(t-1)$ is going to be exaggerated in the next t, deepening the crisis. In the state of prosperity, the autoregressive parameter is significant only in the USA. Its negative value suggests that in times of prosperity downs follow ups and vice versa, impeding a chain-reaction effect. In the period of prosperity, the average rate of return remained at the same level, although the highest was in Germany (1.4%), the lowest – in Poland (1.1%). On the other hand, during crisis, the average of returns was the lowest in Germany (–1.7%), and the highest in Turkey (0.05%), although they weren't significant. At the same time, we should note a very high variance in both regimes on the Turkish market (6.5% and 14%, respectively).

Table 2. Estimated parameters of the Markov-switching model with fixed transition parameters

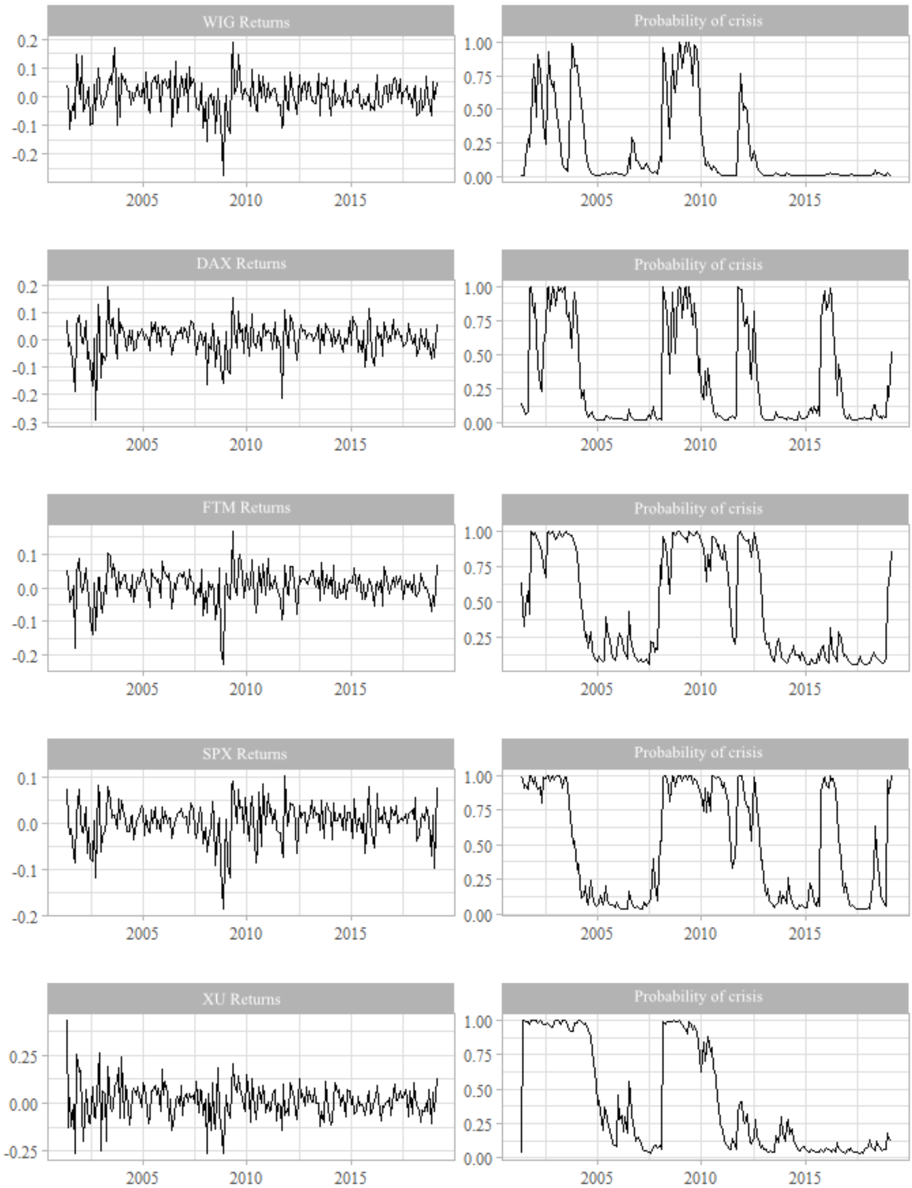
Country	μ_1	b_1	σ_1	μ_2	b_2	σ_2
Poland	0.011***	–0.081	0.043***	–0.001	0.157	0.088***
Germany	0.014***	–0.050	0.035***	–0.017	0.066	0.088***
The UK	0.012***	–0.040	0.032***	–0.007*	0.206**	0.068***
The US	0.013***	–0.167***	0.022***	–0.005	0.167	0.055***
Turkey	0.012***	0.053	0.065***	0.005	–0.202	0.140***

Note. The results of the Markov-switching model parameters, estimated using Hamilton filtering. The model switches between two AR(1) processes, each described by constant μ, an autoregressive parameter b, and a standard deviation of errors σ. The significant parameters at 1% are marked by ***, at 5% by **, at 10% by *.

Source: authors' calculation.

To verify the assumption that the volatility of returns during the crisis period is much greater than in the prosperity period, a restrictive test was performed. As these states are identified primarily by the change in variance, LM tests were performed against models with the restriction that standard deviations in both states are equal. In each country, the difference turned out to be significant (p-value < 0.01). In each case, the null hypothesis of homogeneity of variance was rejected. This proves that returns come from two distributions with significantly different variances. Subsequently, in order to evaluate the goodness-of-fit of the MS model, the test described in the previous section was carried out. For all cases we obtained a p-value ≥ 0.05, so the Markov-switching between the two AR(1) models is here an appropriate description.

Figure 1. Returns' volatility (left panel) and conditional probabilities of being in the second regime (associated with crisis) from the MS model (right panel)



Source: authors' calculation.

Figure 1 shows the volatility of returns (left panel) and the conditional probability of being in the second regime obtained from the MS model (right panel). The graphs in Figure 1 allow the conclusion that the high values of conditional probability indicate the periods when high volatility is observed. The financial literature suggests that the high return volatility is driven mainly by a rising uncertainty in the stock market (Ang & Bekaert, 2002; Forbes & Chinn, 2004; Longin & Solnik, 1995; Ramchand & Susmel, 1998). Therefore, the states display a close link with the mood on stock markets.

The first common period of high volatility can be related to the crash of the dot-com bubble, which was caused by excessive speculation in internet-based companies at the end of the 20th century. After a few peaceful years, the conditional probability of being in the second regime has increased around 2007, which marked the beginning of the world-wide financial crisis, followed by a severe recession. The strong, conditional probability of being in this regime peaked around 2008 when the volatility of returns was particularly high, which was connected with the bankruptcy of the Lehman Bank. The years 2010–2012 was also a period of high volatility in effect of the fiscal problems in the EU.

Table 3. Correlations between indices' returns and exogenous variables

Index	VIX	TED spread	ZEW	Unemployment	CPI	Oil price returns	Long-term interest rate
WIG	-0.495^{***}	-0.071	0.238^{***}	-0.041	0.024	0.259^{***}	-0.268^{***}
SPX	-0.758^{***}	-0.091	0.142^{*}	-0.171^{**}	0.092	0.289^{***}	0.358^{***}
XU	-0.355^{***}	-0.103	0.128^{*}	-0.097	0.005	0.148^{*}	-0.586^{**}
DAX	-0.619^{***}	-0.076	0.115^{*}	-0.120^{**}	0.044	0.158^{*}	0.304^{***}
FTM	-0.672^{***}	-0.202^{**}	0.178^{**}	-0.029	0.079	0.291^{**}	0.125^{**}

Note. The table presents the correlations between indices' returns and exogenous variables. The stationarity of the exogenous data has been verified by means of the ADF test. As they are not stationary (except ZEW), the difference of the first order is used. The significant parameters at 1% are marked by ***, at 5% by **, at 10% by *.

Source: authors' calculation.

Table 3 presents the correlations between the increments of the VIX, TED spread, the ZEW, unemployment, the CPI, oil price returns, and long-term interest rates, with return rates of stock indices. These correlations may serve as an initial step in the analysis, suggesting what can be expected of the coefficients in the final models. A relatively high (as an absolute value) negative correlation can be observed between the VIX and all the considered indices' returns. This correlation is higher for developed markets than for emerging ones. The correlation coefficients with TED spread are rather small (ranging from –0.202 to –0.071) and insignificant (except for the FTM). The correlation coefficients with the ZEW index are moderate (from 0.115 to 0.178).

The lowest coefficient is observed for the DAX index, while the highest for the WIG. This may mean that when the consumer sentiment in Germany is optimistic, the rates of return on the Polish stock exchange are likely to increase relatively more than in Germany. The negative correlation coefficients with increments of the unemployment rate are insignificant. The highest one (as an absolute value) is for the SPX. Insignificant correlation coefficients are obtained for the CPI factor. The oil price returns are relatively highly correlated with stock market indices (the highest correlation coefficient is for the UK and the US, while the lowest for Turkey). Long-term interest rates seem to be the second most important factor (after VIX). The data indicate a positive correlation for developed markets, while a negative one for Poland and Turkey. Moreover, for the latter, it seems to be the most strongly correlated factor (-0.586).

In the next stage of the research, we verify which of the factors affect the transition between states. For this purpose, we use the Markov-switching model with a time-varying matrix transition probability, where transition probabilities p_t^{ii}, $(i = 1, 2)$, are the logistic function of $x_t^T \beta_i$, where: $x_t^T \beta_i = \beta_{0i} + \beta_{1i} Z_t$, and Z_t denotes a given factor. If there is no statistically meaningful impact of this factor on the stock market, then the TVPMS model converges to the Markov-switching model with fixed transition parameters. Therefore, for each case, we tested the null hypothesis of the Markov-switching model with fixed transition parameters against the alternative of the model with time-varying transition parameters.

Table 4. Estimated parameters of the TVPMS model and LM statistics

Factors	μ_1	b_1	σ_1	μ_2	b_2	σ_2	β_{01}	β_{11}	β_{02}	β_{12}	LM
Poland											
VIX	0.019	−0.164	0.049	−0.071	0.232	0.078	8.681	**−2.099**	−4.193	2.300	**39.658**
	0.000	0.048	0.000	0.190	0.078	0.000	0.061	**0.070**	0.503	0.469	**0.000**
TED SPREAD	0.012	−0.170	0.041	−0.007	0.308	0.086	4.847	**−0.317**	0.810	−0.037	**14.411**
	0.002	0.148	0.000	0.381	0.107	0.000	0.008	**0.082**	0.320	0.107	**0.000**
ZEW	0.012	−0.065	0.043	−0.004	0.159	0.084	4.357	**0.128**	2.764	−0.009	**7.465**
	0.003	0.469	0.000	0.729	0.231	0.000	0.015	**0.098**	0.017	0.738	**0.024**
Unemploy-ment	0.013	−0.076	0.044	−0.009	0.179	0.087	3.309	30.511	1.885	**37.295**	**5.582**
	0.003	0.442	0.000	0.560	0.241	0.000	0.000	0.254	0.017	**0.022**	**0.061**
CPI	0.013	−0.074	0.044	−0.008	0.166	0.085	4.262	**11.072**	2.113	0.984	2.066
	0.003	0.442	0.000	0.535	0.266	0.000	0.013	**0.098**	0.009	0.701	0.355
Oil	0.011	−0.050	0.042	−0.005	0.169	0.088	2.871	−3.006	3.887	−6.527	0.439
	0.006	0.546	0.000	0.722	0.255	0.000	0.013	0.412	0.014	0.850	0.803
Interest rate	0.011	−0.076	0.043	−0.005	0.155	0.087	3.662	**−1.963**	2.079	−0.387	**53.034**
	0.018	0.921	0.000	0.722	0.368	0.000	0.001	**0.004**	0.012	0.837	**0.000**

Table 4. Estimated parameters of the TVPMS model and LM statistics (cont.)

Factors	μ_1	b_1	σ_1	μ_2	b_2	σ_2	β_{01}	β_{11}	β_{02}	β_{12}	LM
USA											
VIX	0.022	–0.091	0.025	–0.045	0.319	0.034	4.901	**–2.059**	–0.401	**0.768**	**81.388**
	0.000	0.055	0.000	0.000	0.002	0.000	0.004	**0.003**	0.251	**0.003**	**0.000**
TED SPREAD	0.013	–0.093	0.023	–0.013	0.144	0.061	4.161	**–0.244**	1.451	–0.098	**9.639**
	0.000	0.284	0.000	0.177	0.324	0.000	0.002	**0.012**	0.016	0.241	**0.008**
ZEW	0.012	–0.117	0.023	–0.005	0.174	0.057	3.003	**0.068**	2.576	0.016	**5.363**
	0.000	0.257	0.000	0.406	0.116	0.000	0.000	**0.055**	0.000	0.387	**0.068**
Unemploy-ment	0.013	–0.166	0.023	–0.004	0.166	0.054	4.602	–22.807	2.935	3.407	4.020
	0.000	0.118	0.000	0.543	0.121	0.000	0.080	0.165	0.000	0.768	0.133
CPI	0.011	–0.158	0.022	–0.004	0.172	0.056	3.462	0.150	2.836	0.053	2.004
	0.000	0.143	0.000	0.510	0.132	0.000	0.000	0.381	0.000	0.814	0.367
Oil	0.014	–0.106	0.023	–0.011	0.113	0.058	5.264	**29.204**	1.939	–13.066	**9.272**
	0.000	0.254	0.000	0.189	0.410	0.000	0.001	**0.016**	0.004	0.203	**0.009**
Interest rate	0.019	–0.173	0.028	–0.039	0.211	0.043	4.694	**18.273**	–0.213	**–3.117**	**11.189**
	0.000	0.013	0.000	0.000	0.156	0.000	0.013	**0.019**	0.677	**0.044**	**0.004**
Turkey											
VIX	0.013	0.062	0.066	0.001	–0.218	0.147	4.190	–0.130	3.729	0.256	1.810
	0.015	0.442	0.000	0.949	0.120	0.000	0.000	0.755	0.004	0.186	0.405
TED SPREAD	0.013	0.064	0.066	–0.001	–0.212	0.144	28.397	–0.946	3.618	0.036	**14.634**
	0.014	0.431	0.000	0.945	0.117	0.000	0.829	0.787	0.000	0.203	**0.002**
ZEW	0.012	0.015	0.065	0.009	–0.171	0.144	4.414	**0.052**	4.585	–0.036	**4.678**
	0.030	0.866	0.000	0.660	0.206	0.000	0.000	**0.092**	0.079	0.438	**0.096**
Unemploy-ment	0.009	0.057	0.067	–0.075	–0.416	0.124	14.944	–23.89	4.903	0.157	4.414
	0.026	0.380	0.000	0.072	0.180	0.000	0.547	0.657	0.000	0.804	0.110
CPI	0.012	0.074	0.066	0.002	–0.019	0.136	4.602	–0.790	2.115	3.720	3.464
	0.016	0.656	0.000	0.908	0.164	0.000	0.000	0.864	0.012	0.266	0.176
Oil	0.013	0.041	0.066	0.002	–0.192	0.143	4.484	2.005	6.440	–31.925	3.387
	0.020	0.791	0.000	0.878	0.144	0.000	0.000	0.972	0.034	0.109	0.183
Interest rate	0.050	–0.325	0.034	–0.045	–0.139	0.040	14.312	–42.08	0.814	**2.509**	**41.575**
	0.000	0.001	0.000	0.000	0.736	0.000	0.862	0.734	0.244	**0.070**	0.000
Germany											
VIX	0.025	–0.097	0.039	–0.067	0.151	0.061	4.293	**–1.458**	–0.192	**0.493**	**43.752**
	0.000	0.387	0.000	0.000	0.020	0.000	0.001	**0.004**	0.119	**0.027**	**0.000**
TED SPREAD	0.013	–0.038	0.035	–0.017	0.065	0.090	3.542	**–0.038**	2.060	–0.003	2.219
	0.000	0.274	0.000	0.083	0.252	0.000	0.000	**0.034**	0.000	0.862	0.329
ZEW	0.013	–0.047	0.036	–0.018	0.070	0.090	3.075	**0.071**	2.704	–0.006	**7.299**
	0.000	0.298	0.000	0.075	0.289	0.000	0.000	**0.038**	0.001	0.756	**0.026**
Unemploy-ment	0.012	–0.007	0.036	–0.018	0.075	0.094	3.958	17.331	2.545	**28.959**	**5.720**
	0.000	0.925	0.000	0.188	0.581	0.000	0.001	0.135	0.009	**0.023**	**0.057**
CPI	0.012	–0.016	0.036	–0.017	0.061	0.093	3.004	1.000	4.618	**–6.356**	**5.032**
	0.000	0.850	0.000	0.163	0.635	0.000	0.000	0.491	0.065	**0.093**	**0.080**
Oil	0.013	–0.029	0.038	–0.024	0.041	0.093	3.538	15.226	4.345	–16.230	**5.075**
	0.000	0.711	0.000	0.124	0.779	0.000	0.000	0.122	0.018	0.130	**0.079**
Interest rate	0.012	–0.024	0.037	–0.019	0.068	0.093	4.741	**13.112**	2.388	–0.810	**6.701**
	0.001	0.768	0.000	0.176	0.611	0.000	0.000	0.034	0.000	0.848	**0.035**

Table 4. Estimated parameters of the TVPMS model and LM statistics (cont.)

Factors	μ_1	b_1	σ_1	μ_2	b_2	σ_2	β_{01}	β_{11}	β_{02}	β_{12}	LM
						UK					
VIX	0.022	0.055	0.032	−0.049	0.603	0.043	2.744	**−2.080**	−2.169	1.091	**70.155**
	0.000	0.116	0.000	0.000	0.000	0.000	0.007	0.041	0.052	0.232	**0.000**
TED SPREAD	0.012	0.027	0.035	−0.046	0.632	0.075	2.725	**−0.137**	−0.829	−0.006	**10.513**
	0.000	0.737	0.000	0.071	0.014	0.000	0.003	0.038	0.271	0.9877	**0.005**
ZEW	0.012	−0.002	0.032	−0.007	0.210	0.072	4.901	**0.177**	3.168	0.089	3.949
	0.000	0.620	0.000	0.000	0.100	0.000	0.856	0.078	0.104	0.204	0.138
Unemploy-ment	0.010	0.001	0.033	−0.036	0.753	0.072	2.505	−15.242	2.284	3.575	2.832
	0.002	0.826	0.000	0.081	0.045	0.000	0.045	0.189	0.004	0.785	0.243
CPI	0.011	0.0041	0.032	−0.010	0.188	0.072	1.195	0.009	−1.526	−0.610	3.203
	0.000	0.981	0.000	0.007	0.001	0.000	0.195	0.574	0.059	0.122	0.202
Oil	0.014	−0.014	0.031	−0.032	0.739	0.064	1.213	**9.384**	−1.606	−6.604	**6.686**
	0.000	0.665	0.000	0.017	0.012	0.000	0.069	0.098	0.032	0.101	**0.035**
Interest rate	0.012	−0.047	0.034	−0.019	0.178	0.080	6.071	**24.711**	−0.287	−1.723	**9.707**
	0.000	0.692	0.000	0.343	0.292	0.000	0.006	0.034	0.324	0.473	**0.007**

Note. The table shows the TVPMS model parameters with their p-values, estimated using Hamilton filtering. The model switches between two AR(1) processes, each described by a constant μ, an autoregressive parameter b, and a standard deviation of errors σ. The transition probabilities matrix is described by parameters $\beta_1 = [\beta_{01}, \beta_{11}]$, $\beta_2 = [\beta_{02}, \beta_{12}]$. LM is a test statistic, used to compare models from Table 4 against respective models with fixed transition probabilities, presented in Table 2. The significant beta parameters and high LM values are marked in **bold**.

Source: authors' calculation.

The full set of the estimated model parameters with corresponding p-values is presented in Table 4. For all cases, we obtained statistically significant volatility parameters $\sigma_{i,}(i = 1, 2)$. We want to pay special attention to columns β_{11}, β_{12}, and LM. Parameter β_{11} provides information on how the values of the exogenous variable affect the probability of staying in the first state (prosperity), whereas β_{12} – in the second state (crisis). A significant LM statistic indicated that the adoption of a dynamic transition probability matrix, based on the exogenous variable, actually improved the model.

When analysing the results presented in Table 4, it can be noticed that, in general, the factor of the greatest importance is the VIX. For almost every country we obtained a high value of the LM statistic and p-value $= 0.000$. It turned out insignificant only for the Turkish market. The statistical significance of the β_{11} or β_{12} coefficients shows the influence of this factor on probabilities p_t^{11} or p_t^{22}. For the USA and Germany, the VIX impacts both p_t^{11} and p_t^{22}. The negative sign of parameter β_{11} indicates that an increase in the VIX values weakens the probability of staying in the first regime, whereas the positive sign of parameter β_{12} indicates that a decrease in the VIX values weakens the probability of staying in the second state. In the USA and Germany, a rising VIX index not only indicates a high probability of an oncoming crisis in the state of prosperity but also a low probability of rebounding in the state of crisis. For the Polish and British markets, only parameter β_{11} proves significant.

For all the studied countries, we also performed tests for Granger's causality and found that the value of the VIX has an impact on transition probabilities in the following month.

Similar results were obtained for the TED spread factor. As noted in Table 3, the increases in this indicator are very weakly correlated with returns. Despite this, when analysing results collected in Table 4, we observed a relatively high value of the LM statistic and p-value < 0.05 in most of the countries. However, the fact that only β_{11} is significant, shows that it affects only the p_t^{11} probability, i.e. the probability of staying in the first regime. The negative sign of this parameter implies that an increase in the TED spread values weakens this probability.

In other words: an increasing 'fear index' or credit risk is a sign that the market is more likely to shift into a state of crisis. For the Turkish market, we have not observed the TED spread factor's importance on probabilities p_t^{11} or p_t^{22}. However, the inclusion of this indicator in the model significantly improves the parameter estimation ($LM = 14.634,\ p$-value < 0.05).

The results presented in Table 4 indicate that also the ZEW index plays an important role in market volatility modelling. We can notice that β_{11} differs significantly from zero for all the analysed markets. The positive sign of this parameter indicates that the higher the expectations, the greater the probability that the market will remain in the first state. This means that, in general, investors and analysts have accurate information at their disposal on the state of the economy. The highest β_{11} parameters were observed for Poland and the UK. However, in the case of the UK, an additional inclusion of the ZEW index in the model does not significantly improve the accuracy of the parameter estimation ($LM = 3.949, p$-value $= 0.138$).

As regards the macroeconomic factors, it should be noted that the unemployment rate was important only in Poland and Germany, while CPI solely in Germany. In Poland, unemployment seems to have a significant positive impact on the probability of staying in the second state ($\beta_{12} = 37.295$). In Germany, a similar pattern is observed, i.e. parameter β_{12} also significantly diverges from zero ($\beta_{12} = 28.96$). The estimates of these parameters suggest that when the unemployment rate decreases, the probability of staying in crisis also decreases. For the remaining countries, both beta parameters are insignificant, but signs of their estimates reveal the fact that unemployment has a negative impact on the markets. This trend is common among all the countries, which corresponds with the theoretical expectations. Unemployment is a strong determinant of the condition of an economy and its rapid growth may indicate an economic downturn.

In Poland, the CPI seems to have a positive impact on the probability of staying in the first regime, although the overall model is not significantly better than the MS. In Germany, a considerably negative parameter β_{12} signifies that when the inflation rate decreases, the probability of remaining in the second state increases. The same direction was observed for other developed markets, but both beta parameters are insignificant. Summing up, the CPI growth seems to have a favourable impact on stock exchanges.

This part of the paper is devoted to a discussion on the impact of the oil prices returns on the stock market. This factor is important for all of the analysed developed markets. As we can notice, oil prices returns have a significantly positive impact on the probability of staying in the first state in the USA,[2] Germany and the UK. Oil prices returns also have a slightly negative impact on the probability of remaining in the second state, as expected. The existing positive relationship between oil prices and assets prices was documented by Apergis & Miller (2009), Ferson & Harvey (1994), Huang et al. (1996), Kilian & Park (2009), Narayan & Narayan (2010) and others.

For all the analysed developed markets, long-term interest rates are also important. Statistically significant parameters β_{11} (for the USA $\beta_{11} = 18.273$; for Germany $\beta_{11} = 13.122$; for the UK – $\beta_{11} = 17.711$) indicate a positive impact on the probability of remaining in the first state. For the USA, the rise in long-term rates is related to the decreasing probability of staying in the state of crisis. For other developed markets, negative parameter β_{12} suggests the direction of change in the probability of being in the second state; however, these parameters seem to be insignificant.

We also found long-term interest rates important for the modelling of emerging markets. However, the direction of this relationship is quite the opposite to that of the developed economies. For the Turkish market returns, parameter β_{12} is significantly greater than zero, and parameter β_{11} less than zero, but it is insignificant. This means that interest rates negatively impact the market, especially in bad economic times. Poland follows a similar pattern, however here parameter β_{11} equals significantly less than zero and parameter β_{12} (estimate is greater than zero) is insignificant, so in this case, the relationship in the first regime, i.e. in the period of prosperity, is more meaningful.

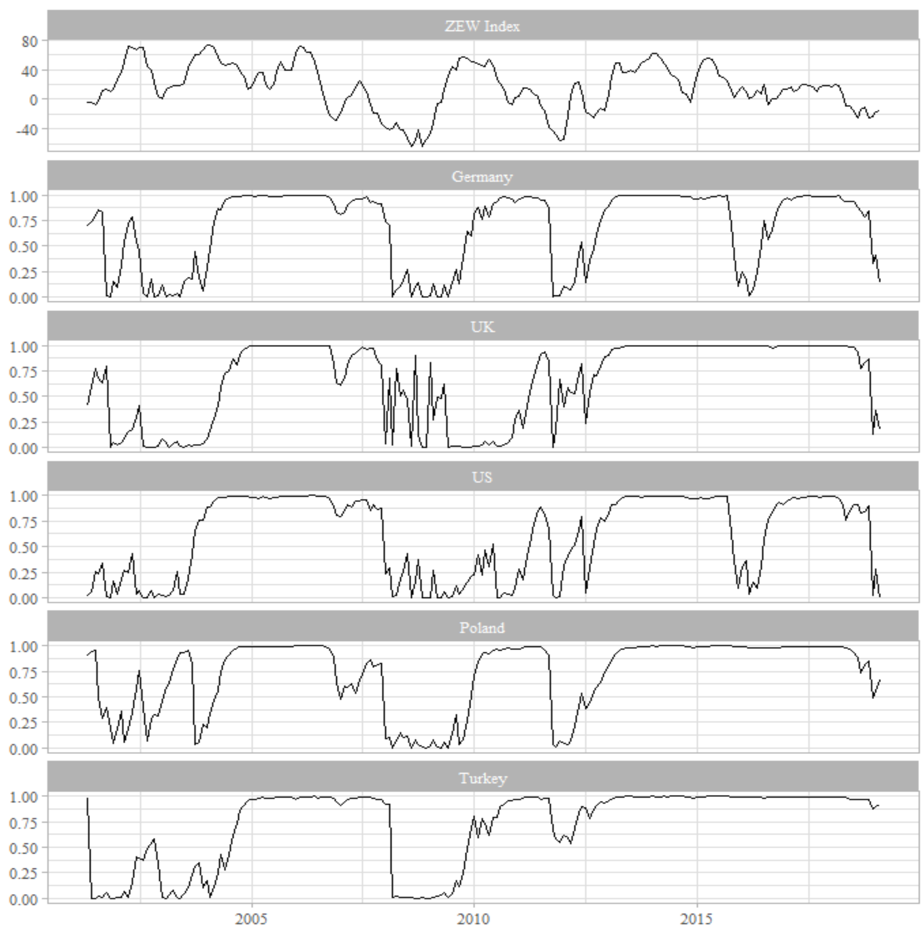
To sum up, most factors are important for developed markets. We have found that high prices of oil, 10-year bonds, and the ZEW index can be connected with a high probability of staying in the first state, whereas an increase in the VIX index

[2] These results are similar to the findings of Chen et al. (1986), according to which the growth rate of oil prices impacts positively the expected returns (however, a significant impact was observed only between 1956-67).

and the TED spread significantly reduces the probability of staying in this state. The ZEW and LTI factors are also important for both emerging markets. Furthermore, for the Polish stock market, as well as for German, domestic macroeconomic factors are significant. A high unemployment rate indicates a high probability of a crisis persisting, whereas a high inflation rate can be connected with a more probable recovery.

Figure 2 shows changes in the ZEW index, compared to the probabilities of staying in the first regime. We can notice that the ZEW index is associated with the probabilities for all the analysed markets. We have noted that the decline in the ZEW index is reflected in the decreases in the probability of staying in the first regime.

Figure 2. The ZEW index and the probability of staying in the first regime (associated with prosperity), obtained from a TVPMS model with the ZEW index as an exogenous variable



Source: authors' calculations.

As expected, in the case of Germany a very high similarity is observed in the changes of the ZEW index and the probability of remaining in the first state. The largest drop in the ZEW index was noted in 2007–2009 and 2001–2012. There was a decrease in the value of this index after 2015 and 2018. These decreases were reflected in the decrease in the probability of remaining in the first regime. A similar pattern of the relationship occurred for the USA, Poland and Turkey.

4. Conclusions

The aim of the study was to check which variables affect regime shifts. Its contribution is the verification of the thesis that both global factors (such as the VIX, TED spread, oil prices, the ZEW index) and selected macroeconomic variables (e.g. the consumer price index, long-term interest rates or the unemployment rate) may be important for the state of volatility of markets. Particular attention has been devoted to the impact of the ZEW sentiment factor on the markets. To the authors' best knowledge, this factor has not been widely examined yet. The applied methodology allowed the analysis of the importance of the factors in each state (prosperity or crisis) separately. The application of the TVPMS model in practice enabled the determination whether the examined factors are of greater importance in the period of prosperity or in the period of crisis. And again, as far as the authors know, such study has not been conducted for the Polish or Turkish market before. There has also been very little research done on the ZEW index so far.

The analysis revealed that there is no uniform and general set of indicators influencing market volatility. In the case of large, developed markets such as the USA, Great Britain or Germany, a wide range of the considered exogenous indicators have some impact on the returns dynamic. We have discovered that high returns of prices of oil, 10-year bonds, and the ZEW index can be related to the high probability of staying in the first state, whereas an increase in the VIX index and the TED spread significantly reduces the probability of remaining in this state. The positive impact of 10-year bonds and the ZEW index on the market was discussed by Hüfner & Schröder (2002), Kvietkauskienė & Plakys (2017) and others.

The ZEW and 10-year bonds indicators have proven important not only for the developed markets that were analysed in this study, but also for the two emerging ones. Although the research showed a positive impact of the ZEW index on market volatility, it also indicated the opposite relationship between 10-year bonds and rates of return than in the case of the developed markets.

Domestic macroeconomic factors play an important role for the Polish and German stock markets. A high unemployment rate indicates a strong probability of a crisis persisting, whereas a high inflation rate usually signals a greater probability of economic recovery.

The article considered three developed and two emerging markets. On their basis, some observations could have been made. However, the formulation of a more general conclusion (relating to the difference between developing and developed markets) requires a much wider study, which the authors decided to carry out in their subsequent research.

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Does the slope of the yield curve of the interbank market influence prices on the Warsaw Stock Exchange? A sectoral perspective

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Abstract. The interest rate curve is often viewed as the leading indicator of economic prosperity in a broad sense. This paper studies the ability of the slope of the yield curve in the term structure of interest rates to impact the sectoral indices on the Warsaw Stock Exchange, using daily data covering the period from 1 January 2001 to 30 September 2020. The results of the research indicate an ambiguous dependence of the logarithmic rates of return of sub-indices on the change of the interbank interest rate curve. The only sectors showing a clear relationship of this type is energy and pharmaceuticals.

Keywords: stock market sub-indices, EGARCH, term structure of the interest rates

JEL: C58, E43, E44

1. Introduction

An interest rate is a primary short-term instrument at work in conventional macroeconomic models. As a part of the monetary transmission mechanism, it is one of many channels through which monetary policy operates (Kuttner & Mosser, 2002). Additionally, the monetary transmission mechanism incorporates the relationship between interest rates and the values of real and financial assets. The monetary transmission mechanism within market-oriented economic systems is defined by official short-term interest rates (the policy instrument) and various financial asset prices together with banks' and other financial intermidiaries' balance sheet variables (intermediate channels of monetary transmission), as well as by real economic activity and prices (final policy objectives). Monetary policy directly affects the interest rate curve (yield curve), which acts as a leading indicator in predicting the macroeconomic activity over longer horizons of time (Khandwala, 2015). The steepness of the yield curve can be perceived as an indicator of the 'health' of an economy, and shows its position in the business cycle. The yield curve becomes steep at the beginning of the business cycle. This is because the central bank usually keeps short-term interest rates low during economic downturns to stimulate the economy. As growth picks up, long-term rates begin to rise, which steepens the yield curve. The same process can be observed on the interbank market. Eventually, short-term

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rates become higher (due to the central bank's rate hikes), and the growth of interbank market longer-term rates begins to slow down, or the rates start becoming lower. Thus, the shape of the yield curve can flatten or even inverse, in which case a negative slope of the interest curve emerges. This situation indicates a possible slowdown or even a recession in the real economy in the following months. The slope is understood as the difference between the values of the longer-term interest rate versus the shorter-term one.

It has been observed that when the slope of the yield curve turns negative, a recession will most probably follow. Conversely, when the slope of the yield curve becomes positive, the economy begins to recover.

Central banks typically deal with uncertainty about the key relationships describing the economy. One of these relationships is the level of short-term interest rates. Uncertainty about the key interest rate describing the economy leads to a disagreement about the effects of the monetary policy and, in turn, to a disagreement as to the appropriate interest rate setting (Traficante, 2013).

Mishkin (1991) explains the association of yield spread and real economic activity in terms of the productivity of capital and the business cycle. In his interpretation, the real yield spread should be perceived as the difference between the long-run and short-run marginal productivity of capital. As far as the interbank market is concerned, the author sees the real yield spread as the difference between the long-term and the short-term interest rate. This difference is also the source of valuable information on the opportunity costs of employing capital over a longer rather than a short term.

The argument is that at the peak of the business cycle, the utilisation of capacity is at a high level, and short-run capital productivity is higher compared to longer-run capital productivity, since, in the long run, the economic activity is likely to slow down. On the other hand, at the trough, productivity in the short run is low and an upswing in the longer run is expected. Thus, there is a positive relationship between the yield spread and real economic activity.

Similar considerations can be observed on the stock exchange. An increase in the slope of the interest rate curve, implying an increase in capital productivity, is visible in rising stock prices.

In this setting, a robust monetary transmission mechanism works properly even when the policymaker does not know the detailed structure of the possible adjustments within the economy. The policymaker is supported in this task by the financial market participants via the process of adjusting prices.

To the best of the authors' knowledge, no prior study has analysed the impact of the interest rate slope on the returns of the sectoral stock indices. Bhowmik & Wang (2020) presented an up-to-date literature review on the application of the GARCH class models for forecasting variance volatility in financial markets. According to

Alberg et al. (2008), the EGARCH model is the best predictor of daily data returns of the Tel Aviv stock market index. All these papers concern the stock market index volatility. No papers could be found, however, that deal with the impact of the slope of the interest curve.

Generally, there are two types of research relevant to this subject: the first type explores the applicability of the EGARCH model to modelling the volatility of financial processes, and the volatility of indices in particular. The second examines the reaction of prices or rates of return of securities (equities) to the central bank's decisions (Ehrmann & Fratzscher, 2004). The purpose of this paper is to test the impact of the slope of the interest rate curve on the returns of the stock market sub-indices. The formulated research hypothesis assumes that the increasing slope of the market interest rate curve has a significant impact on the yield rates of the sub-indices listed on the Warsaw Stock Exchange. In this study, we are interested in the short-term approach, although the estimation of the long-term relationship between the rates of return of sectoral sub-indices and the slope of the interest rate curve is also possible. This aspect, however, lies outside the scope of our considerations here.

This paper is of an empirical nature and is organised in the following way: Section 2 is devoted to the review of the relevant literature on the possible impact of the interbank market on stocks or stock indices. Section 3 describes the research methodology and the estimation procedure, and Section 4 presents the empirical results. Final conclusions are included in Section 5. All estimations are made in the Gretl program.

2. Literature review

The undertaken research relates to two currents in literature. The first is the monetary transmission mechanism as an institutional framework of the monetary policy. The second relates to asset pricing and market efficiency.

Theoretical relationships between monetary policy and financial markets are complex. Dreger & Wolters (2009) argue that a monetary policy shock, for example in the form of changes in the money supply or interest rates, eventually leads to shifts in investors' portfolios continuing until the relationship between liquidity and asset holdings is re-established. In other words, asset prices should react to changes in the interest rates of the interbank market, and such reactions should be observable. For central banks, the transmission of monetary impulses in increasingly integrated financial markets is of great importance (De Santis, 2008). According to Dale & Haldane (1995), the monetary transmission mechanism within market-oriented economic systems is defined by official short-term interest rates (the policy instrument) and various financial asset prices together with banks' and other financial intermediaries' balance sheet variables (intermediate channels of monetary transmission), as well as by real economic activity and prices (final policy objectives). The yield curve plots yield to maturity against the terms for the otherwise similar fixed-income securities. It is commonly assumed that the yield curve contains useful information. Pelaez (1997), however, claims that disagreement exists as to its nature and importance. The results obtained by Argyropoulos & Tzavalis (2016) are consistent with previous macroeconomic studies (Estrella & Hardouvelis, 1991; Gamber, 1996; Hamilton & Kim, 2002; Moneta, 2005; Wheelock & Wohar, 2009), and first and foremost confirm that the slope factor of the yield curve reflects future changes in the business cycle conditions. Assefa et al. (2017) find statistically significant negative effects of interest rates on stock returns in developed countries.

Interestingly, the money and capital markets are closely interrelated, because most corporations, financial institutions and investors are active in both of them. The study of asset pricing lies at the core of financial economics, and the fundamental finance principle asserts that the asset price equals the discounted future streams of cash flows. In the light of the above, two relevant factors have to be pointed out: the uncertainty of the expected cash flows and changes in the discount rate.

It is commonly assumed that stock prices are determined in a forward-looking manner. The stock prices reflect the private sector's expected future discounted sum of returns on the assets. Changes in asset prices can then result from changes in the expected future dividends, the expected future interest rate which serves as a discount rate, or from changes in the stock returns premium.

Stocks as a class of assets are presumed to be sensitive to macroeconomic conditions. Any aggressive change in stock prices can have negative implications for the economy, which makes the causal relationship between macroeconomic variables and stock returns an intriguing topic in empirical finance (Barakat et al., 2016). Gostkowska-Drzewicka & Majerowska (2018) studied the evidence of the industry's effect on companies listed on the Warsaw Stock Exchange.

Participants of financial markets are also often characterised as being forward-looking. Likewise, financial prices can be considered forward-looking with regard to those macroeconomic variables that can affect them and, therefore, often contain valuable information on their expected or future behaviour (Alonso et al., 2001).

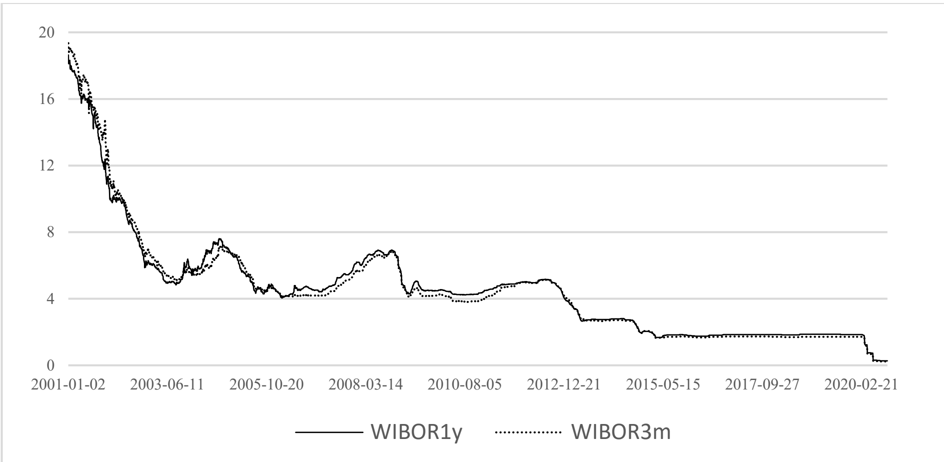
Explaining the relationship between macroeconomic variables and the stock market is important, as the latter has a systematic effect on the former. Economic forces affect discount rates, and through this mechanism, macroeconomic variables become part of the risk factors in equity markets (Chen et al., 1986). In an efficient capital market, stock prices adjust rapidly as new information becomes available; therefore, stock prices reflect all information about the stocks. This means that investors cannot use the readily-provided information to predict stock price movements and make profit by trading shares. In short, an efficient market incorporates new information quickly and completely. Stock prices also reflect the expectations towards the future performance of corporate profit. If stock prices reflect the above assumptions, they should be used as indicators of the economic activity; the dynamic relationship between stock prices and macroeconomic variables can be then used to guide countries' macroeconomic policies (Maysami et al., 2005). The most common means of linking macroeconomic variables with stock market returns is through the arbitrage pricing theory (APT), developed by Ross (1976). According to his theory, multiple risk factors can explain stock returns. The APT assumes that stock prices can be influenced by the behaviour of macroeconomic fundamentals, i.e. there are many channels for the relationships between the stock market and key macroeconomic variables. Chen et al. (1986) found that industrial production, changes in the term structure of interest rates, and changes in risk premiums, were all positively related to the expected stock return. Bernanke & Kuttner (2005) asserted that due to exogenous factors, changes in monetary policy – if not anticipated – affect the volatility of stock prices. Empirical studies point out to the importance of such economic variables as exchange rates, gross domestic product (GDP), basic interest rates, and inflation (Bhuiyan & Chowdhury, 2020). Other studies show explicitly the relationship between stock returns and interest rates (Assefa et al., 2017; Izgi & Duran, 2016; Papadamou et al., 2017). Interbank rates are capable of explaining the adjustments of stock prices, because their changes affect cash flows of companies, and may additionally affect the risk-adjusted discount rate (Flannery & Protopapadakis, 2002). Atanasov (2016) proves that value stocks are highly sensitive to upside movements in interest rate growth, whereas growth stocks rather tend to react to downside movements of interest rates.

3. Data and methodology

The conducted empirical analysis is based on daily data of the WIBOR rates relating to 1-year and 3-month deposits. In practice, the difference in interbank deposit rates for these terms should be seen as a proxy for economic expectations. The greater the positive value of the observed difference, the more favourably interpreted the economic outlook. The 3-month rate is of particular importance for the assessment of the economic outlook, as it is a price-setting parameter for variable-rate loans granted to enterprises. Therefore, any other possible slopes and differences are outside the scope of interest of this study. The analysis extends over the period from 1 January 2001 to 30 September 2020. It produced 4,944 daily observations in total. Figure 1 presents the evolution of these rates (expressed as a percentage) during the considered period. Sharp declines in the initial years can be observed, followed by fluctuations related to the economic crisis of 2008–2009. In the following years, the

WIBOR 1-year and 3-month rates were at a similar level. Subsequent breakdowns were connected with the outbreak of the pandemic and the related economic crisis.

Figure 1. WIBOR 1Y and WIBOR 3M rates



Source: authors' calculation based on data from the Stooq.pl (n.d.) financial website.

Figure 2 shows the differences between 1-year and 3-month rates. Except the initial period of some fluctuation in the banking market (i.e. before Poland joined the European Union in 2004), the slope of the Polish interbank market interest rate curve was mostly positive. The period 2012–2013 deserves particular attention due to the financial crisis in Greece, followed by a banking crisis in Cyprus.

Figure 2. Differences between WIBOR 1Y and WIBOR 3M rates

Source: authors' calculation based on data from the Stooq.pl (n.d.) financial service.

The Warsaw Stock Exchange consists of eight sectors and numerous subsets. Indices, including sector indices, are determined. Currently, 14 sectoral sub-indices are listed. Our analysis focuses on 9 of them, according to the classification adopted by the Stooq.pl portal. These are the sub-indices of the banking, construction, chemical, pharmaceutical, energy, oil and gas, food, and real estate sectors, as well as one macrosector index – the WIG GAMES. The initial listing of particular indices on the stock exchange took place at different times, which resulted in obtaining time series of varied length.

The relevant literature suggests using a GARCH type model for modelling daily returns. The equations of the GARCH(1,1) model, based on Bollerslev (1986), can take the following form:

$$y_t = \mu + \varepsilon_t, \tag{1}$$

$$\varepsilon_t | I_{t-1} = N(0, h_t), \tag{2}$$

$$h_t = \omega + \alpha\varepsilon_{t-1}^2 + \beta h_{t-1}, \tag{3}$$

where y_t is the return series, h_t is the conditional variance, I_t is the set of all available information, and ε_t is the error term. In order to ensure a non-negativity of the conditional variance, the following restrictions are imposed $\omega > 0$, $\alpha \geq 0$, $\beta \geq 0$. The requirement for covariance stationarity of ε_t is $\alpha + \beta < 1$.

Lin (2018) and Wei et al. (2020) apply the exponential GARCH model to describe the volatility asymmetry of returns. The EGARCH model was introduced by Nelson (1991). It allows including the asymmetric impact of positive and negative rates of return on variances. Thus, based on the above-mentioned works, the conditional variance is defined as:

$$lnh_t = \omega + \alpha\{\theta z_{t-1} + \gamma[|z_t| - E(|z_{t-1}|)\} + \beta lnh_{t-1}, \tag{4}$$

where $z_t = \varepsilon_t/\sqrt{h_t}$. It is not necessary to introduce any restrictions on the parameters of such equations, mainly due to the logarithmic form of the last equation.[1]

To test the impact of the interest rate slope on the returns of sub-indices, the additional variable to equation (1) of the above model was introduced. So the equation takes the form:

$$y_t = \mu + \delta x_t + \varepsilon_t, \tag{5}$$

[1] An overview of the GARCH class models can be found, for example, in Fiszeder (2009).

where x_t represents the slope measured by the differences between one-year and three-month market interest rates (WIBOR 1Y minus WIBOR 3M).

Additionally, we introduced the difference x_t, defined above, into the variance, giving:

$$lnh_t = \omega + \upsilon x_t + \alpha\{\theta z_{t-1} + \gamma[|z_t| - E(|z_{t-1}|)\} + \beta lnh_{t-1}. \tag{6}$$

As mentioned before, the application of GARCH class models for financial data can be found in the literature. For example, Ugurlu et al. (2014) modelled stock market returns volatility for data from Bulgaria, the Czech Republic, Hungary, Poland and Turkey, applying the GARCH class models. Fałdziński et al. (2021) used ARCH models to forecast energy commodities. Due to the 'fat tails' of returns distributions, it is suggested to apply the t distribution to the conditional distribution of ε_t (Fiszeder, 2009).

4. Results

To determine the relationship between the logarithmic rates of return of selected stock exchange sub-indices and the slope of the interest rate curve, Pearson's linear correlation coefficients were determined. Additionally, correlation coefficients were calculated between the rates of return and differences of interest rates lagged by one period (one day). The results are presented in Table 1.

Table 1. Descriptive statistics of indices' returns[a]

Indices	Variables	
	x_t	x_{t-1}
WIG – banking	0.023	0.010
WIG – construction	−0.066	−0.078
WIG – chemicals	−0.027	−0.033
WIG – pharmaceuticals	−0.082	−0.078
WIG – energy	−0.087*	−0.090*
WIG.GAMES	−0.053	−0.056
WIG – oil and gas	−0.025	−0.026
WIG – food	−0.019	−0.031
WIG – real estate	0.022	0.017

a Data source: Stooq.pl financial service (n.d.).
* Statistically significant at 0.05 significance level.
Source: authors' calculation.

Generally, the correlations are not statistically significant at the 0.05 significance level. The only exception is the dependency between the return of the WIG – energy and the differences in the interest rates. It is not a surprising result, and it is consistent with the findings of Atanasov (2016). The energy-sector companies tend to be perceived as value stocks. In their financial statements, an important cost item is interest payments on external capital, and on their balance sheets, it is long-term debt with financial institutions. This results from a relatively high degree of financing fixed assets with long-term debt and hence the direct sensitivity to changes in interest rates is observed. Additionally, the value stocks are expected to pay dividends that refer directly to the level of the WIBOR 1Y rate.

In the next step, the proposed model, described by equations (4) and (5), was estimated. Results are presented in Table 2.

Table 2. Results of the estimation of the EGARCH(1,1) model with the slope in the conditional mean equation[a]

Indices	Conditional mean		Conditional variance				Conditional density	
	μ	δ	ω	α	γ	β	η	λ
WIG – banking	0.000	−0.000	−0.197*	0.123*	−0.042*	0.988*	7.437*	0.016
WIG – construction	0.000	0.000	−0.261*	0.146*	−0.027*	0.983*	5.452*	−0.014
WIG – chemicals	0.001*	0.001	−0.303*	0.154*	−0.041*	0.978*	7.388*	0.018
WIG – pharmaceuticals	0.007*	−0.049*	−0.528*	0.327*	−0.042	0.963*	4.894*	0.031
WIG – energy	−0.000	0.001	−0.206*	0.138*	−0.040*	0.988*	6.607*	0.008
WIG.GAMES	0.002	−0.005	−0.833*	0.357*	−0.085	0.927*	5.359*	−0.068
WIG – oil and gas	0.000*	0.000	−0.184*	0.111*	−0.027*	0.988*	8.269*	0.002
WIG – food	0.000*	−0.000	−0.525*	0.254*	−0.030	0.962*	5.318*	−0.012
WIG – real estate	0.000*	−0.000	−0.150*	0.124*	−0.024*	0.994*	6.656*	−0.010

a Data sources: Stooq.pl financial service (n.d.).

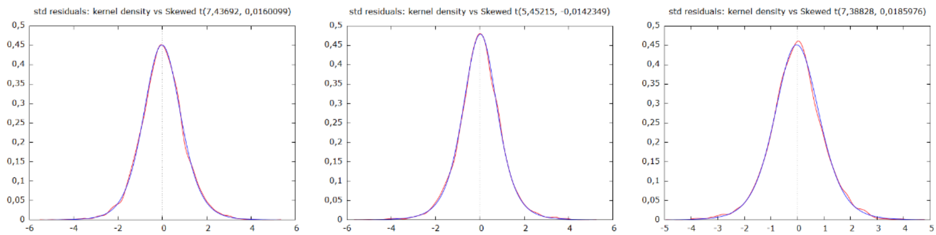
* Statistically significant at 0.05 significance level (the result of 0.000 means that the estimated value of the parameter is below 0.0005).

Source: authors' calculation.

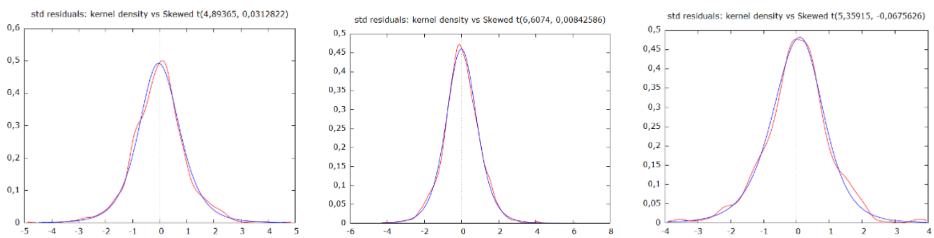
The last two columns in Table 2 feature estimated parameters of the conditional density, assuming skewed t distribution of error terms $t(\eta, \lambda)$. Parameters of the distribution were estimated jointly with the EGARCH parameters. The parameter responsible for the asymmetry of distribution is insignificant in all cases. Figure 3 contains the distribution of the residuals obtained on the basis of the estimated models presented in Table 2. High leptokurtosis of distributions is clearly visible in all the pictures.

Figure 3. Distribution of residuals obtained from the models presented in Table 2

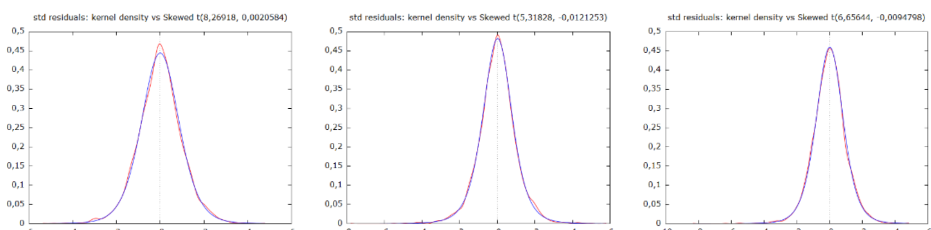
(a) WIG – banking (b) WIG – construction (c) WIG – chemicals



(d) WIG – pharmaceuticals (e) WIG – energy (f) WIG.GAMES



(g) WIG – oil and gas (h) WIG – food (i) WIG – real estate



Source: authors' calculation.

It can be observed that the estimates of most of the structural parameters connected with the volatility were statistically significant. It confirms the rationality of applying this model and is consistent with the results obtained by Wei et al. (2020). The exception is the gamma parameter which proved statistically insignificant for two sectoral sub-indices and one macrosector index. The structural parameter connected with the WIBOR slope was insignificant in all cases, except the one for the WIG – pharmaceuticals. This indicates that rates of return of sub-indices do not respond to changes in the shape of the interest rates curve.

Table 3. Results of the estimation of the EGARCH(1,1) model with the slope in the conditional mean and variance equations[a]

Indices	Conditional mean		Conditional variance					Conditional density	
	μ	δ	ω	υ	α	γ	β	η	λ
WIG – banking	0.000	–0.000	–0.197*	0.000	0.123*	–0.042*	0.988*	7.435*	0.016
WIG – construction	0.000	0.000	–0.261*	0.000	0.146*	–0.027*	0.983*	5.452*	–0.014
WIG – chemicals	0.001*	0.001	–0.302*	–0.003	0.154*	–0.041*	0.978*	7.394*	0.018
WIG – pharmaceuticals	0.003	–0.025	–2.096*	–6.355*	0.590*	–0.085	0.696*	5.203*	0.013
WIG – energy	–0.000	0.001	–0.237*	–0.038*	0.143*	–0.041*	0.984*	6.650*	0.008
WIG.GAMES	0.000	0.001	–0.852*	–0.215	0.358*	–0.088	0.921*	5.313*	–0.068
WIG – oil and gas	0.000*	0.000	–0.184*	0.001	0.111*	–0.027*	0.988*	8.266*	0.002
WIG – food	0.000*	–0.000	–0.528*	0.005	0.255*	–0.030*	0.962*	5.322*	–0.012
WIG – real estate	0.000*	–0.000	–0.150*	–0.004	0.123*	–0.025*	0.994*	6.668*	–0.010

a Data sources: Stooq.pl financial service (n.d.).

* Statistically significant at 0.05 significance level (the result of 0.000 means that the estimated value of the parameter is below 0.0005).

Source: authors' calculation.

Adding the WIBOR slope to the conditional variance, the model was estimated again, and the results are presented in Table 3 (equations (5) and (6)). They generally confirm our previous findings. For this model, the WIBOR slope was insignificant for all the analysed series in the conditional mean, and significant only for two sectoral sub-indices: the WIG – pharmaceuticals and the WIG – energy in the conditional variance.

5. Conclusions

The aim of the study described in this paper was to show whether the slope of the interest rate curve affects the formation of the rates of return of selected sub-indices listed on the Warsaw Stock Exchange. The conducted analysis, based on Pearson's linear correlation coefficient, showed no direct relationship between these factors. The econometric verification of the estimated EGARCH(1,1) model in two versions, i.e. with the slope in the conditional mean equation and with the slope in the conditional mean and variance equations, confirmed the above findings. It can therefore be stated, on the basis of the significance of the structural parameters, that the sectoral sub-index returns did not respond to changes in the curve's slope. It means that our hypothesis, stated in the introduction, needs to be rejected for most of the sub-indices. Under the monetary transmission mechanism, the credit channel operates for sub-indices (and companies) listed on stock exchanges. Based on the analysis carried out for sub-indices of the Warsaw Stock Exchange, it can be concluded that the quotations will react to changes in the slope of the interest rate curve in a heterogeneous manner. At the same time, we can observe a specific information assymetry, i.e. the energy and pharmaceutical industries reacting more strongly than the others. This is because their expected future profits and cash flow are more affected, as these industries (and companies) will experience higher market costs for re-financing debt after the slope of the interbank market interest rate curve has increased, even without any tightening of the monetary policy.

It should be pointed out that the variation in response to the change in the interest rate slope depends on the characteristics of the industry (sub-indices), to which the individual companies are affiliated.

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