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Risk Transmission Between Sovereign Credit Default Swaps and Government Bonds During the Global Financial Crisis. The Case of the Czech Republic, Hungary and Poland

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Abstract

The goal of the paper is to verify the direction of sovereign risk transmission between sovereign CDS and sovereign bond markets in the Central European economies: the Czech Republic, Hungary and Poland. We focus on the hectic crisis period of 2008-2013. On the one hand, the sCDS market is said to react faster to the news than the sovereign bonds market. On the other hand, the bond market is related more closely to the internal situation of the country than the sCDS one and thus can price the sovereign risk more accurate. Moreover, the relationships between the markets can change during crisis time. We find that in the case of most risky and most indebted economy in Hungary there was a feedback between sCDS and sovereign bonds risk. In the case of Poland sCDS market risk Granger caused the risk of sovereign bonds – if we exclude instantaneous causality from the analysis; when it is included, feedback occurred. Eventually, in the case of the Czech Republic the risk of sCDS market Granger caused risk of the bonds market.

Keywords: sovereign credit default swaps, bond yields, Central and Eastern Europe, risk transmission

JEL Classification: G15, G18, F65

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1 Introduction

During hectic crisis times investors, who have in their portfolios debt instruments of various sovereigns, may want to obtain fast and reliable information about the changes of the risk in the given country. One of such indicators, available at daily frequency, are the yields of sovereign bonds. Such indicators are closely linked to the internal situation of the country (Kocsis, 2014), but may not react immediately to the flow of new information, due to possible difficulties with changing the position in this instrument. An alternative indicator of a given country sovereign risk is a sovereign CDS (further: sCDS) instrument. The sCDS market is documented to react fast to the news, however may not directly reflect the changing risk level of the country due to extreme vulnerability to the change of the global risk (see e.g Longstaff *et al.*, 2005; Longstaff *et al.*, 2011; Plank, 2010, Camba-Mendez and Serwa, 2016; Adam, 2013; Będowska-Sójka and Kliber, 2013; Kliber, 2013).

Over the years there appeared many studies trying to determine the leadership of one market over another. For instance, Giannikos *et al.* (2013) find that CDS market dominates other markets in terms of price discovery, while Coudert and Gex (2010) show that in the case of the high-yield countries the CDS market leads the bond one, while in the case of the low-yield ones the bond market leads the sCDS one.

However, the authors of the papers quoted above investigated the long-run relationships between the sCDS prices and bond yields. In our approach we focus precisely on the risk, which is usually approximated with *volatility* of the instruments obtained from the GARCH-type models. Moreover, the authors investigate the relationships between the sCDS spreads and bond spreads (differences between the yield of the given country bond and the yield of a sovereign bond of the least risky economy in the region), while in our study we analyze bond *yields*.

The advantage of using the yields is that we do not introduce new information connected with the sovereign risk of another entity. The obvious drawback is that the relationship between sCDS spread and sovereign bonds are not precisely defined in the economic theory and we can expect either positive or negative correlation between these two variables. In fact, there are two strands in the literature. For instance, Fontana and Scheicher (2016) refer to Merton (1974) model, where the increase of the risk-free rate implies the growth of the value of the company's assets and reduce the risk of insolvency. Thus, changes of the risk-free rate should be negatively correlated with the changes of credit spreads. Fontana and Scheicher (2016) extend this reasoning to the sCDS spreads and assume negative correlation between the spreads and changes of risk-free rate (although in their empirical study they obtain only insignificant relationships between the two). The opposite approach assumes positive correlation between the interest rate and the sCDS spread, since the growth of the spread is an indicator of the growth of the country credit risk. Thus, investors who lend their money to the government in the form of government bonds, expect higher return (yield). Practitioners note also, that during the episodes of market turmoil, the difference between the high- and low-yield bonds is getting even larger: the yield of the safer bonds declines, while at the same time the yield of the riskier ones grows, and the spread between bonds of different risk category grows (see e.g. Jain 2014).

However, when it comes to *risk*, measured by *volatility* of instruments, we can expect that the growth of volatility of sovereign bonds should be accompanied by the growth of volatility of sCDS. The main research question of this paper is: what is the direction of sovereign risk transmission in the V3 (Hungary, Poland and the Czech Republic) group? Is it the sCDS market that leads the bond one or is it the other way around? Answering the question, we add up to the discussion about the role of sCDS market in crisis transmission in the emerging Central and Eastern European countries, extending such work as e.g. Kliber (2011), Adam (2013), Kliber (2014), Kocsis (2014), Camba-Mendez *et al.* (2016) or Ters and Urban (2018).

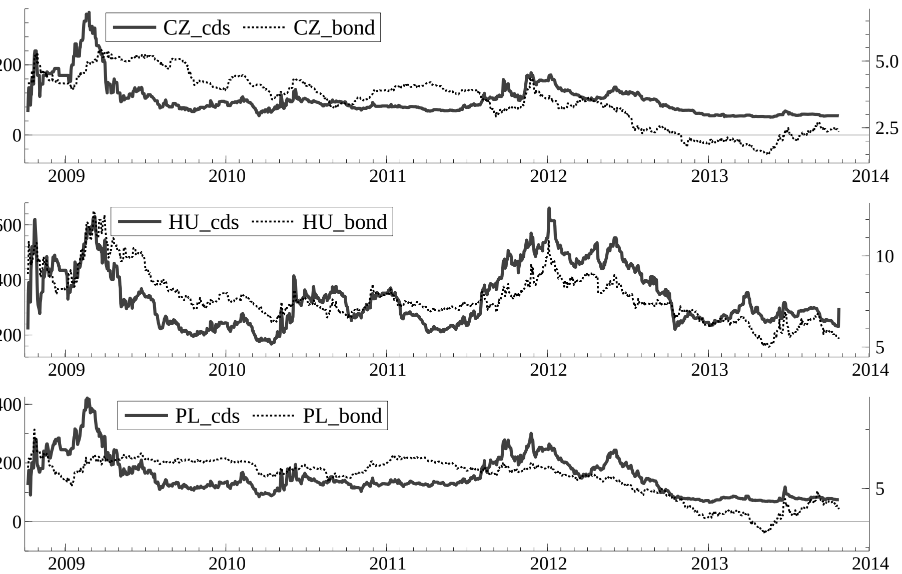
The paper is organized as follows. First, we present the data with the descriptive statistics and we plot the time series. In section 2 we present the methodology and describe estimated GARCH models both for sCDS and bond yields. The Hong test of Granger-type causality between the volatilities of the sovereign bonds and CDS is presented in details. The results discussed in the last section are analyzed with respect to the level of indebtedness and the policies implemented to deal with the public debt in each country, as well as with the level of integration with European markets.

2 Data

We investigated prices of 10-years maturity sCDS quoted in euro for three countries (the Czech Republic, Hungary and Poland) that belong to Visegrad Group: V4 (an alliance of four Central European States: Czech Republic, Hungary Poland and Slovakia; we omit Slovakia as this is the only country that already adopted euro and the risk factors attributed to its bonds may be of slight different nature), as well as the government bonds of 10-years maturity issued by governments of the countries. Our sample covered the period from 01.11.2008 to 24.10.2013 that gives 1316 daily observations. The sCDS data comes from DataStream, while the bond data are taken from www.stooq.pl and CEIC database.

In Figure 1 we present the dynamics of sCDS premia together with the dynamics of the bonds' yields. The greatest similarities between the time series dynamics are observed for Hungary. In the case of Poland and the Czech Republic these two series behave quite differently in each country. Polish sCDS series seem to be much more volatile than the bond yields. In the beginning of the sample the growth of sCDS premia is accompanied with the decline of the bond yield. Started from the second half of 2009 both series seem to follow similar trend, however since 2011 they behave differently: the growth of sCDS premium in 2011 is accompanied by the decrease in yield. Again, the decline of the yield in 2013 is much sharper than the decline of the premium. Similar argumentation applies also to the case of the Czech Republic.

Figure 1: Dynamics of the prices of sCDS and bond-yields: Czech, Hungary and Poland



Note: CZ_cds, HU_cds, PL_cds denote respectively the spreads of the Czech, Hungarian and Polish CDS, while CZ_bond, HU_bond and PL_bond - the yields of the Czech, Hungarian and Polish government bonds

Table 1: The descriptive statistics of the changes of sCDS and bond yields

Statistics:	sCDS			Bond yield		
	CZ	HU	PL	CZ	HU	PL
mean	−0.01	0.06	−0.04	0.00	0.00	0.00
std.deviation	6.25	13.68	8.12	0.06	0.15	0.06
skewness	0.15	1.23	−0.29	1.39	0.00	1.14
Excess kurtosis	25.37	18.13	18.25	14.77	10.42	22.42
	Stationarity test: KPSS					
KPSS test	0.05	0.05	0.07	0.07	0.06	0.07

Note: CZ stands for Czechia, HU for Hungary, PL for Poland. The sample starts on 1.11.2008 and ends on 24.10.2013. The critical values of KPSS test are: 0.74 for α=1% and 0.46 for α=5%

Table 1 presents base descriptive statistics for the sovereign CDS returns and bond yields covered in the study. These statistics show that all mean returns are not significantly different from zero. The volatility proxied by standard deviation is higher for sovereign CDS than for the bonds. The highest volatility is displayed for the Hungarian sCDS, followed by the Polish and Czech one. Among the bond yields, the Hungarian one obtains the highest value. Only the distribution of the Hungarian bond yield is symmetric. Positive estimates of skewness in all but Polish sCDS indicate that in these cases the right tails are longer. All series are characterized by very high excess kurtosis which indicates that their distributions are fat-tailed and confirms the presence of the extreme movements both in sCDS and bond yields. Both skewness and kurtosis show that the distributions of the series are not Gaussian. As the strict stationarity of the series is an important assumption of the Hong (2001) test, we conduct KPSS (Kwiatkowski *et al.* 1992) test for our series. The null hypothesis of stationarity is not rejected in any case. The results are presented in Table 1.

3 Methodology

The goal of the paper is to compare dependencies in volatility of the bond and sCDS market among the three economies. Volatility of financial instruments is especially important when investigating financial markets. It can be related to information flow (see e.g. Ross, 1989). If the information comes in clusters, then the assets can exhibit volatility even if the market instantaneously adjusts to news. Thus, studying volatility spillovers allows for better understanding the patterns of information flow across different markets or market sectors (see also Hong, 2001).
The most common model of volatility of financial series is the GARCH model that captures volatility clustering. The GARCH model was introduced by Bollerslev (1986). We estimated a set of GARCH-type models and based on the diagnostic tests we chose the best one for each series. Following the results of the Pearson's test for goodness of fit, we applied Student t distribution in each case. Thus the following models were estimated:

1. Polish sCDS: ARMA(1,0)-IGARCH(1,1),
2. Czech sCDS: ARMA(0,0)-IGARCH(1,1),
3. Hungarian sCDS: ARMA(0,0)-IGARCH(1,1),
4. Polish bonds: ARMA(1,0)-IGARCH(1,1),
5. Czech bonds: ARMA(0,0)-IGARCH(1,1),
6. Hungarian bonds: ARMA(1,0)-IGARCH(1,1).

The estimated models have the following form:

$$(r_t - \mu) = a_1 (r_{t-1} - \mu) + y_t,$$

$$y_t = \sigma_t \epsilon_t; \quad \epsilon_t \sim iid\, t(\kappa),$$

$$\sigma_t^2 = \omega + \alpha_1 y_{t-1}^2 + \beta_1 \sigma_{t-1}^2; \quad \beta_1 = 1 - \alpha_1,\ \omega > 0$$

where r_t denotes the return of the respective instrument, μ is its unconditional mean, while σ^2 – its conditional variance. The estimation of GARCH models was performed in OxMetrics 7 with G@RCH package (Laurent 2010).
In order to examine the interdependency between bond yields and sCDS spreads we perform Hong test for non-causality in the conditional variance. The Hong test was implemented in R Cran.

3.1 Hong test

The Hong test is an extension and generalisation of the Cheung and Ng (1996) test for non-causality in variance. Lest us consider two strictly stationary and ergodic time series processes: X_t and Y_t, as well as two information sets defined by: $I_t = \{X_{t-j}, j \geq 0\}$ and $J_t = \{X_{t-j}, Y_{t-j}, j \geq 0\}$. According to Granger (1980) Y_t is said to cause X_{t+1} with respect to J_t if :

$$Pr\left(X_{t+1}|I_t\right) \neq Pr\left(X_{t+1}|J_t\right). \tag{1}$$

In practice, definition (1) is too general to test for causality. Therefore, the researchers frequently use the following definition (Cheung and Ng, 1996): Y_t is said to cause X_{t+1} in mean with respect to J_t if :

$$E\left(X_{t+1}|I_t\right) \neq E\left(X_{t+1}|J_t\right). \tag{2}$$

The concept of causality in mean can be extended to the causality in variance. Y_t is said to cause X_{t+1} in variance if:

$$E\left[\left(X_{t+1} - \mu_{x,t+1}\right)^2 |I_t\right] \neq E\left[\left(X_{t+1} - \mu_{x,t+1}\right)^2 |J_t\right], \tag{3}$$

where $\mu_{x,t+1}$is the conditional mean of X_{t+1} (conditioned on I_t). Feedback in variance occurs when X causes Y, and Y causes X.
Let us also suppose that:

$$X_t = \mu_{x,t} + h_{x,t}^{\frac{1}{2}}\varepsilon_t, \quad Y_t = \mu_{y,t} + h_{y,t}^{\frac{1}{2}}\xi_t. \tag{4}$$

In the model above $\mu_{z,t}$ denotes the conditional mean, while, $h_{z,t}$ the conditional variance of variable z. It is worth noting, that there is a difference between the approach of Cheung and Ng (1996) and Hong (2001) in the definition of the set of information, on which the mean and variance are conditioned. As already stated, Cheung and Ng define $\mu_{x,t+1}$ as the conditional mean of X_{t+1}, conditioned on I_t. On contrary, Hong (2001) points out that very rarely $\mu_{x,t+1} = E\left(X_{t+1}|I_t\right)$. Therefore, in the Hong (2001) version of the test: $\mu_{x,t+1} = E\left(X_{t+1}|J_t\right)$, while the conditional variance of each process is measured with respect to its own information set (i.e. $h_{x,t}^{\frac{1}{2}}$ is conditioned on I_{t-1}, while $h_{y,t}^{\frac{1}{2}}$ – on $J_{t-1}\backslash I_{t-1}$). Such a specification suggests that

the VAR model for returns could be implemented in the test. In our study we used the AR-type models for specification of the conditional mean (see the next subsection for the discussion), so in fact we implemented the definition of Cheung and Ng at this step.
Eventually, ε_t and ξ_t are white noise processes with a null mean. Let U_t and V_t denote squares of standardised residuals:

$$U_t = \frac{(X_t - \mu_{x,t})^2}{h_{x,t}} = \varepsilon_t^2, \quad V_t = \frac{(Y_t - \mu_{y,t})^2}{h_{y,t}} = \xi_t^2. \tag{5}$$

The squared innovations are unobservable, but can be estimated consistently using squared residuals standardised by their conditional variances (Hong, 2001).
Let $r_{U,V}(k)$ denote the cross-correlation between U and V, for the k-th lag:

$$r_{U,V}(k) = \frac{c_{U,V}(k)}{\sqrt{(c_{U,U}(0)c_{V,V}(0))}}, \tag{6}$$

where $c_{U,V}(k)$ denotes the covariance between U and V at lag k. Since the processes U and V are independent:

$$\begin{matrix} \sqrt{T}r_{U,V}(k) \\ \sqrt{T}r_{U,V}(j) \end{matrix} \to N\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix}\right),\ k \neq j. \tag{7}$$

Cheung and Ng proposed the following test to verify the causality in the variance. First, we construct the following statistics:

$$S = T\sum_{i=j}^{k} \hat{r}^2_{U,V}(i), \tag{8}$$

which is distributed according to χ^2 distribution with $(k - j + 1)$ degrees of freedom. If the sample is small, the following, corrected version of the statistics is applied:

$$S = T\sum_{i=j}^{k} \varpi_i \hat{r}^2_{U,V}(i), \tag{9}$$

where $\varpi_i = \frac{T+2}{T-|i|}$. The statistics are used to test the null hypothesis of no causality in the variance.
Hong (2001) modified the test statistics (8), pointing out that larger weight should be attributed to more recent correlations. The idea of the test is as follows: volatility tends to cluster, which means that high volatility at the day t tends to be followed by another high volatility at the day $t+1, t+2, \ldots, t+n$, and low volatility tends to be followed by low volatility. Usually also high volatility at the day $t+1$ has higher impact on the volatility at the day t than volatility at the day $t+n$, where

n is a distant day in the past. On the one hand, empirical studies suggest that the impact of very distant volatility is rather minor. Thus, the Cheung and Ng (1996) test may be inefficient if a big M is used, because of the equal weighting of even distant correlations. On the other hand, some financial time series indeed exhibit strong cross-correlations and in such a case tests based upon a small number of past correlations may fail to detect causality. Thus, it is advisable to let M grow with T or include all $T-1$ cross-correlations with properly diminishing weights (Hong 2001). The Hong statistics is of the following form:

$$Q = \frac{S_H(k) - C_{1T}(k)}{D_{1T}(k)} \sim N\,(0,1), \tag{10}$$

where:

$$\begin{aligned} S_H(k) &= T\sum_{i=1}^{T-1} k^2\left(\frac{i}{M+1}\right)\hat{r}^2_{U,V}(i), \\ C_{1,T}(k) &= \sum_{i=1}^{T-1}\left(1-\frac{i}{M+1}\right)k^2\left(\frac{i}{M+1}\right), \\ D_{1,T}(k) &= \sum_{i=1}^{T-1}\left(1-\frac{i}{M+1}\right)\left(1-\frac{i+1}{T}\right)k^4\left(\frac{i}{M+1}\right). \end{aligned} \tag{11}$$

The variables $C_{1,T}$ and $D_{1,T}$ are approximately mean and variance of S_H, M is a positive integer, denoting lag truncation order, while $k(z)$ is the weighting function. Following Osińska (2011) and Łęt (2012), we use slightly modified version of the statistics, putting in the denominator $(M+1)$ instead of M, in order to have non-zero weight for the M-th correlation.

Let us note that S_H in (11) is a special case of (9) where $\omega_i = k^2(\frac{i}{M+1})$. Hong (2001) proposes several weighting functions $k(\cdot)$: the truncated, Bartlett, Parzen and Tuckey-Hanning, which attribute 0 weights for lags greater than M, as well as Daniell and quadratic-spectrall (QS) that have so called unbounded support. Below, we present formulas of the Daniell, Parzen and Tuckey-Hanning kernels (in the research we used all the kernels proposed by Hong, but for the sake of consistency we present only the results obtained for the three mentioned ones):

- Daniell kernel: $k(z) = \frac{\sin \pi z}{\pi z}$,
- Parzen kernel: $k(z) = \begin{cases} 1-6z^2+6|z|^3, & |z| \le 0.5, \\ 2(1-|z|)^3, & 0.5 < |z| \le 1, \\ 0, & \text{otherwise.} \end{cases}$,
- Tuckey-Hanning kernel $k(z) = \begin{cases} 0.5\cdot(1+\cos(\pi z)), & |z| \le 1, \\ 0, & \text{otherwise.} \end{cases}$.

The Q statistics is normally distributed and it should be compared to the upper-tailed critical value of $N(0,1)$. If Q is larger than the critical value, the null hypothesis of non-causality should be rejected.

4 Results

In the following section we present the estimates of the GARCH models and the results of the test of non-causality. We also clarify some controversies which may arise due to the choice of the volatility models.

4.1 Volatility models

In Tables 2-3 we present the estimates of the volatility models. As already mentioned, in all of the cases the best model to describe the dynamics of the series was the IGARCH one. We would like to drive the Readers' attention to the two facts. First of all, IGARCH processes are strictly stationary but not covariance stationary. Secondly, when we look at the estimated number of degrees of freedom of Student distribution, we note that only in the case of the Czech sCDS and the Hungarian bonds it exceeded 4. This means that the unconditional kurtosis does not exist in all the remaining cases. Taking all above into account, we must conclude that the estimates of the standard deviation, skewness and kurtosis presented in Table 1 should be taken with caution and treated as approximations in the sample, rather than the estimates of the population parameters. Another doubt may arise when we consider the way in which

Table 2: Estimates of the IGARCH models for the changes of sCDS

	Poland			the Czech Republic			Hungary		
Parameter	Coeff.	Std.Err.	p-value	Coeff.	Std.Err.	p-value	Coeff.	Std.Err.	p-value
μ	-0.42	0.43	0.33	-0.43	0.22	0.05	-0.67	1.31	0.61
a_1	-0.07	0.07	0.32	–	–	–	–	–	–
ω	6.01	4.16	–	0.97	0.89	–	190.57	116.46	–
ARCH (α_1)	0.28	0.08	0.00	0.30	0.11	0.01	0.49	0.20	0.01
Degr. of fdm (κ)	3.54	0.47	–	4.26	0.79	–	3.10	0.40	–
GARCH(β_1)	0.72	–	–	0.70	–	–	0.51	–	–

Note: we do not provide the p-value for ω and κ, as the test is two-sided. Instead, we compare the magnitude of the error term with the estimated parameter. All the estimated models are of the following form:

$$\begin{aligned} (r_t - \mu) &= a_1 (r_{t-1} - \mu) + y_t, \\ y_t &= \sigma_t \epsilon_t; \quad \epsilon_t \sim iid\ t(\kappa), \\ \sigma_t &= \omega + \ \alpha_1 y_{t-1}^2 + \beta_1 \sigma_{t-1}^2; \quad \beta_1 = 1 - \alpha_1. \end{aligned}$$

Table 3: Estimates of the IGARCH models for the changes of the government bonds

	Poland			the Czech Republic			Hungary		
Parameter	Coeff.	Std.Err.	p-value	Coeff.	Std.Err.	p-value	Coeff.	Std.Err.	p-value
μ	−0.002	0.001	0.058	−0.004	0.002	0.042	-0.004	0.002	0.031
a_1	−0.018	0.061	0.767	–	–	–	−0.106	0.075	0.159
ω	0.277	0.163	–	0.508	0.648	–	1.100	0.689	–
ARCH (α_1)	0.179	0.038	0.000	0.110	0.069	0.113	0.207	0.060	0.001
Degr. of fdm (κ)	3.119	0.457	–	2.974	0.435	–	4.025	0.911	–
GARCH(β_1)	0.821	–	–	0.890	–	–	0.793	–	–

Note: we do not provide the p-value for ω and κ, as the test is two-sided. Instead, we compare the magnitude of the error term with the estimated parameter. All the estimated models are of the following form:

$$\begin{aligned}(r_t-\mu) &= a_1\left(r_{t-1}-\mu\right)+y_t,\\ y_t &= \sigma_t\epsilon_t;\ \ \epsilon_t\sim iid\ t\left(\kappa\right),\\ \sigma_t &= \omega+\ \alpha_1 y_{t-1}^2+\beta_1\sigma_{t-1}^2;\ \ \beta_1=1-\alpha_1.\end{aligned}$$

we model the conditional mean. As Hong (2001) states: the existence of causality in variance do not exclude causality in mean. Investigating the causality in mean could be achieved e.g. through modelling the mean of the series with VAR process, instead of the AR-type one. However, the lack of autocorrelation in most of the series (see: Table 2-3) would result in insignificant estimates of the parameters and over-specification of the model. The latter could affect the results of the test and also introduce some noise into the second step of the model specification (the conditional variance). Therefore, we decided to use AR-type model for the conditional mean specification and limit our study to the investigation of interrelationships in volatility only.

4.2 Causality patterns

In Tables 4-6 we present the results of the non-causality test of Hong. We consider two cases, first one that includes instantaneous causality (feedback) and second one that excludes it. The argument for including instantaneous causality is twofold: on the one hand, in the financial markets investors react very fast and the response of the markets to the inflow of news can be observed even within the same day. Moreover, two markets (in our case: bond and sCDS) can react in the same time to the changes on third market or changes in the so-called global factor. On the other hand, when we take into account the feedback relationships, if we reject the null hypothesis in the case of causality in both directions, we are unable to distinguish the leadership of any market. Therefore, we decided to investigate two kinds of relationship separately.
Table 4 presents the results of Hong test for the Czech Republic. The cases where the null hypothesis of no causality is rejected are bolded. We observe that the

Table 4: Results of Hong test of non-causality – the case of the Czech Republic

Causality direction: sCDS → sovereign bonds

Kernel type	Including instantaneous causality (feedback)						Excluding instantaneous causality					
	Lag number (M)						Lag number (M)					
	1	5	10	20	30	50	1	5	10	20	30	50
truncated (p-value)	1.480 (0.07)	**2.801** (0.00)	**3.299** (0.00)	**2.250** (0.01)	**1.587** (0.06)	**2.276** (0.00)	1.162 (0.12)	**2.652** (0.00)	**3.166** (0.00)	**2.097** (0.02)	1.441 (0.06)	**2.165** (0.00)
Daniell (p-value)	1.347 (0.09)	**2.480** (0.01)	**2.971** (0.00)	**3.460** (0.00)	**3.144** (0.00)	**2.442** (0.00)	1.283 (0.10)	**2.442** (0.01)	**2.885** (0.00)	**3.352** (0.00)	**3.003** (0.00)	**2.284** (0.00)
Parzen (p-value)	1.002 (0.16)	**1.956** (0.03)	**2.518** 0.01)	**3.127** (0.00)	**3.338** (0.00)	**2.979** (0.00)	1.162 (0.12)	**1.962** (0.03)	**2.473** (0.01)	**3.047** (0.00)	**3.233** (0.00)	**2.830** (0.00)

Causality direction: sovereign bonds → sCDS

Kernel type	Including instantaneous causality (feedback)						Excluding instantaneous causality					
	Lag number (M)						Lag number (M)					
	1	5	10	20	30	50	1	5	10	20	30	50
truncated (p-value)	0.432 (0.33)	1.660 (0.05)	0.319 (0.38)	−0.266 (0.61)	−0.351 (0.64)	**−0.655** (<0.01)	−0.321 (0.63)	1.401 (0.08)	0.038 (0.49)	−0.483 (0.69)	−0.529 (0.70)	**−0.796** (<0.01)
Daniell (p-value)	0.872 (0.19)	1.265 (0.10)	1.356 (0.09)	0.722 (0.24)	0.256 (0.40)	**−0.138** (<0.01)	0.024 (0.49)	0.872 (0.19)	1.031 (0.15)	0.409 (0.34)	−0.029 (0.51)	**−0.373** (<0.01)
Parzen (p-value)	0.910 (0.18)	0.792 (0.21)	1.367 (0.09)	1.235 (0.11)	0.821 (0.21)	**0.205** (<0.01)	−0.321 (0.63)	0.106 (0.46)	1.006 (0.16)	0.902 (0.18)	0.492 (0.31)	**−0.092** (<0.01)

Note: The bolded values are these, for which null hypothesis of no causality is rejected.

Table 5: Results of Hong test of non-causality – the case of Hungary

Causality direction: sCDS → sovereign bonds

Kernel type	Including instantaneous causality (feedback)						Excluding instantaneous causality					
	Lag number (M)						Lag number (M)					
	1	5	10	20	30	50	1	5	10	20	30	50
truncated (p-value)	**105.04** (0.00)	**60.26** (0.00)	**44.13** (0.00)	**32.78** (0.00)	**28.35** (0.00)	**22.23** (0.00)	**44.42** (0.00)	**19.37** (0.00)	**13.24** (0.00)	**10.14** (0.00)	**9.59** (0.00)	**7.45** (0.00)
Daniell (p-value)	**113.07** (0.00)	**91.35** (0.00)	**72.01** (0.00)	**54.06** (0.00)	**45.41** (0.00)	**36.75** (0.00)	**44.16** (0.00)	**32.89** (0.00)	**24.02** (0.00)	**17.11** (0.00)	**14.31** (0.00)	**12.03** (0.00)
Parzen (p-value)	**106.67** (0.00)	**106.63** (0.00)	**89.44** (0.00)	**69.01** (0.00)	**58.10** (0.00)	**46.56** (0.00)	**44.42** (0.00)	**40.73** (0.00)	**31.72** (0.00)	**22.62** (0.00)	**18.44** (0.00)	**14.65** (0.00)

Causality direction: sovereign bonds → sCDS

Kernel type	Including instantaneous causality (feedback)						Excluding instantaneous causality					
	Lag number (M)						Lag number (M)					
	1	5	10	20	30	50	1	5	10	20	30	50
truncated (p-value)	**77.44** (0.00)	**44.54** (0.00)	**34.59** (0.00)	**24.50** (0.00)	**21.32** (0.00)	**17.42** (0.00)	**5.38** (0.00)	**2.14** (0.02)	**3.23** (0.00)	**1.648** (0.05)	**2.449** (0.01)	**2.584** (0.01)
Daniell (p-value)	**98.43** (0.00)	**69.12** (0.00)	**53.80** (0.00)	**40.94** (0.00)	**34.29** (0.00)	**27.67** (0.00)	**5.41** (0.00)	**4.17** (0.00)	**3.13** (0.00)	**3.02** (0.00)	**2.64** (0.00)	**2.67** (0.00)
Parzen (p-value)	**104.23** (0.00)	**84.17** (0.00)	**67.56** (0.00)	**51.84** (0.00)	**43.91** (0.00)	**35.12** (0.00)	**5.38** (0.00)	**4.905** (0.00)	**3.834** (0.00)	**3.147** (0.00)	**2.988** (0.00)	**2.60** (0.01)

Note: The bolded values are these, for which null hypothesis of no causality is rejected.

Table 6: Results of Hong test of non-causality – the case of Poland

Causality direction: sCDS → sovereign bonds

Kernel type	Including instantaneous causality (feedback)						Excluding instantaneous causality					
	Lag number (M)						Lag number (M)					
	1	5	10	20	30	50	1	5	10	20	30	50
truncated (p-value)	**47.6** (0.01)	**26.68** (0.01)	**20.08** (0.01)	**13.32** (0.01)	**10.3** (0.01)	**7.45** (0.01)	−0.69 (0.756)	−1.28 (0.9)	−0.56 (0.711)	−1.69 (0.954)	−2.1 (0.98)	−2.29 (0.99)
Daniell (p-value)	**62.77** (0.01)	**42.21** (0.01)	**32.26** (0.01)	**23.9** (0.01)	**19.37** (0.01)	**14.74** (0.01)	−0.79 (0.785)	−1.15 (0.875)	−1.33 (0.908)	−1.13 (0.871)	−1.5 (0.933)	−1.73 (0.001)
Parzen (p-value)	**67.93** (0.01)	**52.47** (0.01)	**42.28** (0.01)	**31.06** (0.01)	**25.88** (0.01)	**19.92** (0.01)	−0.69 (0.756)	−0.94 (0.826)	−1.22 (0.888)	−1.17 (0.879)	−1.14 (0.87)	−1.51 (0.935)

Causality direction: sovereign bonds → sCDS

Kernel type	Including instantaneous causality (feedback)						Excluding instantaneous causality					
	Lag number (M)						Lag number (M)					
	1	5	10	20	30	50	1	5	10	20	30	50
truncated (p-value)	**56.76** (0.01)	**32.05** (0.01)	**25.01** (0.01)	**17.68** (0.01)	**14.52** (0.01)	**10.88** (0.01)	**12.15** (0.01)	**4.59** (0.01)	**4.6** (0.01)	**2.76** (0.00)	**2.18** (0.02)	**1.17** (0.00)
Daniell (p-value)	**67.57** (0.01)	**49.89** (0.01)	**38.79** (0.01)	**29.43** (0.01)	**24.55** (0.01)	**19.39** (0.01)	**11.94** (0.01)	**8.76** (0.01)	**6.17** (0.01)	**4.82** (0.01)	**3.94** (0.01)	**3.07** (0.01)
Parzen (p-value)	**68.73** (0.01)	**60** (0.01)	**48.76** (0.01)	**37.38** (0.01)	**31.63** (0.01)	**25.15** (0.01)	**12.15** (0.01)	**11.08** (0.01)	**8.32** (0.01)	**5.99** (0.01)	**5.11** (0.01)	**3.98** (0.01)

Note: The bolded values are these, for which null hypothesis of no causality is rejected.

changes of sCDS volatility Granger-cause volatility of sovereign bonds, even if the lag 0 correlation is excluded from the test statistics. When we investigate possible causality from sovereign bonds to sCDS, the null hypothesis is not rejected in any case.

In table 5 we present the results obtained for Hungary. The null hypothesis of no causality is rejected in both direction and in the case of all lags. Thus, the volatility of sovereign bonds and sCDS are strongly linked one to another. We can suppose that during the crisis, risk in both markets was driven at the same time by the same set of external factors.

Yet another results were obtained for Poland (table 6). If we take into account the zero-lag correlation, we reject the null hypothesis in the case of causality from sCDS to the bond market. If the correlation is excluded, the hypothesis is not rejected. Thus, we suppose that the immediate relationships are the strongest between these two markets. However, if we investigate causality in the opposite direction, we conclude that volatility of sovereign bonds Granger cause volatility of sCDS, even if zero-lag causality is neglected.

5 Discussion

In order to interpret the obtained results we analyzed the sovereign debt of the three economies, the debt policy applied in each country, size of the sovereign bond markets, and their integration with the EU markets. Let us make the comments on each of them.

5.1 External and internal debt

Yarashevich (2013) points out that there are substantial variations among the V3 group in terms of their debt-to-output ratio. In the analyzed period, the least indebted country in the region was the Czech Republic (Sobjak 2013), where debt-to-GDP ratio amounted to 28.7% in 2008 and has grown to 45.5% in 2012. Yarashevich (2013) notes that when it comes to the external debt, the Czech Republic was the only one post-communist country whose debt-to-output ratio was less than 50% in 2010.

Hungary was in the worst situation at the outbreak of the crisis, as its economic problems had caused a slow-down already in 2006 (Sobjak 2013). Moreover, the country had been under the excessive debt procedure since 2004, and since 2009 its debt-do-GDP ratio approached 80%. When it comes to the external debt, in 2010 the debt-to-output ratio exceeded 100% (Yarashevich 2013).

The amount of gross debt to GDP ratio in 2008 in Poland amounted to 47.1% and was steadily increasing to reach 57% in 2013. The ratio of the external debt to GDP amounted to 56.3% (2010) and it was the second best-prospering economy in the V3 group, after the Czech Republic (Yarashevich 2013).

5.2 Public debt policy before and during the crisis

Novak (2015) presents the review of literature dealing with the public debt policy in CEE countries. Following Nemeth (2015) and Alesina-Tabellini (1990) she points out that in the three countries in the pre-crisis period the increase of the debt-to-GDP ratio was observed during election years. During the financial crisis, external factors dominated the creation of the government debt (Novak 2015). Sobjak (2013) points out similarities of the V3 group: all are export-oriented, open and exposed to shift in the global demand – particularly to the European one. Next, in all three countries, together with the gradual integration of the domestic banking system with the European one, an increase of the number of the local branches of the western banks have been observed. Eventually, they wounded up with a large share of the banking assets, and lots of loans for private and institutional clients had been granted in foreign currencies. All these factors can explain the fact that we observe at least immediate causality from sCDS (reacting mostly to changes in external situation) to domestic bonds in each country.

Despite those similarities, each country implemented a slightly different strategy of dealing with the crisis. Hungary was in the worst situation with the slowdown in 2006. In 2008 it applied and received a loan from IMF. The loan, however, did not help to improve the situation and in 2011 the country was severely hit by the next phase of the crisis, and its credit rating has been cut off by three rating agencies to the junk level. In 2010, as a consequence of the elections, a new party came to power and implemented a set of internal reforms to overcome the crisis. First, Hungary had to maintain budgetary discipline through procyclical behavior and it entered the excessive deficit procedure – see: Novak (2015) (only in 2014 – thus, in the period which is out of scope of our analysis – the Central Bank of Hungary had to introduce a credit easing program in order to facilitate the recovery from recession). Among others, a flat personal income tax, a banking tax and temporal sectoral levies had been introduced. Next, VAT has been increased from 25% to 27% (Sobjak 2013). When it comes to foreign loans, in September 2011 Hungarian government passed a legislation that unilaterally changed the terms and conditions of all currency loans contracts. The cost of the transaction had to be borne entirely by banks. In mid-December 2011 the government and banks agreed to share costs of further arrangements.

On contrary, the policy implemented in the Czech Republic, which was in much better economic condition, was characterized by moderate budgetary spending - due to its mostly negative real GDP growth in the last couple of years. As already stated, the economic growth in the Czech Republic has been mostly dependent on the external demand. Before the crisis export made up even to 80-90% of the Czech GDP (Sobjak 2013). Thus, it is not a surprise, that the country was very vulnerable to the external shocks (and thus we observed causality from sCDS reacting to the external events to the government bonds). However, as noted by Sobjak (2013) the slow pace of recovery of the Czech Republic is also a consequence of its government policy. In 2011 fiscal consolidation measures consisted mostly of restructuring expenditure. Public sector wages were cut and social benefits were eliminated. Also, when it comes to the tax system, tax exceptions and allowances were eliminated. The economy in the analyzed period was characterized by stagnation. The reforms have been, however, much less drastic than in the case of Hungary and this can explain the results obtained in our research, that the main source of the risk to the solvency of the country in the analyzed period, were the external events (volatility transmission from sCDS to government bonds).

Eventually, in Poland GDP was growing through the whole analyzed period. This can be contributed to the fact that the Polish economy is less open than the other CEE ones: export represented only a third of Polish GDP, compared to 60% in Hungary and the Czech Republic in 2013, and 80-90% before the crisis (Sobjak 2013). However, the growth had been much slower compared to the pre-crisis period. The recession in the Euro area economies affected Poland mostly in regard to the decrease in demand for Polish products (Reichardt 2011). As a response to the crisis, Poland released the "Stability and Development Plan" (November 2008) which was aimed to maintain stabilization in public finances and financial system. The actions taken by the government included, however, rising the taxes in 2009 (VAT) and freezing employments in the banking sector. At the same time the government admitted to rely on EU funding to supplement the budget cuts. Borowski (2014) analyzed the structure of public debt in Poland over the period 2005 – 2013 and concluded that the domestic market is the main source of the government funding needs, as its participants constitute the majority of the government bonds' holders. We can thus observe that the risk to the government solvency could have two sources: the internal and the external one. However, as the funding of government needs was mostly domestic, this can justify that the longer-term causality occurred from the government bonds (reacting mostly to internal factors) to sCDS.

5.3 Integration with European markets

The last factor, that could explain the result is the degree of integration of domestic markets with the EU market. Orlowski and Tsibulina (2014) examine (inter alia) integration of government markets of V3 countries with the EU one. They show that the crisis episode disturbed the integration process. However, according to their results, the Czech bond market is now fully aligned with the German one, while the Polish one is now on the convergence path. The Hungarian bonds are characterized by the highest risk premium and the convergence has been destabilized.

This phenomenon also can explain the results obtained in our research. As the Czech bond market is now fully aligned with the German market, and as the sCDS premia are documented to react mostly to external (pan-European) shocks, it is understandable that the external shocks will be first visible in the more liquid sCDS prices, and then transmitted to the less liquid bond market. The Polish bond market is less integrated with the EU, but is on the convergence path. Therefore both markets can react to the same external shocks and we observe feedback in causality. However, as the bond market in Poland is the largest and most liquid one (see e.g. Associacion for Financial Markets in Europe, 2016), it can affect the sCDS dynamics as well. Finally, the Hungarian bonds are not so closely related to the German ones. The condition of fundamentals in the country were the weakest, while the external debt was the largest in the analyzed group.

On the one hand the investors in the sCDS market should carefully observe the condition of fundamentals, as their change can possibly affect the sCDS price. It might react to the changes of the yield dynamics, which is closely related to internal situation. On the other hand, the change in external situation (e.g. global or pan-European shocks) can also affect the bond prices – thus we should also observe causality form sCDS (which prices are vulnerable to the changes in the external situation) to the bond prices.

6 Summary

In the paper we analyze the causality patterns between sovereign CDS and bond yields in three CEE countries, the Czech Republic, Hungary and Poland, in the hectic crisis period from 2008 to 2013. We particularly focus on the analysis of the causality in variance using Hong (2001) test. Volatility of the financial series reflects the inflow of information: the more information flows to the market, the higher volatility is. As both instruments, sCDS and bonds, reflect the credibility of the country, the occurrence of the dependencies between these series and causality in particular, are economically justified. We find that in general sCDS is a Ganger cause to bond yields in all three countries, whereas bond yield is a Granger cause for sCDS in Hungary and in Poland.

From the methodological point of view, neither the application of different kernels in Hong test, nor including or excluding instantaneous causality change the results across our sample, with the exception of Poland in the former case when the CDS-bond relation is taken into account.

When interpreting the results, we need to take into account two facts. First of all, bond yields are documented to be related to the fundamentals of the economy more than the spreads of sCDS (see e.g Kocsis, 2014). The latter are much more prone to international turmoil (see e.g. Longstaff *et al.* 2011, Kliber, 2011; Kliber 2014).

The results of the research can be explained based on the degree of openness of the economies, integration with European market and internal policies of dealing with public debt. In Poland, which is not that much export-oriented as the remaining two economies, we observe limited impact of external events to the volatility of government bonds (only instantaneous causality from the sCDS to bonds). On contrary, in the much more export-oriented Czech Republic and Hungary we can observe causality from sCDS to bond market even when the instantaneous causality is excluded from the analysis. In the case of the Czech Republic, where the fundamentals have been solid, the bond market has been fully integrated with the German one, and the risk of internal crisis was very limited, we do not observe causality in variance from government bonds to sCDS market. On contrary, in the case of unstable Hungary, causality is observed in both directions, as the internal instability also can affect the sCDS prices.

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References

[1] Adam M. (2013), Spillover and Contagion in the Sovereign CDS Market, *Bank & Credit* 44(6): 571–604.

[2] Alesina A., Tabellini B. (1990), A positive theory of fiscal deficits and government debt. *Review of Economic Studies* 57(3): 403–414.

[3] Associacion for Financial Markets in Europe (2016), Government Bond Data Report
`https://www.afme.eu/globalassets/downloads/data/government-bonds/afme-pdb-stn-government-bond-data-report-3q16-v3.pdf`

[4] Będowska-Sójka B., Kliber A. (2013), Economic Situation of the Country or Risk in the World Financial Market? The Dynamics of Polish Sovereign Credit Default Swap Spreads, *Dynamic Econometric Models* 13: 87-106 `http://dx.doi.org/10.12775/DEM.2013.005`

[5] Bollerslev T. (1986), Generalized Autoregressive Conditional Heteroskedasticity. *Journal of Econometrics* 31: 307–327.

[6] Borowski K. (2014), Public Debt Management as Valid State Policy Instrument (in Polish). *Journal of Finance and Financial Law* 3: 21–36.

[7] Camba-Mendez G., Serwa D. (2016), Market Perception of Sovereign Credit Risk in the Euro Area during the Financial Crisis, *The North-American Journal of Economics and Finance*, 37: 168–189.

[8] Camba-Mendez G., Kostrzewa K., Marszal A., Serwa D. (2016), Pricing Sovereign Credit Risk of Poland: Evidence from the CDS Market, *Emerging Markets Finance and Trade*, 52(12): 2687–2705, DOI: 10.1080/1540496X.2016.1216935.

[9] Cheung Y-W., Ng L.K. (1996), A Causality-in-Variance Test and its Application to Financial Market Prices. *Journal of Econometrics* 72: 33–48.

[10] Coudert V., Gex M. (2010), Credit Default Swap and Bond Markets: which Leads the Other?, *Financial Stability Review*, Banque de France 14: 161–167.

[11] Fontana A., Scheicher M. (2016), An Analysis of Euro-Area Sovereign CDS and their Relation with Government Bonds, *Journal of Banking and Finance*, 62: 126–140.

[12] Giannikos C., Guirgis H., Suenn M. (2013), The 2008 Financial Crisis and the Dynamics of Price Discovery Among Stock Prices, CDS Spreads and Bond Spreads for U.S. financial firms, *Journal of Derivatives* 21: 27–48.

[13] Hong Y. (2001), A Test for Volatility Spillover with Application to Exchange Rates. *Journal of Econometrics*, 103: 183–224.

[14] Jain S. (2014), The Relationship between Interest Rates and Credit Spreads, Market Realist `https://marketrealist.com/2014/03/relationship-interest-rates-credit-spreads/`

[15] Kliber A. (2011), Sovereign CDS in Central Europe – Linkages and Independence, *Dynamic Econometric Models*, 11: 111–128.

[16] Kliber A. (2013), Influence of the Greek Crisis on the Risk Perception of European Econo-mies, *Central European Journal of Economic Modelling and Econometrics*, 5(2): 125–161.

[17] Kliber A. (2014), The Dynamics of Sovereign Credit Default Swaps and the Evolution of the Financial Crisis in Selected Central European Economies, *Czech Journal of Economics and Finance* 64(4): 330–350.

[18] Kocsis Z. (2014), Global, Regional and Country-Specific Components of Financial Market Indicators: an Extraction Method and Applications. *Acta Oeconomica* 64 (S1): 81–110.

[19] Kwiatkowski D., Phillips P. C. B., Schmidt P., and Shin Y. (1992), Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root. *Journal of Econometrics* 54, 159–178.

[20] Laurent S. (2010), *G@rch 6.0 help.* Timberlake Consultants Ltd., London (2010).

[21] Łęt B. (2012), The Granger Causality Analysis of Crude Oil Future Price and U.S. Dollar Value (in Polish). *AUNC Ekonomia XLIII* (2): 221–231.

[22] Longstaff F. A., Mithal S., Neis E. (2005), Corporate Yield Spreads: Default Risk or Liquidity? New Evidence from the Credit Default Swap Market, *The Journal of Finance* 60: 2213–2253, DOI: http://dx.doi.org/10.1111/j.1540-6261.2005.00797.x.

[23] Longstaff F. A., Pan J., Pedersen L. H., Singleton K. J. (2011), How Sovereign is Sovereign Credit Risk? *American Economic Journal: Macroeconomics, American Economic Association* 3(2): 75–103. DOI: http://dx.doi.org/10.1257/mac.3.2.75

[24] Merton R.C. (1974), On the Pricing of Corporate Debt: The Risk Structure of Interest Rates, *Journal of Finance* 29(2): 449–470.

[25] Németh A. O. (2015), Political Business Cycles: Theory and Empirical Findings for the CEE Region. [in:] Deficit and debt in Transition. The Political Economy of Public Finances in Central and Eastern Europe, [in:] Benczes I. [ed.] Central European University Press: 35–58.

[26] Novák Z. (2015), Public Debt In The CEECs: Is The Sovereign Debt Crisis Over?. *DIEM : Dubrovnik International Economic Meeting* 2(1): 822-836. Retrieved from http://hrcak.srce.hr/161670

[27] Orlowski L., Tsibulina A. (2014), Integration of Central and Eastern Europaean and the Euro-Area Financial Markets: Repercussion from the Global Financial Crisis, *Comparative Economic Studies* 56(3): 376–395

[28] Osińska M. (2011), On the Interpretation of Causality in Granger's Sense. *Dynamic Econometric Models* 11: 129–139.

[29] Reichardt A. (2011), Poland and the Global Economic Crisis: Observations and Reflections in the Public Sector. *Journal of Finance and Management in Public Services* 10 (1): 38–48.

[30] Sobjak A. (2013), From the Periphery to the Core? Central Europe and the Economic Crisis. *The Polish Institute of International Affairs* 7(55).

[31] Ters K., Urban J. (2018), Intraday dynamics of credit risk contagion before and during the euro area sovereign debt crisis: Evidence from central Europe, *International Review of Economics and Finance* 54: 123–142.

[32] Yarashevich V. (2013), External Debt of Post-Communist Countries, *Communist and Post-Communist Studies* 46(2): 203–216.