Quantitative Finance

ISSN: 1469-7688 (Print) 1469-7696 (Online) Journal homepage: http://www.tandfonline.com/loi/rquf20

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To cite this article: Christoph Wegener, Tobias Basse, Frederik Kunze & Hans-Jörg von Mettenheim (2016): Oil prices and sovereign credit risk of oil producing countries: an empirical investigation, Quantitative Finance, DOI: 10.1080/14697688.2016.1211801

To link to this article: http://dx.doi.org/10.1080/14697688.2016.1211801

Published online: 08 Sep 2016.

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Quantitative Finance, 2016
http://dx.doi.org/10.1080/14697688.2016.1211801

Routledge
Taylor & Francis Group

Oil prices and sovereign credit risk of oil producing countries: an empirical investigation

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(*Received 3 February 2016; accepted 21 June 2016; published online 8 September 2016*)

The low oil price recently has caused fears about the sustainability of public finances in some oil producing countries. We examine the relationship between oil prices and sovereign credit risk examining the CDS market. Analysing data from nine countries (Brazil, Malaysia, Norway, Qatar, Russia, Saudi Arabia, the United Kingdom, the United States of America and Venezuela) we have estimated bivariate VAR-GARCH-in-mean models. The results of our empirical investigations generally speaking do suggest that positive oil price shocks lead to lower sovereign CDS spreads. Thus, our findings support the hypothesis that higher oil prices improve the fiscal stability of oil producing countries.

Keywords: Sovereign debt crisis; Oil prices; Fiscal stability

JEL Classification: G17, H63, C58, Q43

1. Introduction

The price of oil has fallen dramatically in recent times. Since 2014, the spot price for West Texas Intermediate (WTI) crude oil, for example, fell from around 100 USD per barrel to below 30 USD per barrel. There are a number of reasons for this slump of oil prices. Most importantly, there have been technological advances in the production of shale oil (see Borenstein and Kellogg 2014, Cornell 2015). The implications of unconventional oil resources for the market power of the OPEC countries already have been discussed in the literature for a while (most importantly, see Greene *et al*. 2006). However, the shale oil revolution in the US recently has limited the ability of OPEC countries to control oil prices in a surprisingly distinct way. Moreover, fears about economic problems in China have started to materialize (see Cesa-Bianchi and Stratford 2016, Tyers 2016). Chinese equity markets have reacted quite strongly to these worries. A slowdown to growth in the biggest economy of Asia clearly could have negative effects on the demand for oil pushing down oil prices (see Cesa-Bianchi and Stratford 2016). Additionally, some observers have been quite optimistic about a return of Iranian oil to the market. After finding a solution to control Iran's controversial nuclear programme the lifting of sanctions has opened the global oil market to supply from Teheran. Meanwhile, low oil prices even have caused more and more concerns about sovereign credit risk of oil producing countries. This quite special type of credit risk describes the phenomenon that governments are unable or unwilling to repay their debt (see, e.g. Ang and Longstaff 2013, Jeanneret 2015). We analyse the relationship between oil price changes and sovereign credit risk in oil producing countries testing for Granger causality in bivariate VAR-models considering GARCH effects. Our measure of sovereign credit risk is the price of sovereign credit default swap contracts (also called CDS spread). This is a very popular indicator trying to quantify the market perception of sovereign credit risk (see, e.g. Aizenman *et al*. 2013, Ang and Longstaff 2013). We examine the price of WTI oil.

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The paper is structured as follows. Section 2 presents some ideas with regard to sovereign credit risk and the fiscal solvency of countries. Section 3 introduces the data examined and presents the results from some preliminary statistical analysis. Section 4 discusses the relevant methodological issues. The findings of our empirical investigations are then presented and evaluated in Section 5. Section 6 concludes by analysing our results and by discussing some directions for further research.

2. Sovereign credit risk and fiscal solvency

History has shown that governments can default on the debt they have accumulated over time (see, e.g. Velde and Weir 1992, Dincecco 2009). Kotlikoff (2006) has argued convincingly that the proper way to discuss the sustainability of public finances is to examine the lifetime fiscal burdens facing current and future generations and that if these burdens exceed the resources of those generations the country's fiscal policy has to be regarded as unsustainable. This difficult situation could lead to a political environment where the government simply has to consider the option to not pay its creditors.† The relevant economic and legal questions surrounding sovereign debt default have been survey by Panizza *et al*. (2009). The recent discussion has focused quite strongly on the European government debt crisis (see, e.g. Basse *et al*. 2012, Sibbertsen *et al*. 2014). In this context, difficulties with high government expenditures and costly bank rescue programmes have been analysed intensively. Now, the falling oil price becomes an additional relevant fiscal problem.‡

With regard to oil-exporting countries El Anshasy and Bradley (2012), for example, have noted that some governments tend to rely heavily on oil revenues to finance government expenditure. Other oil producing nations have a stronger non-oil tax base. We examine data from both types of countries. In any case, changes to oil prices can affect public finances in a relevant way and therefore should be of some importance with regard to the fiscal solvency of these countries. In fact, El Anshasy and Bradley (2012) have presented empirical evidence indicating that in the long run higher oil prices induce larger government size. This should increase the need to fund public expenditure. Moreover, fluctuations to the oil price seem to lead to more uncertainty about future government revenues. Devlin and Titman (2004), for example, have noted that one main problem countries with a strong fiscal dependence on oil have to face is associated to the volatility of oil prices. Especially, in countries with a weak non-oil tax base, falling oil prices should have the ability to cause significant increases to sovereign credit risk. There are a number of relevant issues to be discussed in this context. Most importantly, the establishment of savings and stabilization funds has been suggested to increase the fiscal stability of oil-exporting countries because these funds could help to shield public finances from oil price volatility (see, e.g. Devlin and Titman 2004).

Additionally, exchange rates also might react to oil price swings adding to the fiscal problems. Beckmann *et al*. (2016), for example, recently have shown that currencies of oil importers and oil exporters display a different dependency structure against the US dollar. Basher *et al*. (2016) have argued that this is an extremely important topic because an increase to the oil price could cause an appreciation of the currencies of oil producing countries affecting their terms of trade and competitiveness. Vice versa falling oil prices might lead to a depreciation of the currencies of oil producing countries causing additional fiscal problems (especially with regard to debt issued in foreign currencies). Basher *et al*. (2016) have documented significant movements to exchange rates in oil exporting economies after oil demand shocks but only have reported quite limited evidence indicating that oil supply shocks affect the FX market. Moreover, there is clear empirical evidence that oil price uncertainty can hurt economic growth (see, e.g. Elder and Serletis 2011, Rahman and Serletis 2012). In fact, Elder and Serletis (2011) have examined data from the

†Political consideration, for example, seem to be very important trying to understand the default of Argentina.

‡The low oil price, for example, resulted in a fall of oil revenues in Saudi Arabia which led to a budget deficit and forced the kingdom to think about higher taxes and cuts to government spending. Other oil producing countries face even more serious fiscal problems. Press reports indicate that Venezuela at the moment seems to have not enough money to pay for its money produced by the bank note company De La Rue. The fiscal problems in Qatar even forced job cuts at the government funded media network Al Jazeera.

Table 1. p-values of the unit root test.

Variable	Short	Long
WTI	0.72	0.71
UK	0.66	0.62
US	0.67	0.82
Norway	0.56	0.47
Saudi Arabia	0.99	0.99
Brazil	0.91	0.94
Malaysia	0.48	0.56
Russia	0.30	0.30
Qatar	0.96	0.97
Venezuela	0.87	0.93

Table 2. Lag length selection.

Variable	BIC
UK	1
US	1
Norway	1
Saudi Arabia	1
Brazil	1
Malaysia	1
Russia	1
Qatar	1
Venezuela	1

Table 3. p-values of the ARCH-LM test.

Country	p-value
UK	0.001
US	0.001
Norway	0.001
Saudi Arabia	0.001
Brazil	0.001
Malaysia	0.001
Russia	0.001
Qatar	0.001
Venezuela	0.001

Table 4. Results VAR(1)-(GARCH(1,1)) estimation.

Country	$\hat{c}_1$	$\hat{\gamma}_{11}$	$\hat{\gamma}_{12}$	$\hat{\omega}_1$	$\hat{f}_1$	$\hat{g}_1$	$\hat{c}_2$	$\hat{\gamma}_{21}$	$\hat{\gamma}_{22}$	$\hat{\omega}_2$	$\hat{f}_1$	$\hat{g}_1$	$\hat{\lambda}_{11}$	$\hat{\lambda}_{12}$	$\hat{\lambda}_{21}$	$\hat{\lambda}_{22}$
UK	–0.079	*–0.055*	*0.070*	**0.008**	**0.037**	**0.962**	0.014	0.007	**0.093**	**0.009**	**0.152**	**0.847**	0.045	0.064	0.003	**–0.113**
US	**–0.142**	**–0.069**	–0.009	**0.007**	**0.035**	**0.964**	**0.303**	*–0.093*	**0.068**	**0.472**	**0.194**	**0.785**	0.080	0.036	–0.172	*–0.130*
Norway	–0.087	**–0.062**	0.006	**0.008**	**0.036**	**0.962**	0.005	–0.002	–0.007	**0.002**	**0.105**	**0.891**	0.081	0.036	0.001	–0.055
Saudi Arabia	–0.047	**–0.060**	0.019	–	–	–	0.086	**–0.113**	**–0.077**	–	–	–	–	–	–	–
Brazil	–0.045	–0.066	–0.002	**0.280**	**0.761**	0.003	**0.256**	**–0.132**	0.025	**0.227**	**0.119**	**0.878**	0.000	0.000	0.000	0.003
Malaysia	**–0.167**	**–0.064**	–0.005	**0.009**	**0.035**	**0.963**	**0.477**	**–0.237**	**0.071**	**0.181**	**0.102**	**0.882**	0.152	0.022	–0.200	**–0.262**
Russia	**–0.045**	**–0.070**	–0.005	**0.111**	**0.310**	0.001	0.088	**–1.126**	–0.052	**0.128**	**0.125**	**0.874**	0.000	0.000	0.000	–0.001
Qatar	–0.080	**–0.067**	0.013	**0.006**	**0.025**	**0.973**	**0.290**	**–0.067**	**0.081**	**0.139**	**0.205**	**0.791**	–0.018	**0.081**	**–0.390**	–0.032
Venezuela	–0.044	**–0.067**	0.000	–	–	–	5.330	**–9.394**	**0.106**	–	–	–	–	–	–	–

Note: Bold numbers indicate significant results on a 5% level of significance.

US and have argued convincingly that the extreme volatility of oil prices has contributed to the intensity of the decline in manufacturing activity in the years 2008 and 2009. This is of some importance examining government finances because lower economic growth should also hurt tax receipts and therefore certainly is an additional burden for fiscal stability.

3. Data and initial empirical results

We examine daily observations of CDS spreads of eight countries (United Kingdom (UK), United States (US), Norway, Saudi Arabia, Brazil, Malaysia, Russia, Qatar and Venezuela) and the oil price from 26 January 2011 to 26 January 2016 ($T = 1305$). These time series are obtained from Thomson/Reuters. CDS spreads are a very popular measure of sovereign credit risk (see Ang and Longstaff 2013, Pianeti and Giacometti 2015). Excluding the US all CDS contracts are denominated in USD. US CDS spreads are denominated in EUR.

First of all, we investigate the trending behaviour of our time series using the unit root test by Phillips and Perron (1988). The p-values of the test are reported in table 1. Here, *short* refers to the case where the truncation lag parameter is set to $p^* = 4 \times (T/100)^{0.25}$ and *long* to $p^* = 12 \times (T/100)^{0.25}$.

The results here indicate integrated time series of order one. Therefore, we use first differences of all variables to avoid spurious results and estimate bivariate Vector Autoregressive Models (VAR) for the CDS spreads and the oil price in first differences. In it's general form a VAR(p) model (see Sims 1980) is given by

$$y_t = C + \sum_{j=1}^{p} \Gamma' y_{t-j} + u_t. \tag{1}$$

Here, y_t is a vector of variables, C is a vector of constants, the matrix Γ contains autoregressive parameters up to lag p and u_t is an error term. Note that all variables in this setting are treated endogenously. After investigating the trending behaviour above, we should determine the appropriate lag length p of the model. We focus on the Bayesian Information Criterion (BIC). The results for the bivariate VAR are reported in table 2 .

The BIC indicates one lag for all cases. This selection might be provoked by daily data—another phenomenon which is often present in data with a high frequency is (Generalized) Autoregressive Conditional Heteroscedasticity ((G)ARCH) (see Engle 1982, Bollerslev 1986). Thus we use the multivariate ARCH-LM test (see Hamilton 1994) which is based on the regression

$$\operatorname{vec}\left(\hat{u}_t \hat{u}_t'\right) = \Omega + G_1 \operatorname{vec}\left(\hat{u}_{t-1}\hat{u}_{t-1}'\right) + \cdots + G_q \operatorname{vec}\left(\hat{u}_{t-q}\hat{u}_{t-q}'\right) + v_t \tag{2}$$

where vec is the column-stacking operator, $\left(\hat{u}_t \hat{u}_t'\right)$ are the estimates of $\left(u_t u_t'\right)$, Ω and G_i are coefficient matrices and v_t is an error process. The null hypothesis is defined as $H_0 : B_1 = B_2 = \cdots = B_q = 0$ against $H_1 : B_1 \neq 0 \cap B_2 \neq 0 \cap \cdots \cap B_q \neq 0$ and the test statistic is distributed as $\chi^2\left(qN^2(N+1)^2/4\right)$ with N variables. The p-values of the test are reported in table 3.

4. Modelling strategy

As noted above, there is empirical evidence for GARCH-effects in the residuals of the VAR models. Thus, we use a VAR(1)-GARCH(1,1)-in-mean model. This model allows conditional heteroscedastic errors and investigates the influence of the conditional standard deviations of the residuals on the conditional mean. In it's general form, the model can be written as

$$y_t = C + \Gamma' y_{t-1} + \Lambda' h_t + u_t \tag{3}$$

with y_t as a 2×1 vector of the first differences of the CDS spread of country i and the first differences of the oil price. Furthermore, $C = (c_1, c_2)$ is a constant term and Γ and Λ are 2×2 matrices. Here, Γ contains autoregressive parameters and Λ captures the effects of the conditional standard deviations $h_t = \left(\sqrt{h_{11,t}}, \sqrt{h_{22,t}}\right)$ on the conditional mean of y_t. $u_t = \left(e_{1,t}, e_{2,t}\right)$ is a residual vector. The multivariate GARCH(1,1) is based on the presentation in Engle and Kroner (1995)

$$H_t = \Omega + F \operatorname{vec}\left(u_{t-1}u_{t-1}'\right) + G H_{t-1} \tag{4}$$

where, in it's unrestricted form, Ω is 4×1, F and G are 4×4 matrices. H_t is the conditional variance with $h_t = H_t^{\frac{1}{2}}$. To ensure that H_t is positive definite, we extract the diagonal from the two matrices $\operatorname{vec}\left(u_t u_t'\right)$ and H_t such that

$$C = (c_1, c_2), \quad \Gamma = \begin{pmatrix} \gamma_{11} & \gamma_{12} \\ \gamma_{21} & \gamma_{22} \end{pmatrix}, \quad \Lambda = \begin{pmatrix} \lambda_{11} & \lambda_{12} \\ \lambda_{21} & \lambda_{22} \end{pmatrix},$$

$$\Omega = (\omega_1, \omega_2), \quad F = \begin{pmatrix} f_1 & 0 \\ 0 & f_2 \end{pmatrix}, \quad G = \begin{pmatrix} g_1 & 0 \\ 0 & g_2 \end{pmatrix}.$$

In order to estimate this model, we maximize the appropriate likelihood function numerically.

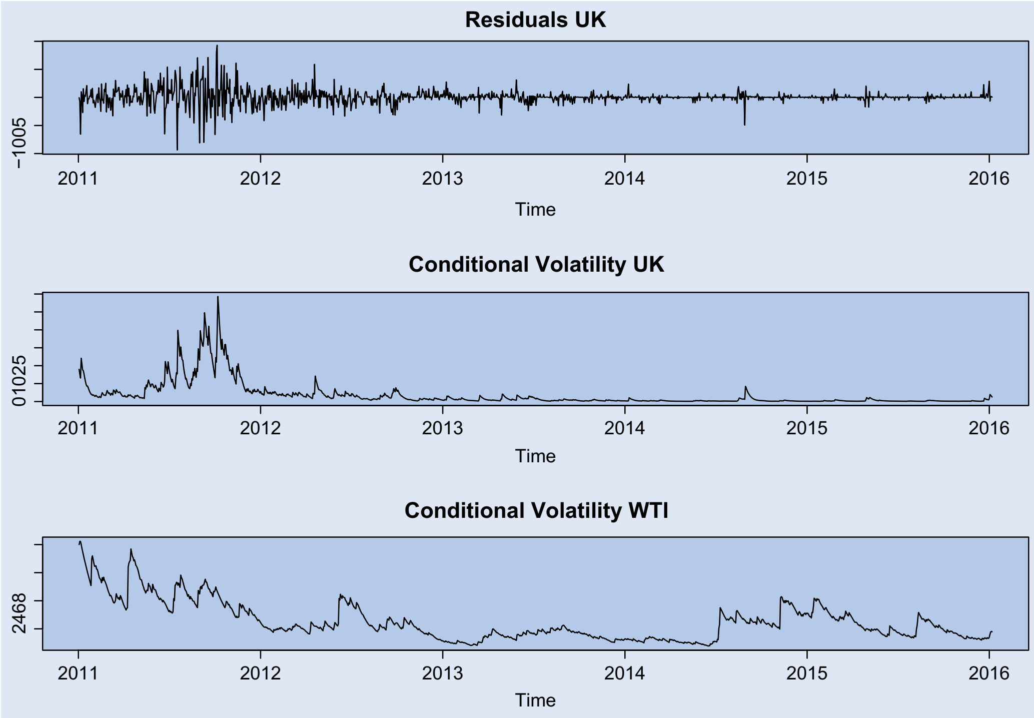

Figure 1. Estimated residuals and conditional volatilities of the UK model.

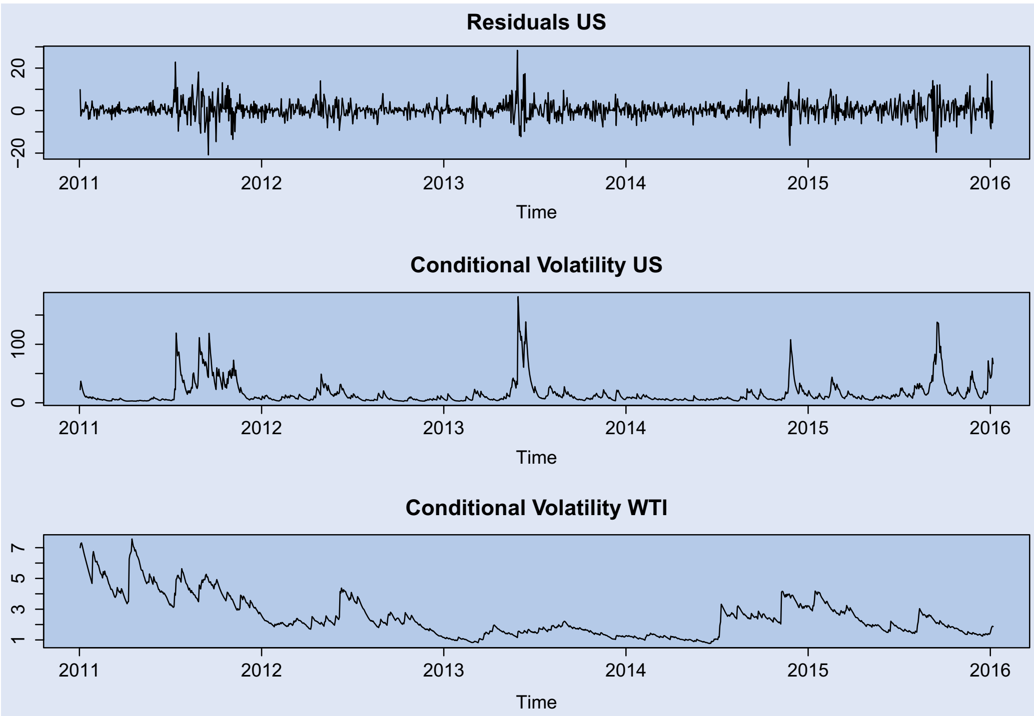

Figure 2. Estimated residuals and conditional volatilities of the US model.

5. Empirical results

Our empirical investigations have produced a number of interesting results reported in table 4. Most importantly, there is clear empirical evidence supporting the hypothesis that increasing oil prices lead to lower sovereign CDS spreads. This effect is captured by γ_{21}.† The UK is the only exception. For this quite diversified economy with a rather strong non-oil tax base a positive coefficient is observed. However, this reaction of the CDS spreads to a changing oil price is statistically

†Note that the hat indicates estimates of the coefficients of model (3) and model (4).

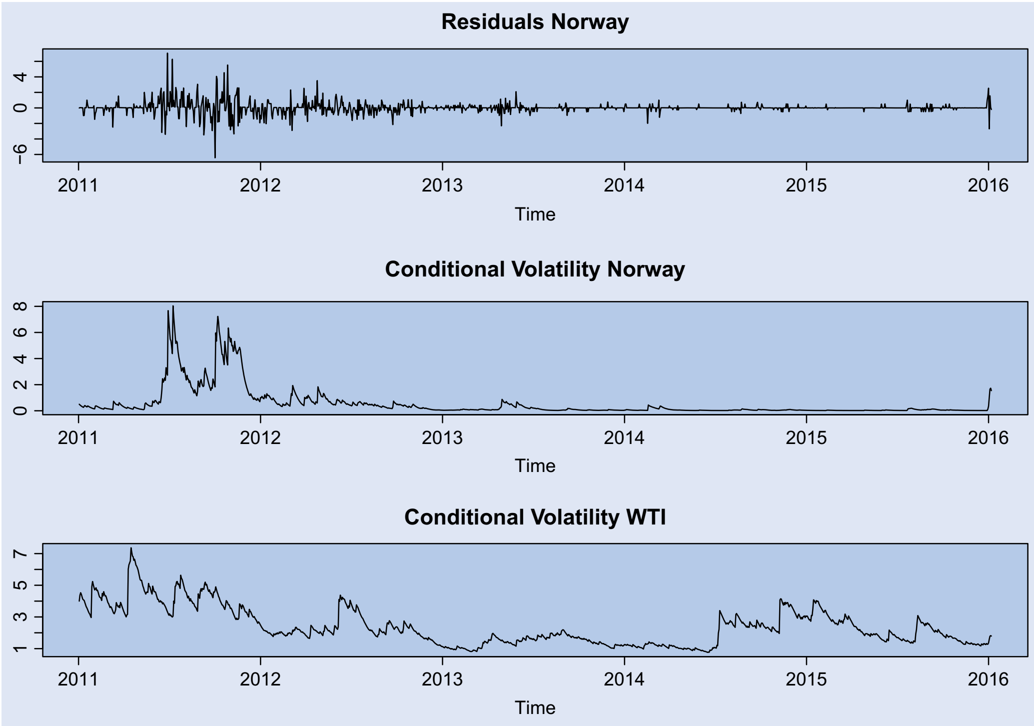

Figure 3. Estimated residuals and conditional volatilities of the Norway model.

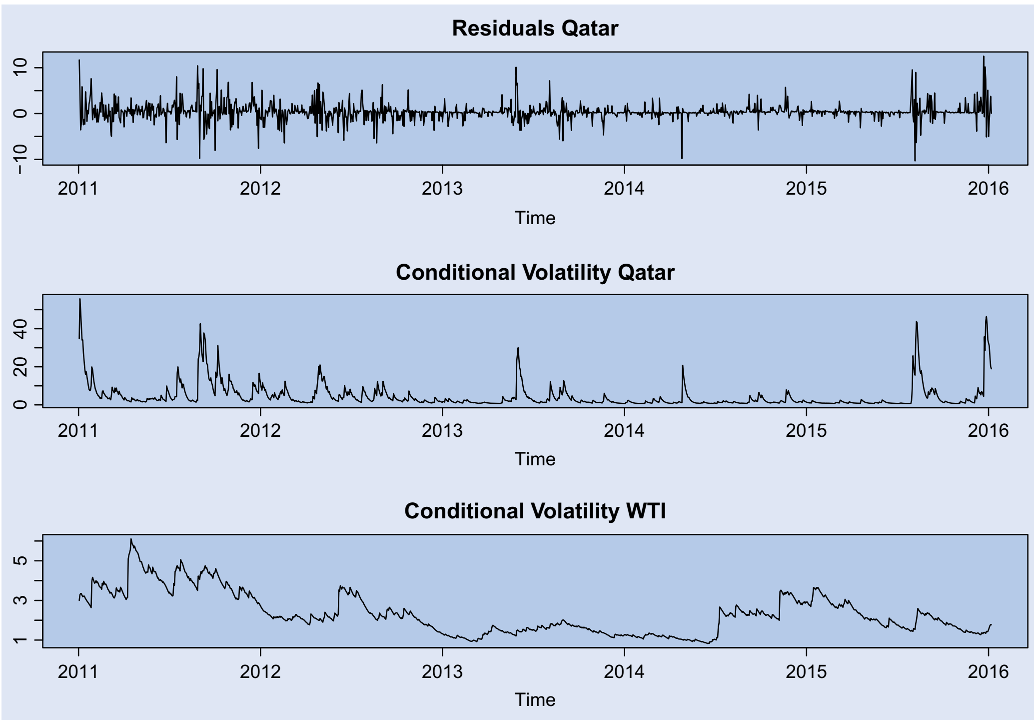

Figure 4. Estimated residuals and conditional volatilities of the Qatar model.

insignificant. Most of the negative reactions of CDS spreads to oil price changes are statistically significant. The findings reported here clearly have to be interpreted as an indication for the fact that increases to the crude oil price help to improve the fiscal stability of oil producing countries. Vice versa falling oil prices obviously seem to have negative consequences for government finances. Interestingly, the reaction of CDS spreads to oil price movements is not significant in the three western industrialized countries US, UK and Norway. A further interesting observation is that only the CDS spread for Qatar seems to react in a negative way to a higher volatility of the residuals of the oil price. This is captured by λ_{21} and might be interpreted against the background that h_t captures ups and downs of the residuals symmetrically. Furthermore, in general, the other variables of GARCH in the mean (λ_{11}, λ_{12} and λ_{22}) are not statistically significant. Only in the case of Qatar the

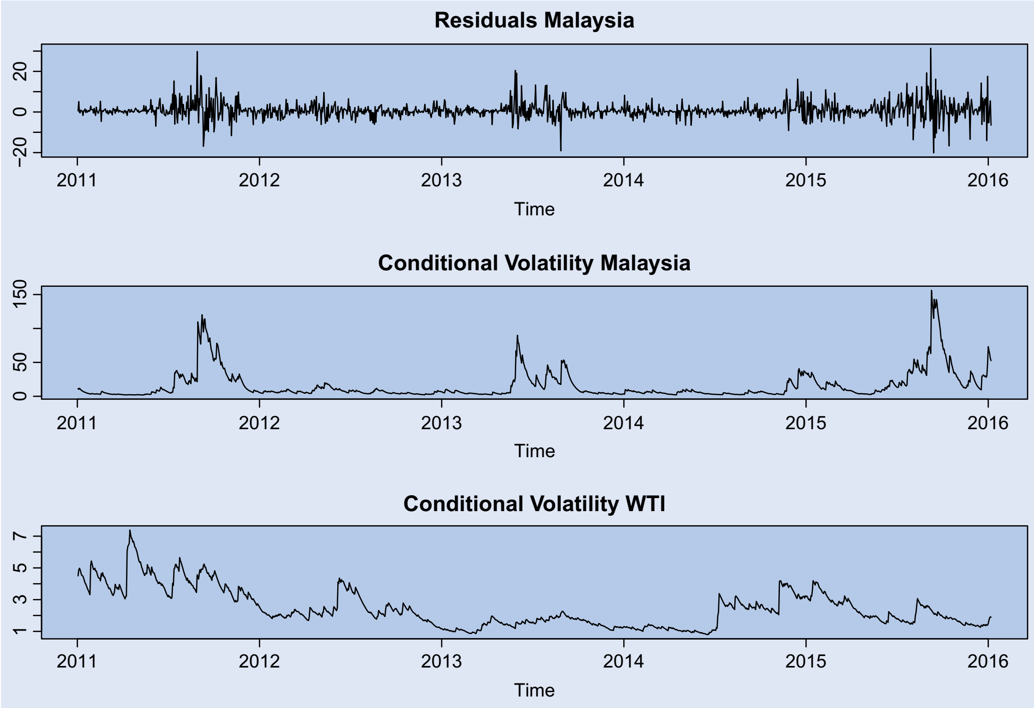

Figure 5. Estimated residuals and conditional volatilities of the Malaysia model.

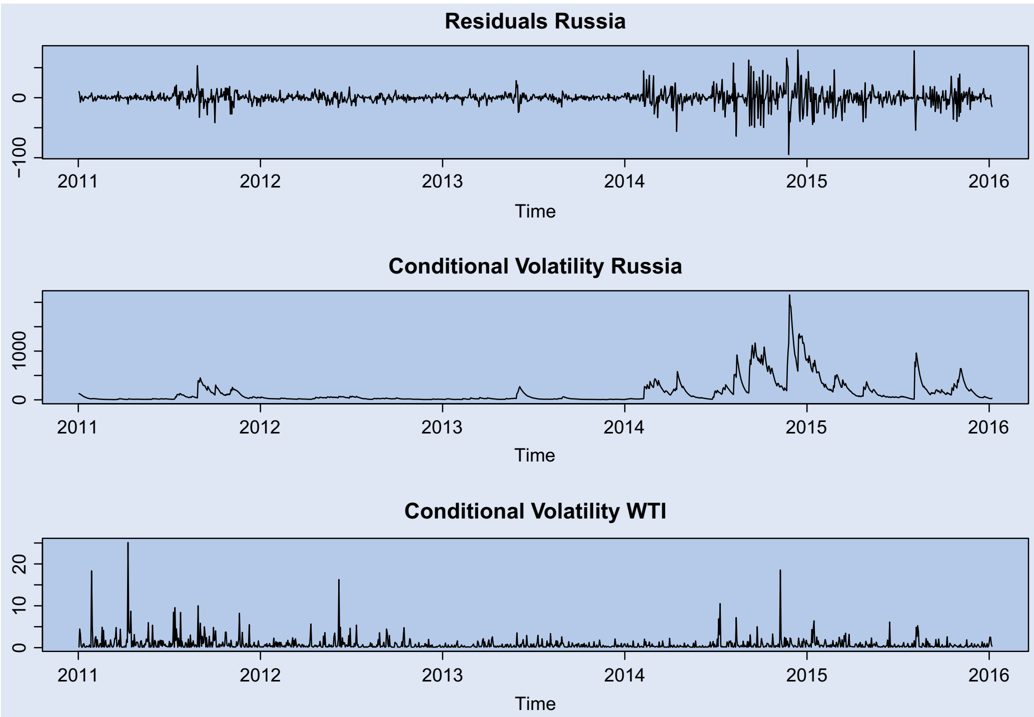

Figure 6. Estimated residuals and conditional volatilities of the Russia model.

effect of the volatility in the CDS error (λ_{12}) is significant for the conditional mean of the CDS.

All residuals of the CDS spread equations, it's conditional variances and the conditional variance of the WTI for the respective model are presented in figures 1–7. In general, the figures and the significant GARCH coefficients (f_1, f_2, g_1, g_2) support the hypothesis of GARCH effects in the residuals. Most importantly, the peaks of the conditional variance in the WTI errors are not coincident with it's counterpart of the CDS residuals. However, in the case of Saudi Arabia and Venezuela GARCH is only significant in the WTI equation. This result is supported by figure 8 for Saudi Arabia—in the case of Venezuela (see figure 9) this might be caused by structural breaks. Thus we skipped the GARCH term and reported the results of the plain VAR. Unsurprisingly, γ_{21} is significant.

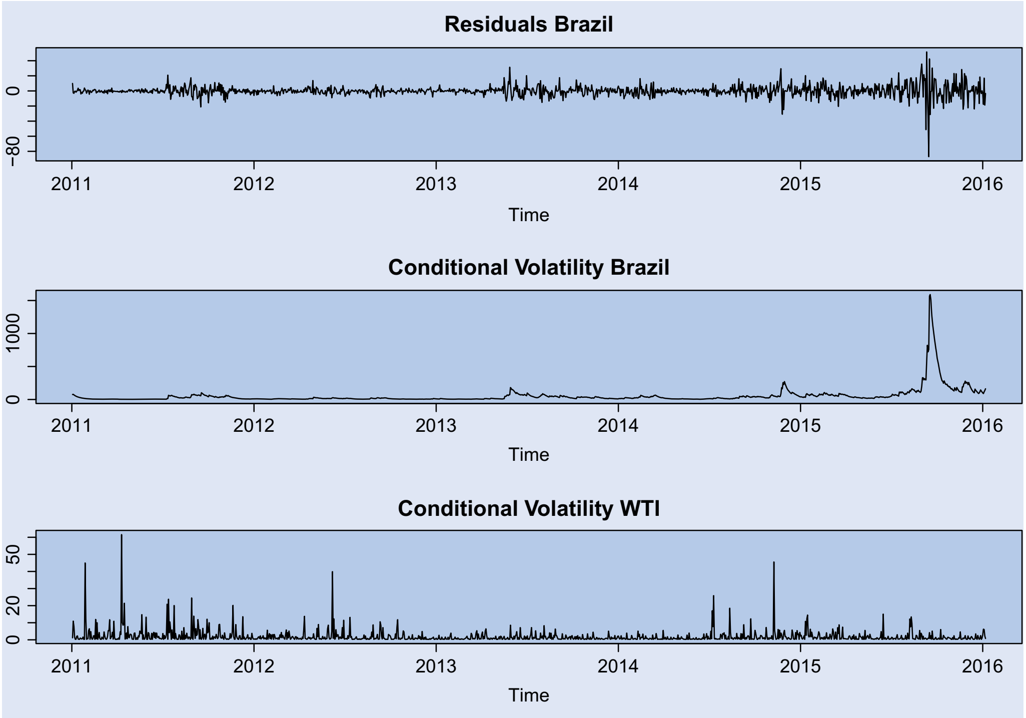

Figure 7. Estimated residuals and conditional volatilities of the Brazil model.

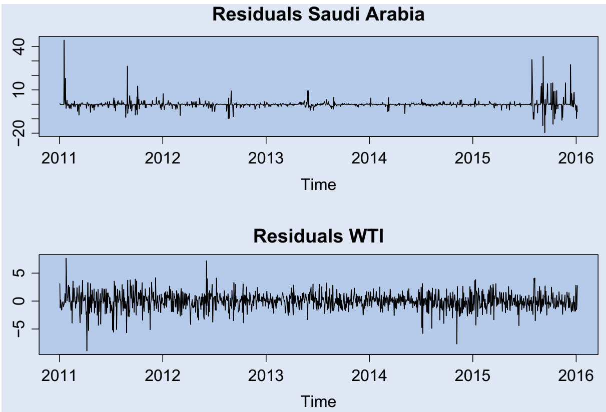

Figure 8. Estimated residuals of the Saudi Ariabia model.

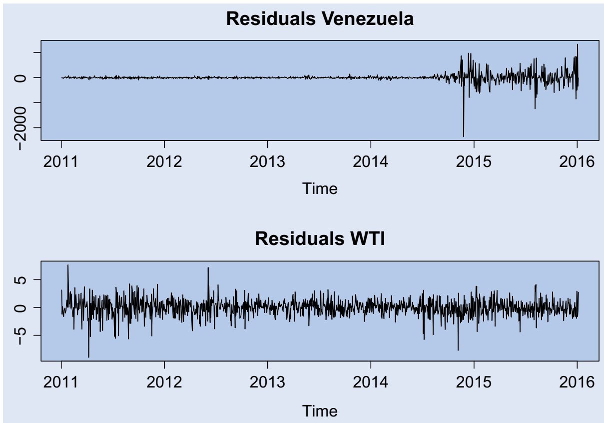

Figure 9. Estimated residuals of the Venezuela model.

6. Conclusion

The results reported above in general do indicate that positive oil price shocks lead to lower sovereign CDS spreads. Interestingly, the UK as a quite diversified economy is the only exception. However, the positive coefficient here is not statistically significant. Most of the negative reactions of CDS spreads to oil price changes are statistically significant. These findings clearly have to be interpreted as an indication for the fact that increases to crude oil price do help to improve the fiscal stability of oil producing countries and vice versa. This, of course, is no surprise at all because positive oil price shocks should increase oil revenues and thereby reduce sovereign credit risk. Summarizing our results crude oil price do seem to matter for the market perception of sovereign credit risk in oil producing countries. Obviously, further empirical research is needed. Most importantly, data from additional countries should be examined. Moreover, the question of asymmetric effects in the State of Qatar is of some interest. Our empirical findings with regard to this problem are inconclusive. Thus, further empirical investigations may be helpful. There are numerous interesting additional research questions. Cointegration (tests are not reported to conserve space) in general seems to be of no relevance in the context examined here. However, this could be a result of structural change distorting cointegration tests. Therefore, further empirical work testing for cointegration among sovereign CDS spreads and oil prices considering structural change might be interesting.

Disclosure statement

No potential conflict of interest was reported by the authors.

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