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The views expressed in this paper are those of the authors and do not necessarily reflect the views of the Bank of Finland.

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The co-movements along the forward curve of natural gas futures: a structural view

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Abstract

This paper studies the co-movements between the daily returns of forwards on natural gas traded in the NYMEX with maturity of 1, 2 and 3 months. We identify a structural multivariate BEKK model using a recursive assumption whereby shocks to the volatility of the returns are transmitted from the short to the long section of the forward curve. We find strong evidence of spillover effects in the conditional first moments, for which we show that the transmission mechanism operates from the shorter to the longer maturity. In terms of reduced form conditional second moments, the shortest the maturity, the higher the volatility of the return, and the more the returns become independent from the others and follow the dynamics of the underlying commodity. The evidence from the structural second moments indicates that the longer the maturity is, the higher the uncertainty about the returns. We also show that the higher the structural variance of a maturity relative to that of another maturity, the stronger the correlation between the two.

Keywords: natural gas prices, forward markets, GARCH, structural VAR

JEL classification numbers: C22, G19

Luonnonkaasufutuurien käyttö termiinituottojen keskinäisen vaihtelun rakenteellisessa mallintamisessa

Suomen Pankin keskustelualoitteita 26/2008

Fabrizio Spargoli – Paolo Zagaglia
Rahapolitiikka- ja tutkimusosasto

Tiivistelmä

Tässä tutkimuksessa tarkastellaan empiirisesti raaka-ainefutuurien termiinituottojen päivittäisiä vaihteluita. Työssä käytetty havaintoaineisto koostuu New Yorkin raaka-ainepörssissä Nymexissä kaupatuista, maturiteetiltaan yhden, kahden ja kolmen kuukauden luonnonkaasufutuureista. Tuottojen vaihtelua selittävän moniulotteisen aikasarjamallin identifiointi perustuu oletukseen, että tuottojen satunnaiset vaihtelut välittyvät lyhyen maturiteetin tuotoista pidemmän maturiteetin tuottoihin. Keskimääräisiin tuottoihin liittyvät estimointitulokset tukevat voimakkaasti heijastusvaikutuksia niin, että häiriöt välittyvät lyhyistä pidempiin maturiteetteihin. Tuottojen varianssia selittävän eli tuottojen keskimääräiseen vaihteluun käytetyn supistetun muodon mallin estimointi puolestaan viittaa siihen, että termiinituottokäyrän ns. lyhyen pään heilahtelujen voimistuessa tuottojen keskinäinen riippuvuus heikkenee ja niiden dynamiikka seuraa läheisesti itse raaka-aineen dynamiikkaa. Tuottojen varianssia selittävän rakenteellisen mallin estimoidut ominaisuudet toisaalta tukevat ajatusta, että tuottoihin liittyvä epävarmuus kasvaa tuottojen maturiteetin pidentyessä. Tuottojen varianssin rakenteelliset estimaatit viittaavat lisäksi siihen, että kahden erimaturiteettisen tuoton välinen korrelointi voimistuu, kun näistä kahdesta toisen tuoton suhteellinen varianssi kasvaa eli kun tuottojen keskimääräinen vaihtelu tietyssä maturiteetissa kasvaa suhteessa muihin maturiteetteihin.

Avainsanat: luonnonkaasun hinta, termiinimarkkinat, GARCH, rakenteellinen VAR

JEL-luokittelu: C22, G19

Contents

1 Introduction

Commodity markets have lost their original purpose of trading and delivery of physical goods nowadays. In fact, they have become the arena for investors interested in futures and forward contracts for hedging. In that sense, understanding the transmission mechanism of volatility across segments of futures markets is a key step for setting up proper hedging strategies. Although a large attention to this issue has been devoted for the oil markets (eg see Lin and Tamvakis, 2001), only a small number of studies is available for the markets for natural gas. For instance, Ewing, Malik and Ozfidan (2002) use the BEKK-GARCH model of Engle and Kroner (1995) to study the correlation between stock indices on oil and natural gas companies traded on the American Stock Exchange. Their results document a strong pattern of volatility transmission between the two indices.

In the present paper, we study the sources for hedging in the forward curve for natural gas on the New York Mercantile Exchange. In particular, we investigate how the joint movements of prices along the maturity structure respond to exogenous shocks, and how the nature of the shocks to prices affects the joint movements between the market segments.

A large number of multivariate models are available to estimate time-varying correlations and volatilities, the most commonly used being the Dynamic Conditional Correlation model of Engle (2002).[1] However the measures derived from these models convey no information about the sources of the correlation between markets. In other words, in order to understand the joint dynamics of the daily prices of natural gas futures, we face the issue of identification, also known as ‘identification problem’, which arises from the lack of knowledge of causal relations. Dealing with the identification problem yields a distinction between ‘structural’ and ‘reduced-form’ conditional volatilities and correlations. The structural moments disentangle the effects of exogenous shocks from the endogenous response, whereas the reduced-form second moments are do not address the identification problem.

Our model includes the forward prices with maturity of one month, two months and three months. We obtain identification by assuming that the conditional second moments of the variables have a recursive structure. This implies that the second moments of the returns on the shortest maturity depend only on their own autoregressive and innovation terms, while those on the longest maturity are a function of all the autoregressive and innovation terms in the system. Summing up, this amounts to estimating a BEKK-GARCH model with parameters that are identified from the reduced-form estimates.

[1] For instance, see Marzo and Zagaglia (2008) for a recent application of the DCC model to energy futures markets.

We find strong evidence of spillover effects for both the first and the second moments along the forward curve. With regard to the first moments, we find that a transmission mechanism operates from the longer to the short maturities. The results also show that the shorter the maturity, the higher the volatility of the return is, and the more the return becomes independent from the other returns and follows the dynamic of the underlying commodity. This consideration holds, however, for the reduced-form moments. For what concerns the structural moments, we find returns on the 3 months maturity display the highest volatility, which implies a large uncertainty on the returns. We find also that the higher the structural variance of a return relative to that of another return, the stronger the correlation is between the two.

This paper is organized as follows. In the first section we outline our structural GARCH model along with the identification technique. In the second section we compare our methodology to others used in the literature and in the third we present the main results. Concluding remarks are presented in section four.

2 The structural GARCH model

Let us assume that the evolution of the variables can be summarized by a structural vector-autoregressive (VAR) model

$$Ax_t = \psi + \Phi(L)x_t + \eta_t$$

where x_t is the vector of the endogenous variables, that is the one month, two months and three months natural gas forward prices, ψ is a vector of constants, A is the matrix of structural parameter, which is normalized to have 1 on the main diagonal, and η_t is the vector of structural shocks, which are assumed to be distributed as a $N(0,\eta_t)$. Therefore, structural innovations are assumed to exhibit conditional heteroskedasticity, which in this set-up is modeled as a BEKK-GARCH (Engle and Kroner, 1995), that is

$$h_t = CC' + Gh_{t-1}G' + T\eta_{t-1}\eta'_{t-1}T'$$

Furthermore, we assume the G and T matrices to be lower triangular.[2] This hypothesis, which is not new in the GARCH literature, implies that the conditional variances of the returns have a recursive structure, which seems to be

[2] The assumptions about the lags of the autoregressive, the moving average term and the degreeof generality are made in order to keep the estimation feasible, given the high number of parameters in a BEKK-GARCH and because they are common in the literature.

reasonable given that term structure variables are going to be used.[3] In particular, it is likely that the conditional variance of the shorter term return depends only on its own lags, whereas the conditional variance of the longer term return depends on its lags and the lags of the others. In other words, shocks to returns' volatility are more likely to be transmitted from the short to the long section of the forward curve than the other way around. Moreover, the model allows for a non zero covariance among structural innovations, which seems reasonable given that term structure variables are likely to be affected by common factors which determine a comovement among them.

As usual in VAR models, estimation through OLS leads to asymptotically biased estimators for the structural parameters owing to the simultaneity among the endogenous variables in the system. In other words, some of the regressors are not exogenous because their source of variation is represented by the dependent variable in the same equation through another equation in the system. In order to gain consistency, an identification procedure which makes the source of variation in the regressors exogenous is required. However, most of the common techniques, which rely on restrictions on the joint behaviour of the variables in the system,[4] can hardly be applied in this set-up because the high frequency of the data does not allow disentangling reasonable links among the variables.

Yet, a feasible technique exhists in this case, which relies on heteroskedasticity in order to identify the VAR. This idea has been originally introduced by Wright (1928) and recently developed by Rigobon (2003). The heteroskedasticity approach to identification amounts to using the information from time-varying volatility as a source of exogenous variation in the endogenous variables. To see this, starting from the assumptions about the structural model, it is worth considering the reduced form VAR model, which is given by

$$x_t = c + F(L)x_t + v_t$$

where $c = A^{-1}\psi$, $F(L) = A^{-1}\Phi(L)$ and $v_t = A^{-1}\eta_t$ are the reduced-form innovations, whose variance-covariance matrix is a combination of the variance-covariance matrix of the structural-form innovations, that is

$$H_t = Bh_tB'$$
$$H_t = BCC'B' + BGh_{t-1}G'B' + BT\eta_{t-1}\eta'_{t-1}T'B'$$

[3] The use of this recursive assumption is not new in the literature on structural GARCH. For instance, Cassola and Morana (2006) order term structure variables from short to long maturities to identify the co-movements in volatility in the Euro area money market.

[4] There is a wide literature about this topic, starting from Sims (1980). Basically, identification is achieved imposing either exclusion restrictions (imposing a recursive order, like in Sims (1980), or not, like in Bernanke (1986), Blanchard and Watson (1986) and Sims (1986)) or long run restrictions.

In this formulation the variance-covariance matrix of the reduced-form innovations is a function of the structural innovations, which the econometrician does not know. However, we can use the equality $\eta_t = Av_t$ to show that

$$\eta_t \eta_t' = A v_t v_t' A'$$

and to represent H_t in terms of the reduced-form innovations as

$$H_t = BCC'B' + BGAH_{t-1}A'G'B' + BTAv_{t-1}v_{t-1}'A'T'B'$$

It should be noted that there are 24 parameters in the reduced form GARCH model as well as in the structural form one because of the recursiveness assumption of the heteroskedastic process and the shape of the matrix A. Therefore, the structural parameters of the VAR model can be identified with no need for additional restrictions on the links between the variables.

After estimating the model, we compute impulse-response functions. In structural GARCH models, these functions show the impact that a shock produces on the conditional second moments of the variables in the system. However, differently from the impulse response functions for a standard VAR, the impulse responses of a structural GARCH depend both on the magnitude of the shock and on the period during which the shock itself takes place. This is due to the fact that the residuals enter the model in quadratic form. Hence, differently from the case of linear models, the magnitude of the effects of a shock is not proportional to the size of the shock itself. This allows us to compute a distribution of impulse responses following each shock. To that end, we use the concept of Volatility Impulse Response Functions – VIRF – proposed by Hafner e Herwartz (2006). The impulse-response function for a vech-GARCH model can be written as

$$V_t(\xi_0) = E[vech(H_t)|\xi_0, I_{-1}] - E[vech(H_t)|I_t]$$

The response at time t of the variances and covariances following a shock η in t=0 – denoted as $V_t(\eta_0)$ – is equal to the difference, conditioned on the information set at time -1 (I_{-1}) and on the shock η_0, of the variance (or covariance) at t from its expected value conditional on the information set of period -1.[5]

[5] Details on the analytical formulas used for the calculation of the VIRFs can be found in Spargoli and Zagaglia (2007).

2.1 Comparison with the literature on identification through volatility

In Rigobon (2003) and Rigobon and Sack (2003b. 2004), identification is obtained through regimes of volatility, that is, subsamples across which there are shifts in the volatility pattern. Therefore, there is no more a unique variance-covariance matrix, but as many matrices as the number of regimes, each one providing a set of equations which can be used to recover the structural parameters. However, it is sometimes not clear how to identify volatility regimes from the data, so that the procedure may turn out to be arbitrary and determine a loss in efficiency.

To solve this problem, Rigobon and Sack (2003a) consider a similar identification technique allowing for continous regime changes instead of discrete ones. In other words, these authors postulate a GARCH process for the variance of the structural innovations, which are also assumed to have a zero mean and to be uncorrelated. These hypothesis impose structure on the reduced form variance-covariance matrix, which can be estimated using reduced-form residuals, and allows the identification of the structural parameters.

The formulation of Rigobon and Sack (2003a), however, does not guarantee that the variance-covariance matrix is positive-definite, which is a problem typical of every vector – vech – GARCH. In order to cope with this problem, Spargoli and Zagaglia (2007) rely on the multivariate BEKK-GARCH of Engle-Kroner (1995) to analyse the co-movements between oil futures prices on the NYMEX and ICE. They assume the structural-form innovations to have zero mean, to be correlated and to follow a BEKK-GARCH process and estimate the restricted reduced-form model. Identification of the structural parameters is achieved like in Rigobon and Sack (2003a) through restrictions on the conditional variance-covariance matrix of the reduced-form innovations.

The identification technique closest to the one used in this paper is in Rigobon (2002), which studies the contagion effects of the 1997 Mexican crisis on Argentina, Columbia, Venezuela and Brazil. The author still assumes a heteroskedasticity process (both an ARCH and a GARCH model in the vech form, with no structural correlation) for the structural innovations, but structural parameters are identified from the reduced form parameters, which are a function of the structural ones. In other words, structural parameters are recovered by solving a system with as many equations as unknowns.

In this paper, identification is grounded on the same basis. However, there are two main differences from Rigobon (2002): a BEKK-GARCH process is assumed for the structural innovations and a recursive structure is imposed on the volatility. The former guarantees the positive-definiteness of the variance-covariance matrices by construction, while the latter captures a reasonable hypothesis about the transmission of shocks among term structure variables and allows the identification of structural parameters.

In conclusion, it is worth mentioning a further point about the identification technique adopted in this paper. As usual for structural vector autoregressions, identification is achieved imposing restrictions on structural parameters. However, these restrictions do not involve the links among endogenous variables, but the conditional variances of their structural innovations. Therefore, this technique seems to be much more suitable for dealing with high frequency data than the traditional ones.

3 Results

We estimated the model using daily data from the 19th of January 1994 to the 27th of April 2007 obtained from Platt's. We compute the returns in percentage points from the two series and we obtained a total of 3307 observations. The time series are plotted in figure 1 and Table 1 presents some summary statistics. The returns exhibit a typical behaviour for financial time series. In particular, they have a mean and a median very close to zero and exhibit a remarkable leptokurtosis, as shown by the very high kurtosis index. The null hypothesis of normality is strongly rejected by the Jarque-Bera test. Table 1 shows also that the longer the maturity, the lower is the dispersion of the returns. In fact, their stardard deviation is the highest for the 1 month maturity and the lowest for the 3 months maturity, and the maximum and minimum values are bigger for the shortest maturity. Furthermore, the distribution of the returns is positively skewed for the 1 month and 2 months maturities and negatively skewed for the 3 months maturity.

As regards the estimation procedure, we first of all estimated a VAR model including the returns on the 1 month, 2, and 3 months maturity as endogenous variable in order to obtain reduced-form residuals. Then, we used these residuals as the innovations in our reduced-form BEKK-GARCH model, which has been estimated through maximum likelihood. Given the number of parameters involved in the estimation and the nonlinearity of the likelihood fuction, special care must be used to address the presence of kinks and local maxima. Therefore, we have chosen to run a number of initial steps through simulated annealing in order to obtain robust estimates of the initial points for the maximization step. In the second round, we have used gradient-based optimization methods conditional on the initial point from simulated annealing.

We estimate a VAR model for the conditional mean including a constant and two lags of the endogenous variables. The optimal lag is chosen by looking at a number of criteria reported in tables 2 and 3. The estimates of the reduced-form

VAR and BEKK models are in tables 4 and 5. This provides us with a system of 24 nonlinear equations that can be solved in the 24 unknowns (that is the 24 structural parameters)

$$vech(C^*C^{*'}) = vech(BCC'B')$$
$$vec(G^*) = vec(BGA)$$
$$vec(T^*) = vec(BTA)$$

where asterisks denote the reduced-form parameters. We have computed the standard errors of the structural parameters using the delta method outlined in Appendix. The BEKK in structural form is reported in table 6. We obtain the following representation of the conditional second moments

$$
\begin{aligned}
h_t = &\begin{bmatrix} 0.51 & 0 & 0 \\ 0.43 & 0.38 & 0 \\ 0.44 & 0.39 & 0.36 \end{bmatrix} \begin{bmatrix} 0.51 & 0 & 0 \\ 0.43 & 0.38 & 0 \\ 0.44 & 0.39 & 0.36 \end{bmatrix} \\
&+ \begin{bmatrix} 0.65^* & 0 & 0 \\ -0.05^* & 0.87^* & 0 \\ 0.06^* & 0.05 & 0.89 \end{bmatrix} \eta_{t-1}\eta'_{t-1} \begin{bmatrix} 0.65^* & 0 & 0 \\ -0.05^* & 0.87^* & 0 \\ 0.06^* & 0.05 & 0.89 \end{bmatrix} \\
&+ \begin{bmatrix} 1.23 & 0 & 0 \\ 0.13 & 0.50^* & 0 \\ -0.18^* & 0.16 & 0.25 \end{bmatrix} h_{t-1} \begin{bmatrix} 1.23 & 0 & 0 \\ 0.13 & 0.50^* & 0 \\ -0.18^* & 0.16 & 0.25 \end{bmatrix}
\end{aligned}
$$

We should stress that proposition 2.1 of Engle and Kroner (1995) guarantees that the BEKK model is identified because the diagonal elements of C, as well as G_{11} and T_{11}, are positive. From the results, we can see that there is no significant coefficient in the constant matrix. As regards the autoregressive component, we find two coefficients which are significantly different from zero. The highest in absolute value is the 2,2 element, while the other one, which has a negative sign, is the 3,1 element. As regards the moving average matrix, we find two out of six coefficients which are not significantly different from zero. The highest coefficients in absolute value are the 1,1 and 2,2 elements, while the 3,1 and 2,1 elements are significant but small.

As regards the conditional mean parameters, we get the following estimates

$$
\begin{aligned}
r_t^{1m} &= -0.15^* r_t^{2m} - 0.04 r_t^{3m} + \beta_{11} r_{t-1}^{1m} + \gamma_{11} r_{t-2}^{1m} + \beta_{12} r_{t-1}^{2m} + \gamma_{12} r_{t-2}^{2m} \\
&\quad + \beta_{13} r_{t-1}^{3m} + \gamma_{13} r_{t-2}^{3m} + \eta_{1t} \\
r_t^{2m} &= 1.03^* r_t^{1m} - 0.05^* r_t^{3m} + \beta_{21} r_{t-1}^{1m} + \gamma_{21} r_{t-2}^{1m} + \beta_{22} r_{t-1}^{2m} + \gamma_{22} r_{t-2}^{2m} \\
&\quad + \beta_{23} r_{t-1}^{3m} + \gamma_{23} r_{t-2}^{3m} + \eta_{2t} \\
r_t^{3m} &= -0.13 r_t^{1m} + 1.01^* r_t^{2m} + \beta_{31} r_{t-1}^{1m} + \gamma_{31} r_{t-2}^{1m} + \beta_{32} r_{t-1}^{2m} + \gamma_{32} r_{t-2}^{2m} \\
&\quad + \beta_{33} r_{t-1}^{3m} + \gamma_{33} r_{t-2}^{3m} + \eta_{3t}
\end{aligned}
$$

The parameters of the equations for the conditional mean, for which stars denote 5% significance, indicate the direct effect that a structural shock to a return causes to the conditional mean of the other returns. We find the largest spillover effects between the returns on adjacent maturities along the forward curve, so that a transmission mechanism linking the shortest and the longest maturity is clearly recognizable. Furthermore, we find no evidence of significant spillover effects between the returns on maturities which are not close to each other. In particular, a 1 basis point increase in the return of the one-month natural gas forward causes a 1.03 basis points increase in the return of the two-month forward which, in turn, determines a 1.01 basis point increase in the return on the three months forward. There is also evidence of a negative, yet smaller in magnitude, spillover effect working the other way around. In particular, we find that a 1 basis point increase in the return on the three-month maturity causes a 0.05 basis point decrease on that on the two-month maturity which, in turn, generates a 0.15 basis point decrease in the return on the shortest maturity. Summing up, the results from the estimated structural coefficients suggests the existence of spillovers among the returns of the forward curve which are mainly due to interactions in term of conditional mean rather than conditional second moments.

From the estimated structural coefficient it is possible to calculate the conditional second moments of the returns for both the structural and reduced form. In particular, the former give a representation of the dynamics of the structural innovations as such, which means that they do not incorporate the indirect effects due to spillovers among the returns. Figures 2 and 3 plot the conditional structural variances and correlations. The conditional structural variance of r_t^{3m} is the highest over the sample except for its middle section, where the conditional structural variance of r_t^{1m} overcomes it. The conditional structural variance of r_t^{3m} shows frequent and high peaks, while that of r_t^{2m} is the lowest over the sample. The three conditional variances show peaks at the same points of the sample.

Figure 3 shows that the returns on the three maturities are strongly correlated, and that there are frequent peaks that push the correlations to extreme levels. It is difficult to detect a pattern in the dynamics of these conditional moments, given their frequent oscillations. However, one can say that the structural correlation between r_t^{1m} and r_t^{2m} becomes positive, and oscillates around 0.5. after 1996. Yet, there are frequent peaks that make it negative and reach -1. Combining the evidence from the conditional structural correlations with that on the conditional structural variances, we can notice that the bigger a conditional variance at a maturity relative to that of another maturity, the stronger the correlation is between the corresponding returns. For example, in the period between observation 1994 and 1996 the correlation between r_t^{1m} and r_t^{2m} oscillates around a value between 0 and -0.5 and the structural variance of r_t^{2m} is almost equal to the one of r_t^{1m}. In the subsequent range of observations, however, the latter becomes much bigger than the former and, at the same time, the correlation between the two returns oscillates around a higher mean in absolute value. The same considerations apply also for the conditional structural correlation between r_t^{3m} and r_t^{1m}: when the conditional structural variance of r_t^{3m} is bigger than r_t^{1m} – ie in the period between 1994 and 1996 – the correlation is higher in absolute value than in the following period where the two structural variances are approximately equal. This also holds for the conditional structural correlation between r_t^{2m} and r_t^{3m}. This suggests that the comovement between the returns is driven by the dynamics of the most volatile.

Figures 4 and 5 plot the reduced-form conditional variances and correlations that incorporate the linkages among the returns. The reduced-form conditional variances are generally smaller than those of the structural form, which means that the spillovers among markets contribute to a reduction of the volatility of structural innovations. The size of the reduced-form conditional variances seems to be the inverse of that of the structural-form variances. This supports the view that forward prices are more volatile for short maturities, given they are the most traded and liquid along the maturity structure. Therefore, even if the structural conditional variance of r_t^{1m} is the lowest over the sample, the reduced-form variance is the highest because of the spillovers and linkages among the returns.

Figure 5 shows that the correlation between r_t^{1m} and r_t^{2m} seems to have three regimes. The first regime goes from 1994 to 1996, when it oscillates around a mean value between 0.5 and 1. The second regime is from 1996 to 2004, where it fluctuates around a mean value around 0. The third regime is similar to the first. The same consideration holds for the reduced-form conditional correlation between r_t^{3m} and r_t^{1m}. This means that the return on the shortest maturity is independent from those on the other maturities, which can be explained by the fact that the closer the expiration date of a derivative product, the more its price follows the price of the underlying commodity. The reduced-form conditional correlation between r_t^{3m} and r_t^{2m} has smaller oscillations and has a mean value comprised between 0.5 and 1. Furthermore, it is perfectly positive in most part of the sample. These facts, together with a structural-form conditional correlation with a negative mean value, could be interpreted as a suggestion that the two-month and three-month maturities are held for hedging purposes.

Now we turn to the analysis of the persistence of the effects of the shocks, which we carried out through volatility impulse responses. As explained earlier, given that GARCH are non-linear in the innovations, the effect of a shock depends both on the size and the timing. Therefore, our use of VIRFs is twofold. On the one hand, we can plot traditional impulse responses after a specific shock occurred at a specific point in time. On the other hand, we can compute the distribution of VIRFs, that is we can calculate impulse responses for each shock and then determine their frequency. This should be done for each time horizon of the VIRF.

Figure 6 shows the impulse responses on a potentially significant date, namely the second Gulf war shock, which takes place on the 20th of March 2003. The shock is absorbed very quickly, given that the effect on all the conditional moments vanishes after 3 or 4 days. The shock has a negative impact on the conditional variances, in particular on that of r_t^{1m}, and on the correlation between r_t^{3m} and r_t^{2m} while it has an impact of positive sign on the correlations between r_t^{1m} and r_t^{2m}, and on those between r_t^{3m} and r_t^{1m}. The finding about the reaction of the conditional variance of r_t^{1m} confirms that the returns on the shortest maturities are more volatile. The response of the conditional correlation shows that a shock to the returns on the one-month maturity determines an effect of the same sign as that on the two and three-month maturity, which can be interpreted as a transmission mechanism of volatility shocks. However, this does not hold for the maturities at two and three months. This suggests that the transmission process takes place directly from the short maturity to the rest of the forward curve.

Turning to the distribution of the VIRFs, figures 7 and 8 report the 1st, 10th, 25th, and the 50th, 75th, 90th and 99th percentiles. At a first glance, we can again notice that the effect of the shocks tend to absorbed very quickly, given that after 3 or 4 days all the percentiles become close to zero. It should be noted also that the immediate impact of the shock has a great dispersion, because the extreme percentiles of the distribution are very far from each other for all the VIRFs. It is interesting to analyze the median of the VIRF distribution, in order to understand whether the shocks have a positive or a negative impact on the conditional moments. From Figure 7, it is evident that the shocks exert mainly a negative impact on the conditional variance of r_t^{2m} and r_t^{3m} and on the correlation between r_t^{2m} and r_t^{3m}, given that even the 75th percentile is negative. As regards the other moments, the distribution of their VIRFS is symmetric because the 50th percentile is approximately zero. Therefore, the shocks generate effects of positive and negative sign in the same proportion.

4 Conclusion

We study the relation between the returns on one-month, two-month and three-month maturities of natural gas forwards traded in the New York Mercantile Exchange. We estimate a BEKK-GARCH model (Engle and Kroner, 1995) from which we identify the parameters of a structural model VAR model with heteroskedasticity in the structural innovations. In this way, we obtain estimates of the spillovers among the three returns both in term of the first and second conditional moments.

We find that the evidence about conditional second moments is in line with that concerning forwards in general: the shorter the maturity, the higher the volatility of the return, and the more the return becomes independent from the other returns, and follow the dynamics of the underlying commodity price. We find also that the returns on the three-month maturity are those with the highest volatility, which could be interpreted as a consequence of the greater uncertainty that characterizes the factors guiding longer maturities. Another result is that the co-movement between the returns is driven by the dynamics of the most volatile. We detect a transmission mechanism that runs from the short to the long section of the forward curve. Finally, we also show that the effects of the shocks on the conditional second moments have a very little persistence, given that they vanish after 4 or 5 days.

References

Cassola, N – Morana, C (2006) **Co-movements in Volatility in the Euro Market.** ECB Working Paper, No. 703. December.

Engle, R F (2002) **Dynamic Conditional Correlation: a Simple Class of Multivariate Generalized Autoregressive Conditional Heteroskedasticity Models.** Journal of Business and Economic Statistics, 20(3), 339–350.

Ewing, B T – Malik, F – Ozfidan, O (2002) **Volatility Transmission in the Oil and Natural Gas Markets.** Energy Economics, 24(6), 525–538.

Engle, R F – Kroner, K F (1995) **Multivariate Simultaneous Generalized ARCH.** Econometric Theory, 11(1), 122–150.

Hafner, C M – Herwartz, H (2006) **Volatility Impulse Responses for Multivariate GARCH Models: An Exchange Rate Illustration.** Journal of International Money and Finance, 25(5), 719–740.

Lin, S X – Tamvakis, M N (2001) **Spillover Effects in Energy Futures Markets.** Energy Economics, 23(1), 43–56.

Marzo, M – Zagaglia, P (2008) **A Note on the Conditional Correlation between Energy Prices: Evidence from Futures Markets.** Forthcoming in Energy Economics.

Rigobon, R (2003) **Identification through Heteroskedasticity.** Review of Economics and Statistics, 85(4), 777–792.

Rigobon, R (2002) **The Curse of non Investment Grade Countries.** Journal of Development Economics, 69, 423–449.

Rigobon, R – Sack, B (2003a) **Spillovers across U.S. Financial Markets.** Unpublished manuscript, MIT Sloan School of Management.

Rigobon, R – Sack, B (2003b) **Measuring the Reaction of Monetary Policy to the Stock Market.** Quarterly Journal of Economics, 118(2), 639–669.

Rigobon, R – Sack, B (2004) **The Impact of Monetary Policy on Asset Prices.** Journal of Monetary Economics, 51(8), 1553–1575.

Spargoli, F – Zagaglia, P (2007) **The Co-movements between Futures Markets for Crude Oil: Evidence from a Structural GARCH Model.** Unpublished manuscript, Stockholm University, August.

Wright, P (1928) **The Tariff on animal and Vegetable Oils, The institute of Economics.** The Macmillan Company, New York.

Appendix

The Delta method

Let θ be the reduced form parameters and $g(\theta)$ the function which maps those into the structural form parameters. By taking a first order Taylor approximation of this function around the reduced form estimate θ_0. We get

$$g(\theta) - g(\theta_0) \cong \Delta_\theta g(\theta_0)(\theta - \theta_0)$$

Let Ω be the variance-covariance matrix of the reduced form parameters evaluated in θ_0. The variance of the structural parameters is therefore given by

$$V\big(f(\theta) - g(\theta_0)\big) = \Delta_\theta g(\theta_0)\Omega\Delta_\theta g(\theta_0)'$$

In the paper, we calculate numerical derivatives given the high nonlinearity of the $g(\theta)$ function.

Tables 1–6

Table 1. **Summary statistics of the returns**

	1 month	2 months	3 months
Mean	0.0114	0.0095	0.0084
Median	0	0	0.079
Maximum	38.313	24.13	19.57
Minimum	-31.32	-20.61	-26.74
Std. Dev.	3.93	3.366	2.94
Skewness	0.603	0.40	-0.130
Kurtosis	10.527	7.629	9.164
Jarque-Bera	8021.701	3046.41	5255.301
Probability	0	0	0
No. obs.	3313	3313	3313

Table 2. **Reduced form VAR lag selection criteria**

Lag	LogL	LR	FPE	AIC	SC	HQ
0	25673.06	NA	3.34e-11	-1560964	-1560408	-1560765
1	25753.52	1607234	3.20e-11	-1565310	-1563085	-1564513
2	25792.18	77.15761*	3.14e-11*	-15.67113*	-15.63219*	-15.65719*
3	25796.62	8856206	3.15e-11	-1566836	-1561274	-1564845
4	25798.84	4420187	3.16e-11	-1566424	-1559192	-1563835
5	25801.43	5170080	3.17e-11	-1566034	-1557134	-1562848
6	25806.54	1015058	3.18e-11	-1565797	-1555229	-1562014
7	25810.14	7147179	3.19e-11	-1565469	-1553231	-1561088
8	25818.26	1612186	3.19e-11	-1565416	-1551509	-1560437

* indicates lag order selected by the criterion

LR: sequential modified LR test statistic (each test at 5% level)

FPE: Final prediction error

AIC: Akaike information criterion

SC: Schwarz information criterion

HQ: Hannan-Quinn information criterion

Table 3. **Reduced form VAR lag Wald test**

	Equation 1	Equation 2	Equation 3	Joint
Lag 1	3172318	5073861	3412858	1913949
	[5.99e-07]	[0.166468]	[0.332241]	[0.000000]
Lag 2	1876626	9595461	2371787	7325090
	[0.000306]	[0.022337]	[0.498908]	[3.50e-12]
Lag 3	3835597	6175865	5143290	9436920
	[0.279773]	[0.103361]	[0.161600]	[0.397958]
Lag 4	2380706	2101958	2328731	4238994
	[0.497236]	[0.551517]	[0.507039]	[0.894990]
Lag 5	1219282	2496316	2523415	4269831
	[0.748384]	[0.475957]	[0.471074]	[0.892772]
Lag 6	4848730	3855526	2576565	9996472
	[0.183215]	[0.277493]	[0.461613]	[0.350771]
df	3	3	3	9

Table 4. **Reduced form VAR estimation output**

	Equation 1	Equation 2	Equation 3
1st lag Equation 1	-0.211245	-0.011981	-0.041453
	(0.04409)	(0.03801)	(0.03354)
	[-4.79164]	[-0.31516]	[-1.23599]
2nd lag Equation 1	-0.105337	-0.000985	-0.014208
	(0.04407)	(0.03800)	(0.03352)
	[-2.39044]	[-0.02591]	[-0.42382]
1st lag Equation 1	0.173529	-0.057573	0.067575
	(0.07965)	(0.06868)	(0.06059)
	[2.17857]	[-0.83823]	[1.11520]
2nd lag Equation 1	-0.026044	-0.133273	-0.022582
	(0.07938)	(0.06845)	(0.06039)
	[-0.32809]	[-1.94701]	[-0.37394]
1st lag Equation 1	0.042591	0.051355	-0.040767
	(0.06255)	(0.05394)	(0.04759)
	[0.68087]	[0.95209]	[-0.85667]
2nd lag Equation 1	0.136699	0.137681	0.034856
	(0.06234)	(0.05375)	(0.04742)
	[2.19286]	[2.56135]	[0.73500]
Constant	0.000401	0.000386	0.000413
	(0.00068)	(0.00058)	(0.00051)
	[0.59352]	[0.66314]	[0.80296]

Standard errors in () and t-statistics in []

Table 5. **Reduced form BEKK-GARCH estimates**

Parameter	Coefficient	Std error	P value
C^*_{11}	1.241	0.54313	0.01119
C^*_{21}	0.88924	0.48203	0.03259
C^*_{31}	0.53891	0.02662	0.00857
C^*_{22}	0.034808	0.85398	0.48375
C^*_{32}	-0.021953	0.13455	0.5648
C^*_{33}	0.065602	0.45551	0.44275
T^*_{11}	1.0943	0.76172	0.07547
T^*_{21}	-0.13526	0.93874	0.55728
T^*_{31}	-0.1854	0.28794	0.74015
T^*_{12}	-0.51141	0.73587	0.75644
T^*_{22}	0.79667	0.74063	0.14108
T^*_{32}	0.34793	0.27594	0.10372
T^*_{13}	-0.22695	0.03743	0.7594
T^*_{23}	-0.36592	0.06498	0.4857
T^*_{33}	0.10044	0.02423	1.7e-05
G^*_{11}	0.67192	0.01132	0.01284
G^*_{21}	0.068378	0.07936	0.19448
G^*_{31}	0.078996	0.02762	0.00213
G^*_{12}	0.20773	0.67045	0.37835
G^*_{22}	0.88629	0.01285	0.39458
G^*_{32}	-0.008632	0.00086	0.85720
G^*_{13}	-0.018566	0.86871	0.50852
G^*_{23}	-0.061878	0.15036	0.65964
G^*_{33}	0.86973	0.05248	0.00385

Table 6. **Structural form BEKK-GARCH estimates**

Parameter	Coefficient	Std error	P value
C_{11}	0.50861	2.7385	0.42634
C_{21}	0.43181	2.8163	0.43908
C_{31}	0.44143	2.6102	0.43286
C_{22}	0.38185	2.724	0.44426
C_{32}	0.39785	2.0879	0.42444
C_{33}	0.3627	1.3123	0.39113
T_{11}	-1.0298	0.24677	0.99998
T_{21}	0.12696	0.078	0.05185
T_{31}	0.14087	0.07201	0.02527
T_{12}	-1.0015	0.12688	0.93854
T_{22}	0.03962	0.03527	0.13066
T_{32}	0.047867	0.01085	5.3e-06
T_{13}	0.65397	0.01819	0.04859
T_{23}	-0.047549	0.0151	0.99917
T_{33}	0.060895	0.00544	0.00501
G_{11}	0.87415	0.12787	4.9e-12
G_{21}	0.054971	0.1569	0.36305
G_{31}	0.89849	0.59222	0.06467
G_{12}	1.2318	5.1524	0.40553
G_{22}	0.12727	1.7928	0.47171
G_{32}	-0.18377	0.01487	0.59681
G_{13}	0.50275	0.18147	0.00282
G_{23}	0.16212	3.0073	0.47851
G_{33}	0.25413	0.56898	0.32758

Figures 1–8

Figure 1. **Plot of the data series**

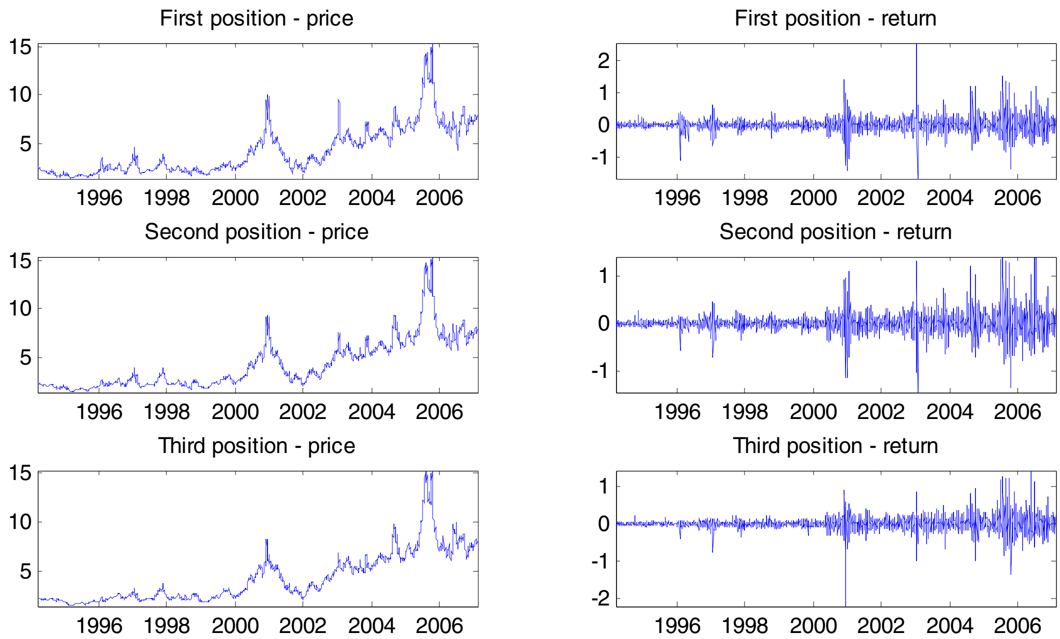


Figure 2. **Structural conditional variances**

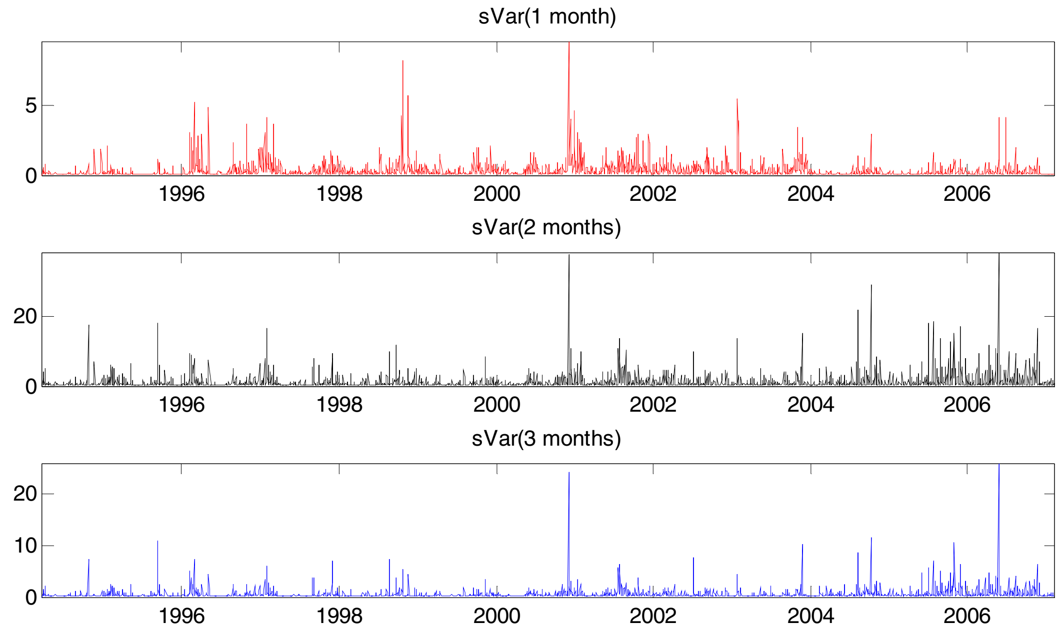


Figure 3. **Structural conditional correlations**

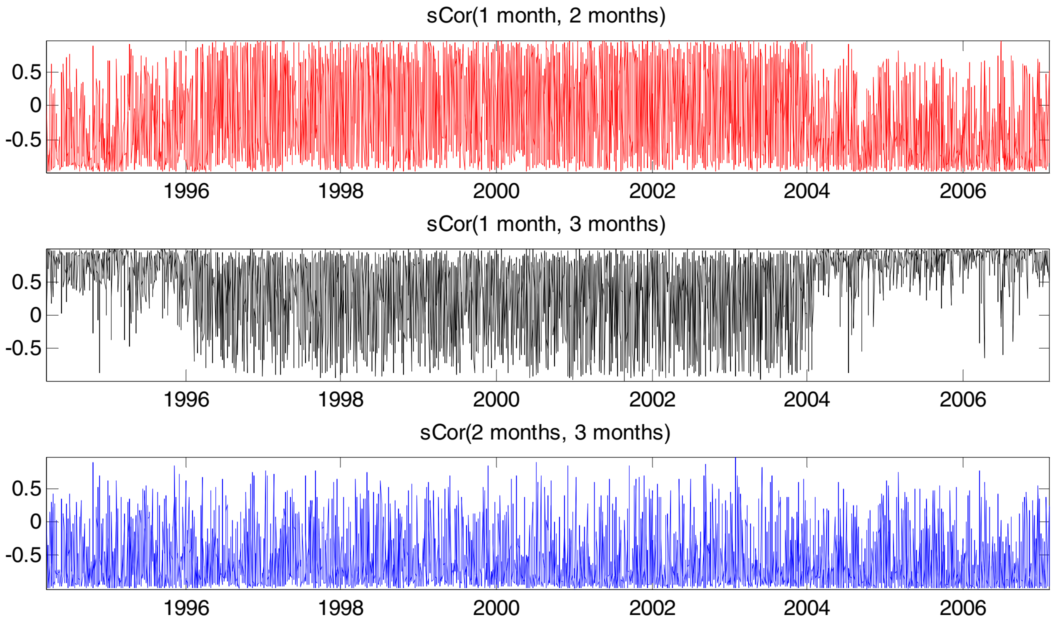


Figure 4. **Reduced-form conditional variances**

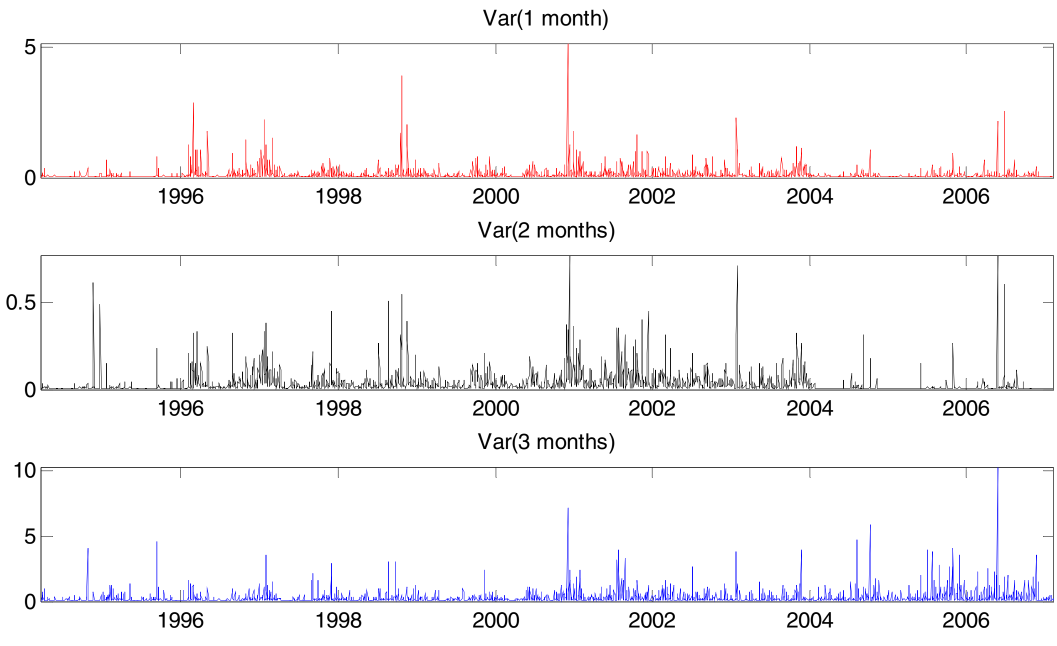


Figure 5. **Reduced-form conditional correlations**

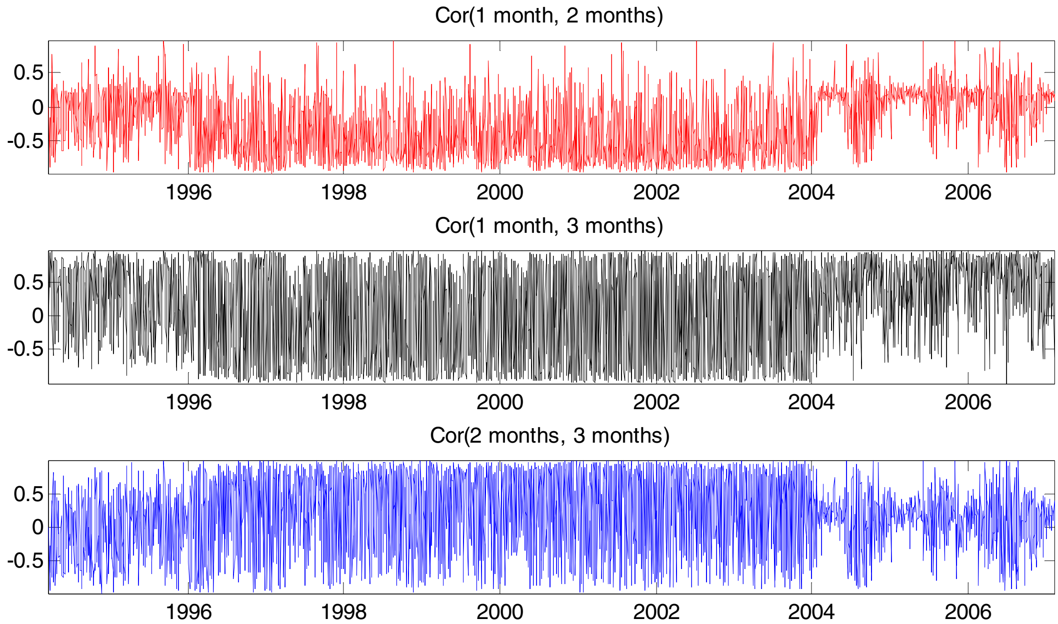


Figure 6. **VIRFs for the second Gulf-war shock**

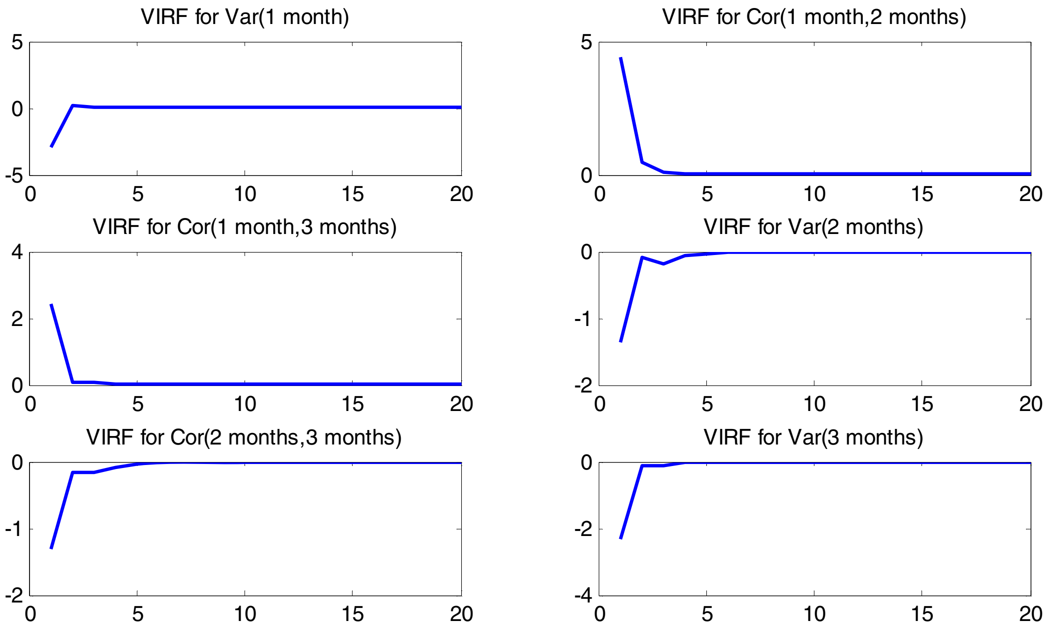


Figure 7. **1st, 10th and 25th percentiles of the VIRF distribution**

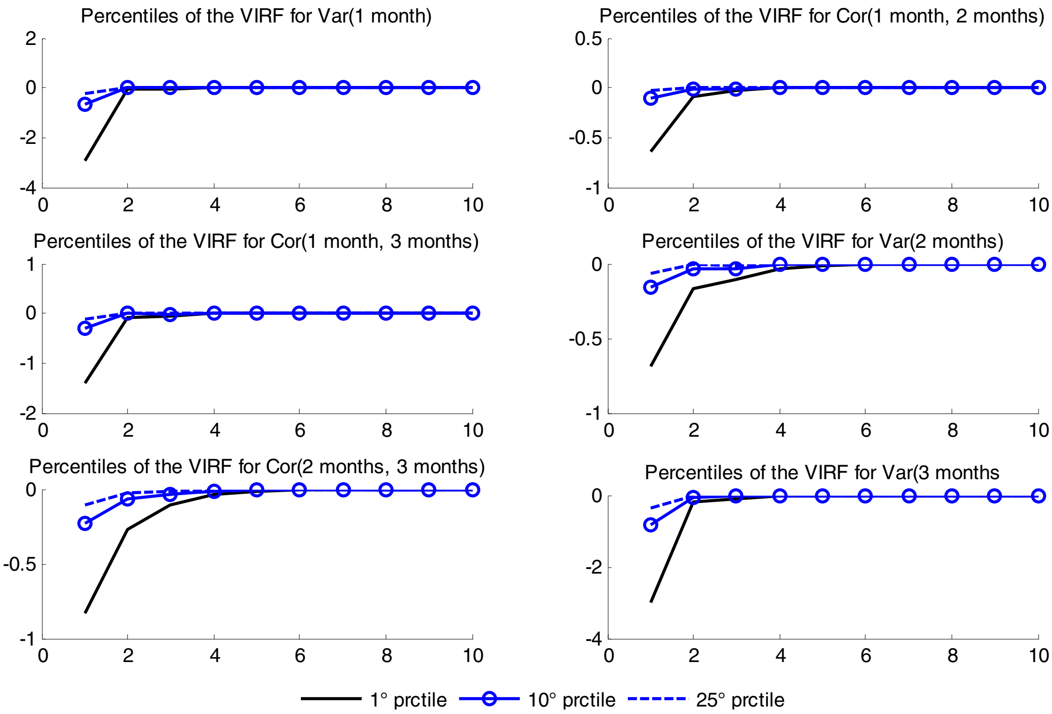
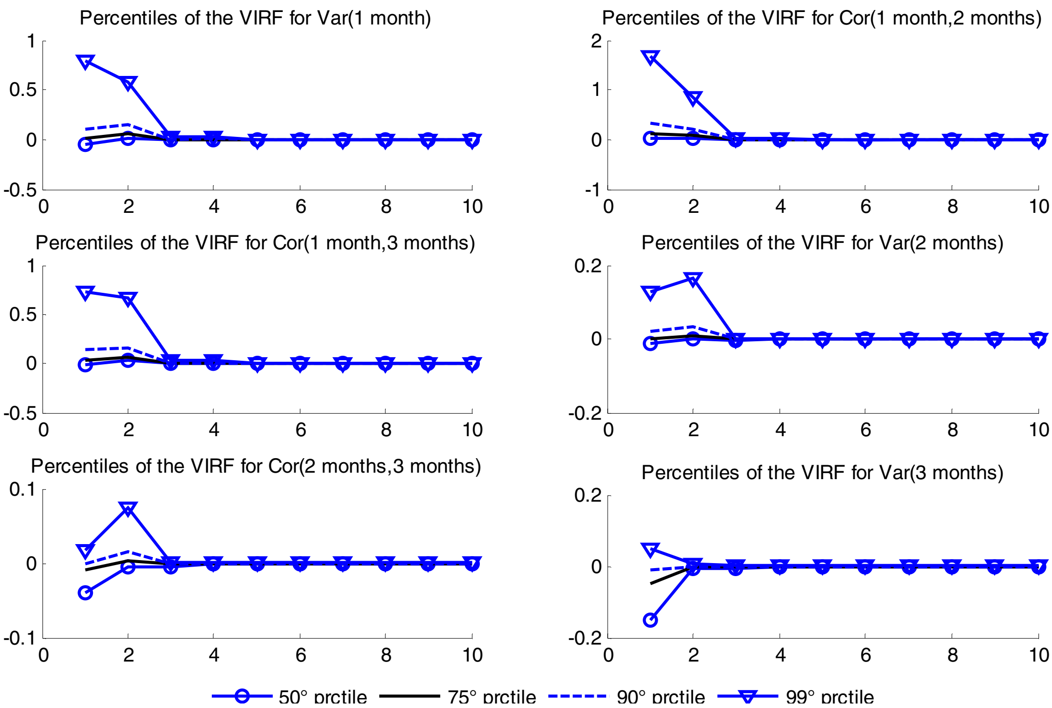


Figure 8. **50th, 75th, 90th and 99th percentiles of the VIRF distribution**



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