Macroeconomic and Financial Effects of Oil Price Shocks: Evidence for the Euro Area*

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Abstract

The paper investigates the macroeconomic and financial effects of oil prices shocks in the euro area since its creation in 1999, with a special focus on the recent slump. The analysis is carried out episode by episode, within a time-varying parameter framework, consistent with the view that "not all the oil price shocks are alike", yet without imposing any a priori identification assumption. We find evidence of recessionary effects triggered not only by oil price hikes, but also by oil price slumps in some cases, likewise for the most recent episode, which is also rising deflation risk and financial distress. In addition through uncertainty effects, the current slump might then depress aggregate demand by increasing the real interest rate, as ECB monetary policy is already conducted at the zero lower bound. The increase in real money balances following the slump points to the accommodation of the shock by the ECB, concurrent with the implementation of the Quantitative Easing policy ($Q.E.$). In so far as $Q.E.$ failed to generate inflationary expectations within the expected environment of soft oil prices, the case for a more expansionary use of fiscal policy than in the past might become compelling, in order to counteract deflationary and recessionary threats.

Keywords: oil price shocks, oil price-macroeconomy relationship, risk factors, semiparametric dynamic conditional correlation model, time-varying parameter models.
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1 Introduction

Crude oil price dynamics since mid-2000s have surely raised new interest on the oil price-macroeconomy relationship, particularly the 2008 boom-bust episode, comparable in real terms with the early 1980s shock (US$ 140 July 2008; US$ 40 December 2008). The oil price has persisted at rather high levels for about five years thereafter (90 US$), until the recent oil price slump, which has lead to a swift 50% oil price contraction since June/October 2014 (40 US$).

Despite the ensuing (potentially) sizable real effects for oil importing countries, we are unaware of any empirical assessment carried out using post-2009 data for the euro area (EA). The latter issue is particularly relevant since the EA has so far only partially recovered from the subprime financial crisis, newly falling in recession in July 2011 - through February 2013-, as the sovereign debt crisis deepened. Hence, the unfavorable and long lasting oil price developments since late 2009 might have contributed to its scattered and sluggish recovery. It is also unclear whether the recent oil price slump will enhance economic growth in the EA. On the one hand, if sufficiently sustained, it might be expected to support economic recovery by stimulating both aggregate demand and supply, i.e., through reducing energy bills and energy input costs and increasing total factor productivity, and through monetary policy accommodation also (Mohaddes and Pesaran, 2016; see also Morana, 2013b). On the other hand, by occurring in an environment of weak economic growth, high deflation risk, and where the policy interest rate is already at the zero lower bound, the oil price slump might also exercise recessionary effects through deepening deflationary dynamics, i.e. by increasing real interest rates and macroeconomic and financial uncertainty.[1]

In this respect, the available evidence is rather thin and, by neglecting recent economic developments, might not yield accurate guidance. For instance, Jiménez-Rodríguez and Sánchez (2005) estimate a small-scale structural vector autoregressive model (SVAR) for the euro area economy over the period 1972 through 2001. Consistent with the "reallocation effect" (Hamilton, 2011), they point to a non-linear impact of oil price shocks on real GDP, as oil price increases lead to stagflation, while oil price declines do not have a statistically significant impact. Peersman and van Robays (2009) also estimate a small-scale constant parameter SVAR model for the euro area over the period 1986 through 2008. They find that the macroeconomic effects of oil price shocks crucially depend on their source, i.e., on whether they are oil supply disturbances, shocks to flow oil demand, or precautionary oil demand shocks. In particular, all of the three types of shocks are inflationary already in the short-run, as well as recessionary at different horizons, flow oil supply and demand shocks in the medium-term only, while precautionary oil demand disturbances also in the short-term. Similar evidence is provided by Forni et al. (2012), who estimate a small open economy DSGE model for the euro area over the period 1995-2007. On the other hand, Hahn and Mestre (2011) estimate a small-scale time-varying parameter SVAR model with stochastic volatility for the euro area, over the period 1970 through 2009. They point to weaker stagflationary effects of both supply and demand oil shocks since the mid-1990s than over the three previous decades (see also ECB, 2010, for similar evidence), yet an unchanged relative contribution of both shocks

[1] The weak and scattered recovery from the sovereign debt crisis and the recent contraction in energy and food prices have put price stability at risk in the EA, inducing deflationary dynamics from December 2014 through March 2015, and then again in September 2015, despite the sizable depreciation of the currency and the implementation of the Quantitative Easing policy ($Q.E.$) since January 2015.

to the determination of oil price fluctuations.

In this paper we further assess the transmission of oil price shocks in the EA. In particular, we employ a new large-scale time-varying parameter framework, based on the semiparametric dynamic conditional correlation model (SP-DCC) of Morana (2015). Consistent with Jiménez-Rodríguez and Sánchez (2005), in order to allow for an asymmetric response of the economy to positive and negative oil price shocks, we separately assess various episodes of persistent price changes. While we do not attempt to categorize the original source of oil price shocks, by evaluating their effects episode-by-episode our analysis is however fully consistent with the view that "not all the oil price shocks are alike", yet with the advantage of not imposing any identification assumption. Moreover, relatively to previous studies, we aim at a thorough understanding of the consequences of oil price shocks, not only concerning their effects on real activity fluctuations and price stability, but also concerning their potential contribution to financial distress and to changes in market expectations about the future economic outlook. Hence, the information set includes, in addition to standard macroeconomic variables, a new financial conditions index for the euro area (Morana, 2016) and the European market, size, value and momentum factors (Fama and French,1993; Charart, 1999). We believe the latter concern is well justified in light of the progressive "financialization" of commodity markets since the early 2000s (Gkanoutas-Leventis and Nesvetailova, 2015) and the sizable contribution of oil market shocks to the determination of risk factor fluctuations (10% to 20% at various horizons; Morana, 2014).

Overall, our findings yield new insights on the macro-financial impact of oil price disturbances for the euro area. For instance, we find strong evidence of asymmetric real effects of oil price shocks, as net oil price increases would have determined a contraction in industrial production over the sample investigated, while net price decreases would have been expansionary only in the early and mid-2000s. Moreover, the latter appear to increase with the magnitude of the shock and the level achieved by the oil price itself. Deflationary dynamics can also be noted following net oil price contractions; in particular, the recent oil price slump might have imparted both a recessionary and deflationary bias and worsened financial conditions, consistent with the observed increase in the real short term rate and uncertainty effects. An increase in real money balances can also be noted, coherent with the deepening of the expansionary stance during 2014, eventually culminated with the introduction of the Quantitative Easing policy ($Q.E$) in January 2015 by the ECB.

The rest of the paper is organized as follows. In Section 2 we introduce the econometric methodology. In Section 3 we present the data, while in Section 4 we discuss the empirical results. Finally, Section 5 concludes.

2 Econometric Methodology

In order to assess the linkage between macro-financial variables and the oil price the semiparametric dynamic conditional correlation model (SP-DCC) of Morana (2015) is employed. The latter model is defined by the following equations

$$\mathbf{y}_t = \boldsymbol{\mu}_t(\boldsymbol{\delta}) + \boldsymbol{\varepsilon}_t \tag{1}$$

$$\boldsymbol{\varepsilon}_t = \mathbf{H}_t^{1/2}(\boldsymbol{\delta})\mathbf{z}_t \tag{2}$$

where $\mathbf{y}_t$ is the $N \times 1$ column vector of the variables of interest, $\boldsymbol{\mu}_t(\boldsymbol{\delta})$ is the $N \times 1$ conditional mean vector $E(\mathbf{y}_t|I_{t-1})$, $\boldsymbol{\delta}$ is a vector of parameters, I_{t-1} is the sigma field; $\mathbf{H}_t(\boldsymbol{\delta})$ is the $N \times N$ conditional variance-covariance matrix $Var(\mathbf{y}_t|I_{t-1})$. Moreover, the random vector $\mathbf{z}_t$ is of dimension $N \times 1$ and assumed to be *i.i.d.* with first two moments $E(\mathbf{z}_t) = \mathbf{0}$ and $Var(\mathbf{z}_t) = \mathbf{I}_N$. Concerning the specification of the conditional variance-covariance matrix $\mathbf{H}_t(\boldsymbol{\delta})$, we assume that the elements along its main diagonal, i.e., the conditional variances $Var(y_{i,t}|I_{t-1}) \equiv h_{i,t}$ follow a GARCH(1,1) process

$$h_{i,t} = \omega_i + \alpha_i \varepsilon_{i,t-1}^2 + \beta_i h_{i,t-1} \qquad i = 1, ..., N \tag{3}$$

subject to the usual restrictions to ensure that the conditional variances are positive almost surely at any point in time.

Concerning the definition of the conditional covariances, a nonparametric specification is posited, grounded on the *polarization* identity

$$Cov(A, B) \equiv \frac{1}{4}\left[Var(A + B) - Var(A - B)\right] \tag{4}$$

given that $Var(A \pm B) = Var(A) + Var(B) \pm 2Cov(A, B)$.

Accordingly, the off-diagonal elements of $\mathbf{H}_t$, $Cov(y_{i,t}, y_{j,t}|I_{t-1}) \equiv h_{ij,t}$, are

$$h_{ij,t} = \frac{1}{4}\left[Var_{t-1}(y_{i,t} + y_{j,t}) - Var_{t-1}(y_{i,t} - y_{j,t})\right] \qquad i, j = 1, ..., N \qquad i \neq j. \tag{5}$$

By defining the transformed variables $y_{ij,t}^{+} \equiv y_{i,t} + y_{j,t}$ and $y_{ij,t}^{-} \equiv y_{i,t} - y_{j,t}$, and assuming a GARCH(1,1) specification for their conditional variance processes $Var_{t-1}(y_{ij,t}^{+}|I_{t-1}) \equiv h_{ij,t}^{+}$ and $Var_{t-1}(y_{ij,t}^{-}|I_{t-1}) \equiv h_{ij,t}^{-}$ as well, we then have

$$h_{ij,t}^{+} = \omega_{ij}^{+} + \alpha_{ij}^{+} \varepsilon_{ij,t-1}^{+2} + \beta_{ij}^{+} h_{ij,t-1}^{+} \qquad i, j = 1, ..., N \qquad i \neq j \tag{6}$$

$$h_{ij,t}^{-} = \omega_{ij}^{-} + \alpha_{ij}^{-} \varepsilon_{ij,t-1}^{-2} + \beta_{ij}^{-} h_{ij,t-1}^{-} \qquad i, j = 1, ..., N \qquad i \neq j \tag{7}$$

where $\varepsilon_{ij,t}^{+} = \varepsilon_{i,t} + \varepsilon_{j,t}$ and $\varepsilon_{ij,t-1}^{-} = \varepsilon_{i,t} - \varepsilon_{j,t}$.

2.1 Estimation of the SP-DCC model

Consistent and asymptotically normal estimation of the SP-DCC model is performed in two steps. Firstly, the conditional variances $h_{i,t}$, $i = 1, ..., N$, i.e., the elements along the main diagonal of $\mathbf{H}_t$, and $h_{ij,t}^{+}$, $h_{ij,t}^{-}$, $i, j = 1, ..., N$, $i \neq j$, are estimated equation by equation by means of QML, using conditional mean residuals; this yields $\hat{h}_{i,t}$, $i = 1, ..., N$, and $\hat{h}_{ij,t}^{+}$, $\hat{h}_{ij,t}^{-}$, $i, j = 1, ..., N$, $i \neq j$. Then, in the second step the off-diagonal elements of $\mathbf{H}_t$, $h_{ij,t}$, $i, j = 1, ..., N$, $i \neq j$, are estimated nonparametrically by computing

$$\hat{h}_{ij,t} = \frac{1}{4}\left[\hat{h}_{ij,t}^{+} - \hat{h}_{ij,t}^{-}\right] \qquad i, j = 1, ..., N \qquad i \neq j. \tag{8}$$

By further defining

$$\hat{D}_t = diag\left(\hat{h}_{1,t}^{1/2}, ..., \hat{h}_{N,t}^{1/2}\right)$$

the conditional correlation matrix R_t is then estimated as

$$\hat{R}_t = \hat{D}_t^{-1} \hat{H}_t \hat{D}_t^{-1}.$$

In order to ensure well behaved conditional covariance and correlation matrices, an ex-post correction can be implemented if needed (see Morana (2015) and the Appendix for further details). Monte Carlo evidence provided in Sbrana and Morana (2016) strongly supports the SP-DCC model, yielding a superior performance, in terms of accuracy of estimated conditional correlations, than standard DCC approaches.

2.2 Displaced conditional covariances and correlations

Conditional cross-covariances at time t displaced s periods between any two variables i,j, $Cov(y_{i,t}, y_{j,t+s}|I_{t-1}) \equiv h_{ij,ts}$, can be computed as

$$h_{ij,ts} = \frac{1}{4}\left[Var_{t-1}(y_{i,t} + y_{j,t+s}) - Var_{t-1}(y_{i,t} - y_{j,t+s})\right] \qquad i,j = 1,...,N \qquad i \neq j,\ s\ \neq 0 \tag{9}$$

and therefore their conditional correlation at time t, displaced s periods, is

$$\rho_{ij,ts} = \frac{h_{ij,ts}}{h_{i,t}^{1/2} h_{j,t+s}^{1/2}} \tag{10}$$

and estimated by

$$\hat{\rho}_{ij,ts} = \frac{\hat{h}_{ij,ts}}{\hat{h}_{i,t}^{1/2} \hat{h}_{j,t+s}^{1/2}} \tag{11}$$

which might also be constrained, if needed, in the range $[-1, 1]$ by applying the sign-preserving bounding transformation in (18).

From (9) the conditional slope parameter $Cov(y_{i,t}, y_{j,t+s}|I_{t-1})/Var(y_{i,t}|I_{t-1}) \equiv \beta_{ij,ts}$ can also be computed as

$$\beta_{ij,ts} = \frac{h_{ij,ts}}{h_{i,t}} \equiv \rho_{ij,ts} \times \frac{h_{j,t+s}^{1/2}}{h_{i,t}^{1/2}} \tag{12}$$

and estimated by

$$\hat{\beta}_{ij,ts} = \frac{\hat{h}_{ij,ts}}{\hat{h}_{i,t}} \equiv \hat{\rho}_{ij,ts} \times \frac{\hat{h}_{j,t+s}^{1/2}}{\hat{h}_{i,t}^{1/2}}. \tag{13}$$

By setting $s = 0$, the special case of the static regression slope parameter $\beta_{ij,t}$ is then obtained, yielding

$$\hat{\beta}_{ij,t} = \frac{\hat{h}_{ij,t}}{\hat{h}_{i,t}} \equiv \hat{\rho}_{ij,t} \frac{\hat{h}_{j,t}^{1/2}}{\hat{h}_{i,t}^{1/2}}. \tag{14}$$

3 The Data

Our information set is monthly and spans the period 1999:1 through 2015:6. In addition to the real WTI oil price return (o; in € per Barrel), a comprehensive set of macroeconomic and financial variables is employed. In particular, real activity is measured by the industrial production growth rate (g); nominal conditions by the harmonized CPI index inflation rate (π); monetary policy stance conditions by the real Eonia interest rate (s) and the rate of growth of real M3 (m). Moreover, the real effective € exchange rate return (e), the EA current account balance (ca) in changes, the return on the real IMF non-energy commodities price index (c) in €, and the Morana (2016) EA financial condition index (fc) are included. The latter subsumes information contained in various interest rate spreads and measures of uncertainty/risk, well tracking the phases of the euro area business and financial cycles.

Revisions in market expectations about the economic outlook have been finally included using the Fama-French (1993) European size (smb), value (hml) and market (mkt) factors, plus the Charart (1997) European momentum factor. Available evidence for the US indeed shows that positive innovations to size and value factors reveal expectations of favorable changes in macroeconomic prospects, while the opposite holds for positive innovations to momentum (Morana, 2014). The rationale is that small firms have limited access to external capital markets and are more vulnerable than large firms to adverse changes in credit conditions. Improved credit and, in general, macroeconomic prospects may then be associated with a rise in the profitability of small stocks, resulting in a higher size factor. Similarly, firms with high book-to-market ratios are likely to suffer more from a higher debt burden and are more vulnerable to adverse changes in monetary policy and interest rates. Improved economic conditions may then be associated with higher profitability of value stocks, resulting in a larger value factor. Moreover, if firms with stronger fundamentals outperform firms with weaker fundamentals during economic downturns and fundamentals are persistent and reflected in stock returns, positive momentum should be observed during recessions. Consistent empirical evidence is provided by Morana (2014).

4 Empirical Results

As described in the methodological section, the estimation of the SP-DCC model is performed using conditional mean residuals. The latter are computed from univariate autoregressive (AR) models with lag length selected according the BIC information criterion and serial correlation tests. GARCH(1,1) models are then estimated using the whitened residuals. Hence, the conditional mean model is specified as

$$\mathbf{y}_t = \boldsymbol{\mu}_t(\boldsymbol{\delta}) + \boldsymbol{\varepsilon}_t \tag{15}$$

$$\boldsymbol{\varepsilon}_t = \mathbf{H}_t^{1/2}(\boldsymbol{\delta})\mathbf{z}_t \tag{16}$$

where $\mathbf{y}_t = \begin{bmatrix} o_t & g_t & \pi_t & e_t & ... & mom_t \end{bmatrix}'$ with $N = 13$; the generic ith element in $\boldsymbol{\mu}_t(\boldsymbol{\delta})$ is $\mu_{i,t}(\boldsymbol{\delta}) = \boldsymbol{\delta}_{i0} + \boldsymbol{\delta}_{i1}y_{i,t-1} + \boldsymbol{\delta}_{i2}y_{i,t-2} + ... + \boldsymbol{\delta}_{ip}y_{i,t-p}$, $i = 1, ..., N$; $\mathbf{z}_t$ is $i.i.d.$ with first two moments $E(\mathbf{z}_t) = \mathbf{0}$ and $Var(\mathbf{z}_t) = \mathbf{I}_N$. OLS estimation of the conditional mean model yields the residuals $\hat{\boldsymbol{\varepsilon}}_t = \mathbf{y}_t - \boldsymbol{\mu}_t(\hat{\boldsymbol{\delta}})$, which are then employed for the estimation of the conditional variance model. The elements along the main diagonal of the conditional variance-covariance matrix $\mathbf{H}_t(\boldsymbol{\delta})$ are posited to follow a GARCH(1,1) specification as in (3), and similarly for each of the 78 $N \times (N-1)/2$ distinct composite processes $h_{ij,t}^{+}$ and $h_{ij,t}^{-}$.

In all cases an IGARCH(1,1) specification was eventually selected and estimated by cross-validation, using the Riskmetrics exponential smoothing form of the model for computational easiness. The results show that all the models are well specified and yield standardized residuals consistent with white noise properties. In Table 1 we report details for the estimated AR-IGARCH models for the original variables included in the vector $\mathbf{y}_t$. For reasons of space, we do not report details for each of 156 IGARCH(1,1) models estimated for the composite variables. Yet, we provide a summary of the findings in Figure 1, where the frequencies and the cumulated frequencies of the estimated IGARCH parameters, as well as the p-value of the test for serial correlation and conditional heteroskedasticity for the standardized residuals are reported. As shown in Table 1 and Figure 1, the estimated lagged conditional variance parameter β falls in the range 0.70-0.95 for the original series, and 0.65-0.95 for the composite variables, with mean value close to 0.9 for both set of series.

Very accurate is also the estimation of the conditional correlation processes. In fact, by comparing the original and (ex-post) transformed conditional correlations[2], we find that the average Theil's U index, across the sample of 78 conditional correlation processes, is just 0.09, with standard deviation equal to 0.05 (not reported). This implies that the original and transformed (well-behaved) correlation processes are very close, i.e., the transformation required to make well-behaved the sequence of conditional correlation matrices leaves largely unchanged their original values, consistent with accurate estimation of second moments delivered by the SP-DCC model.

4.1 Contemporaneous conditional correlations

Given that variables are whitened using the AR filter, their contemporaneous and displaced correlations bear the interpretation of semipartial (or part) correlations. Hence, they yield information on the interdependence between two variables once their own past is controlled for, i.e., they relate the unexpected change in two variables relative to each own past. In Figure 2 we plot the estimated (contemporaneous) conditional correlations relating the WTI real oil price return o with each of the macroeconomic (g, π, e, c, m, s, c, ca) and financial (mkt, smb, hml, mom, fc) variables considered. The latter are denoted as g/o, π/o, e/o, c/o, m/o, s/o, c/o, ca/o, mkt/o, smb/o, hml/o, mom/o, fc/o. Shaded areas correspond to periods of economic recession for the EA economy, as measured by the OECD chronology.[3] Some descriptive statistics are reported in Table 2.

As shown in Figure 2, the estimated conditional correlations are sizable in all cases, particularly concerning industrial production growth and inflation. For inflation the linkage is in general positive, 0.34 on average over the sample investigated, consistent with the available literature (Jiménez-Rodríguez and Sánchez, 2005; Peersman and van Robays, 2009; Forni et al., 2012; Hahn and Mestre, 2011), pointing to a sizable transmission of oil price changes to consumer prices already within the same month.[4] On the other hand, for industrial production the linkage is rather unstable, almost null on average (0.07) as conditional correlations take both sizable positive and negative values over time. The alternating regime feature of g/o can then possibly account for the delayed impact of oil price changes on real economic activity usually found by means of constant parameter specifications.[5] Hence, the transmission of oil price changes to the economy appears to

[2]The parameter in the sign preserving transformation k was set equal to 12, as resulting from the solution of the minimization problem presented in (19). The estimated average positive eigenvalues used for nonlinear shrinkage are 0.0655, 0.0445, 0.0716, 0.1011, 0.2216, 0.3849, 0.5674, 0.7638, 0.8094, 1.4923, 1.6588, 2.9148 and 4.6880. Further details are available upon request from the author.

[3]The timeline is as follows. For the early 2000s recession: February 2001 (start) and July 2003 (end). For the late 2000s (Great) recession: March 2008 (start) and June 2009 (end). For the early 2010s recession: July 2011 (start) and February 2013 (end).

[4]This is consistent with previous evidence pointing to a direct pass-through of oil prices into pre-tax prices of liquid fuels accomplished within two to three weeks (ECB, 2010).

[5]For instance, Peersman and van Robays (2009) and Forni et al. (2012) find a delayed response of

be much faster than usually held.

Moreover, g/o tends to be negative at the beginning of recessions and positive toward their end. This is consistent with the finding that oil price hikes usually lead recessions (Hamilton, 2013).[6] It also suggests that a fall in the oil price at the end of a recession is not expansionary, consistent with asymmetries in the mechanism of job creation and destruction (Hamilton, 2011). Also, for the early and late 2000s recessions the largest positive values for π/o occur at the end and all through the recession, pointing to stagflationary effects of oil price shocks.

Negative is also the average linkage between real oil price dynamics and real effective exchange rate (-0.13), current account (-0.31), real money balances (-0.11) and real short-term rate (-0.31) changes, consistent with a scenario where real oil price hikes lead to *i*) a contemporaneous worsening in the current account due to the increase in the price of imported energy goods; *ii*) a monetary policy accommodation, leading to a reduction in the real short-term rate; *iii*) a reduction in real money balances due to the inflationary effect of the shock (dominating the monetary accommodation); *iv*) a depreciation of the real effective exchange rate, consistent with the monetary policy response to the shock and the concurrent nominal depreciation of the euro currency (see Morana, 2016). Average figures however hide important properties of the data. For instance, since 2006 e/o has been persistently positive, rather than negative, almost through the end of the Great Recession; similarly, ca/o turns positive toward the end of the Great Recession. The same pattern can also be noted toward the end of the investigated sample, in correspondence of the oil price slump (October 2014) and the implementation of $Q.E$ by the ECB since January 2015.

On the other hand, the conditional correlations for commodity prices (0.22) and the financial condition index (0.11) are positive on average. The former finding is consistent with an increase in the real oil price leading to higher real commodity prices through *i*) higher production costs of commodities due to the associated increase in the price of chemical and petroleum derived inputs; *ii*) the expectation of a linkage between energy and commodity markets, as agricultural commodities have been increasingly used to produce energy over time; *iii*) the nominal depreciation of the euro currency; iv) increased financialization of commodity markets leading to a stronger comovement in their prices. Moreover, the latter finding is consistent with higher real oil prices leading to financial stress. Interestingly, both c/o and fc/o turn negative at the end of sample, as well as during the transition to the Great Recession (since 2006 for c/o and mid-2007 for fc/o) and during the crisis itself.

Relevant insights are also yield by the conditional correlations computed for the risk factors. On average, the conditional correlations are positive for the size (0.18) and momentum (0.27) factors, while null for the market (0.05) and value (0.02) factors due to sign compensations over the sample investigated. Similar behavior is shown by smb/o, hml/o and mkt/o turning negative since 2006 during the transition to the crisis, the Great Recession itself and the recent oil price slump; on the other hand, mom/o turns negative only during the Great Recession and the recent oil price slump, i.e., in correspondence of the two largest oil price drops. This is consistent with the expected signalling properties of risk factors (Morana, 2014), since *smb* and *hml* might be expected to increase (de-

real activity to both oil demand and supply shocks (about three years), while a quicker impact is found for precautionary/speculative demand shocks.

[6] Indeed, figures from Hamilton (2013) show that this has been the case in 10 out 11 postwar US recessions to date.

crease) in the expectation and during an expansion (recession); *mom* to increase in the expectation and during a recession. In particular, the positive *mom/o* correlation over most of the sample is coherent with the recessionary bias exercised by oil market shocks since late 1990s (Morana, 2013b).

4.2 Dynamic effects of oil price shocks

In this Section we assess how oil price shocks have actually transmitted to the macroeconomy. While we do not attempt to categorize the original source of oil price shocks, by evaluating their effects episode-by-episode our analysis is, however, fully consistent with the view that "not all the oil price shocks are alike", yet with the advantage of not imposing any a priori identification assumption.

We then start with selecting oil price episodes of interest over the period 1999-2015. In Figure 3 we contrast the real WTI oil price level series with its net changes computed according to Hamilton (1996, 2003). The latter are grounded on the rationale that the real effects of price changes depend on how current oil prices compare with their historical path. High costs of monitoring energy expenditures and frictions with regard to adjusting consumption might account for the reluctance of economic agents to respond to small oil price changes, and therefore only sizable fluctuations might be expected to affect economic activity. In particular, a *net oil price increase* is then defined as the amount by which the log real oil price in month t exceeds its maximum over the previous year (i.e., the last twelve months) and oil price increases less than this benchmark are assumed to have no effect. Similarly, a *net oil price decrease* can be defined as the amount by which the log real oil price in month t falls below its minimum over the previous year and oil price decreases less (in absolute value) than this benchmark do not have effects.

As shown in Figure 3, over the time span considered both sizable net price increases and contractions have occurred. Apart from the early 2000s episode (5.1%), sizable net price increases can be detected near the end of year 2002 and in early 2003 (3.4%; 2002:9-2003:2), concurrent with the oil supply disruptions caused by strikes in Venezuela and the US intervention in Iraq. Yet, global oil supply was not significantly affected by both events (Hamilton, 2013), and oil market supply side conditions, consequently, were not a major determinant of oil price fluctuations in 2003 (Morana, 2012). Moreover, the 2003 oil price shock episode was not long lasting, actually resolved within the same year (-2.4% 2003:4-2003:5), and the ensuing global real effects were also weak (Morana, 2013b).

Then, over the period 2004 through 2006, the real oil price almost doubled, moving from an average value of about 25€ per barrel to an average value of about 50€ per barrel. Five clusters of sizable net price changes can be detected, yielding an overall 18% net price increase over the period 2004:4 trough 2006:7, followed by a -4.3% net oil price contraction in 2006:10-2007:1. The real oil price has kept raising thereafter, peaking at about 86€ per barrel in June 2008, following a 11% net price increase over the period 2007:9 through 2008:6, to swiftly plunge to about 30€ per barrel, within five months, in December 2008. Evidence reported in Morana (2013a) show that the 2004-2008 oil price swing was mostly driven by macro-financial factors, consistent with rapid global economic growth and an upsurge in financial speculation in the oil futures market driving oil demand upward initially (while oil supply was stagnating), and then downward during the contraction, as the Great Recession set in. The overall contribution of the *third oil price shock* to the depth of the Great Recession was however moderate, oil market shocks jointly accounting for about 10% of the contraction in global real activity over the period 2008:2-2008:4 (-1.3% out of -15%).

Since late 2009 a new run-up in real oil prices can be noted, concurrent with the recovery of the global economy from the Great Recession shifting oil demand upward and some episodes of oil supply disruption caused by social and political distress in the Middle-East (Arab spring; IS conflict). Albeit not monotonic, the run-up lasted for about five years, until the reversal in the oil supply-demand imbalance triggered the recent price slump (World Bank, 2015).[7] As shown in the plot, a cluster of six sizable net price changes can then be noted over the period 2009-2015, yielding an overall 12.1% net price increase initially (2009:11 through 2013:8) and then a -13.5% net oil price contraction eventually (2014:10 through 2015:1). From an average value of about 70€ per barrel, persisting over a period of almost four years, a swift 50% plunge to about 40€ per barrel was then scored within four months, from October 2014 through January 2015. Since then, the real oil price has mostly stagnated, lately falling at even lower values (about 30€ per barrel).

In light of the above evidence, five periods of positive net oil price changes and five periods of negative net oil price changes can be selected. Of particular interest for our study are then the last six episodes, covering the sustained mid-2000s real oil price run-up, the 2008 boom-bust episode, and the post-Great Recession dynamics through the recent oil price slump.

4.2.1 Conditional correlations and the oil price-macroeconomy relationship

From (12) the sequence of time-varying conditional slope parameters (multipliers) can then be computed as

$$\hat{\beta}_{oj,ts} = \frac{\hat{h}_{oj,ts}}{\hat{h}_{o,t}} \equiv \hat{\rho}_{oj,ts} \times \frac{\hat{h}_{j,t+s}^{1/2}}{\hat{h}_{o,t}^{1/2}} \quad t, s = 0, 1, 2, ...M \tag{17}$$

allowing to trace the dynamic response of the selected macroeconomic and financial variables ($j = g$, π, s,m, e, ca, c, fc, smb, hml, mkt, mom) to unexpected changes in the real oil price (o).

In our application median dynamic responses of the various macroeconomic variables for each episode of interest are computed, which we scale by the size of the corresponding median net price change in order to make them comparable across episodes. Moreover, $M = 12$, i.e., the (cumulated) response of the various macroeconomic and financial variables is computed from one-month up to the following twelve months. Due to sample size limitations, the analysis for the 2014 oil price slump episode is carried out using a smaller forecasting horizon ($M = 6$). The measure of uncertainty for the dynamic responses at any forecast horizon k is computed by the standard deviation of the responses for horizon k, generated across each set of five episodes of interest. Results are reported in Figures 4-9.

[7] Since 2011, U.S. shale oil production has persistently surprised on the upside. Moreover, expectations of global oil demand have been revised downward several times. The failure to agree on production cuts in November 2014 and maintaining unchanged production levels in the face of increased US production, then suggests a change in the OPEC cartel's policy objective, from targeting an oil price band to maintaining market share. Moreover, since the second half of 2014 the U.S. dollar has sizably appreciated against major currencies, possibly contributing to the oil price decrese by depressing its demand in countries that have experienced an erosion in the purchasing power of their currency. See also Baumeister and Kilian (2015) for further evidence on the role of excess supply and oil market specific factors.

Industrial production and inflation As shown in Figure 4, net oil price increases would have determined a decrease in industrial production in all of the episodes investigated. The median contraction is also fairly similar across episodes, in the range -0.1% to -0.3%, within 1 year, apart from the 2008 oil price boom, which has triggered a stronger contraction (-0.65%). On the contrary, net price decreases would have exercised some expansionary effect only in the early 2000s and mid-2000s: a 1% net oil price contraction would have triggered a 0.1% industrial production increase, within 1 year, in the latter cases; a median negative short-lived response can be noted for the other episodes (up to -0.05%), including the current oil price slump, consistent with uncertainty and deflationary effects (see below).

The asymmetric response to positive and negative oil price shocks is consistent with previous evidence of Jiménez-Rodríguez and Sánchez (2005); however, we also find that the contraction in industrial production deepens with the magnitude of the shock and the level achieved by the oil price itself, having been twice as large during the third oil price shock than for any of the other four episodes in the sample. Also, sizable real effects of oil price increases can be noted at much shorter horizons than documented by standard constant parameter specifications. For instance, Peersman and van Robays (2009) and Forni et al. (2012) find a delayed response of GDP growth to both oil supply and demand shocks (about three years), while a quicker response is found for precautionary/speculative demand shocks. Comparison with their estimates is not straightforward due to the different response variables used, i.e., industrial production rather than GDP, and the unity of measure of the shocks, i.e., *net* oil price changes rather than actual changes. With this caveat in mind, Peersman and van Robays (2009) point to a -0.3% real GDP contraction within four (twenty) quarters following a 10% real oil price hike generated by a precautionary oil demand shock (flow oil supply shock); the response to a global flow oil demand shock of the same magnitude is weaker (-0.10 % within twenty quarters) and even positive within four quarters (0.2%).

Concerning the response of the consumer price index (CPI) level to net oil price increases, a weak inflationary response can be noted at the outset in all cases, followed by some sizable deflation apart from the 2008 oil price boom (0.05% within four months). Moreover, the response of CPI to net oil price contractions is in general (weakly) deflationary at the outset, as well as at longer horizons for the current oil price slump. Concerning comparison with previous available evidence, the same caveat as above applies. In this respect, Peersman and van Robays (2009) point to a 0.1% CPI increase within four quarters following a 10% oil price increase triggered by a precautionary oil demand shock; a 0.3% increase within twelve quarters following either a flow oil demand or supply shock of the same magnitude.

Monetary policy variables Concerning the monetary policy response to oil price shocks (Figure 5), a small real short-term rate contraction can be noted following any of the net oil price increase episodes, pointing to some monetary policy accommodation of their inflationary effects. As previously noted (Figure 4), apart from the 2008 boom episodes, at longer horizons net oil price increases tend to be deflationary; the reduction in the nominal short-term rate, implied by the contraction in the real short-term rate during deflation, is then consistent with monetary policy counteracting the deflation threat. The concurrent contraction in real money balances points to a new money market equilibrium characterized by a lower nominal interest rate and lower real money supply and demand (due to recessionary effects).

Also, the 2008 boom episode would have been fully accommodated, given its inflationary impact and the observed reduction in the real short-term rate; this is also in light of real money balances remaining almost unchanged, implying an increase in the nominal money supply, given the rise in the price index level. Similar interactions can be noted during the mid-2000s oil price run-up.

On the other hand, a small real short-term rate increase, without monetary tightening, can be noted following any of the net oil price decrease episodes. Of particular interest is the ECB response to the 2014 oil price slump. In light of its deflationary effects (Figure 4), the increase in the real short term rate is consistent with an accommodation of the shock; coherently, real money balances increase following the deepening of the expansionary monetary policy stance during 2014, eventually culminated with the introduction of $Q.E.$ in January 2015. As nominal interest rates are currently at the zero lower-bound, the real short term rate increase determined by the deflationary effects of the oil price slump, jointly with uncertainty effects, might account for its (weak) recessionary impact. Hence, in so far as $Q.E.$ failed to generate recovery and inflationary effects, within the expected environment of persistently weak oil prices, the case for a more expansionary use of fiscal policy than in the past might become compelling, in order to counteract the deflationary and recessionary threat to the euro area.

Finally, comparison with available evidence in the literature is not straightforward, as previous contributions are based on data extending through the 1970s, when there was neither a euro area nor a single monetary policy; moreover, they do not consider post-2008 oil price developments. With this caveat in mind, available results point to a somewhat accommodative or neutral response to oil price shocks, as no sizable effects on the real short-term rate are detected by Jiménez-Rodríguez and Sánchez (2005) and little contribution of oil price disturbances to the determination of the shock of the interest rate equation is found by Hahn and Mestre (2011). Similar conclusions can be drawn from Peersman and van Robays (2009), pointing to a nominal short-term interest rate increase following an oil price hike determined by either flow oil supply or global oil demand shocks, while a nominal short-term interest rate contraction would follow a precautionary oil demand shock. Yet, the real interest rate is mostly unchanged in the former two cases, despite the nominal interest rate hike, while it contracts in the latter case.

Extrapolation of the results of Peersman and van Robays (2009) would then imply that the current oil price slump, while being deflationary, would not be recessionary in the short-term, actually being expansionary in the medium-term; we do not regard the latter implications as reliable, not controlling for the fact that ECB monetary policy is currently managed at the zero lower-bound.

Real effective exchange rate, commodity prices and current account As shown in Figure 6, net real oil price changes exercise mostly symmetric effects on the real effective exchange rate. Consistent with previous evidence of Jiménez-Rodríguez and Sánchez (2005), the euro depreciates (appreciates) following a net real oil price increase (decrease) in four out of five episodes. With reference to the recent oil price slump, the anomalous (relatively to the past) real depreciation of the euro is consistent with the concurrent implementation of $Q.E.$

Moreover, a negative correlation with non-energy commodity prices can be noted for almost all price hikes (apart from the 2010s) and contractions (apart from the 2008 bust). The latter evidence is, however, consistent with the recessionary effects imparted by oil price hikes, which would also decrease the demand for non-energy inputs.

Also symmetric is the response of the current account to net oil price changes, worsening (improving) following a net oil price increase (contraction) for all the episodes in the sample (Figure 7); in this respect, particularly sizable is the response of the current account to the third oil price shock (about -2% within 12 months) as well as for the 2002:9-2003:2 episode (-1% within twelve months). Interestingly, the current account improvement associated with the recent oil price slump also yields an indirect measure of its negative aggregate demand impact, as reflected in the associated increase in national savings.

Financial conditions Concerning overall financial conditions (Figure 7), among price hikes, only the 2008 boom episode would have had a destabilizing impact; on the other hand, higher instability can be associated with all the oil price contraction episodes.

Moreover, similarities can be noted concerning the response of risk factors to oil price developments (Figures 8 and 9). In fact, concerning the early 2000s episodes, net oil price increases would have determined a persistent decline in all risk factors (apart from *hml* for the 2000:8-2000:11 episode).[8] Moreover, little response is shown by all risk factors to the mid-2000s oil price increase, while very sizable is their contraction during the third oil price shock. In this respect, only *mom* proves to be resilient to oil price developments, as a positive momentum is expected to lead recessions, as well as to persist at least through their early phase (Morana, 2014).

On the other hand, concerning net oil price decreases, a positive response of *mkt*, *smb* and *hml* can always be noted, apart from the 2008 bust episode for *mkt*. Also, *mom* was negative during the early 2000s oil price contractions, yet positive for the other episodes, including the 2008 oil price bust, signalling financial and economic distress. The positive response of *mom* to the most recent oil price contraction further corroborates the evidence of rising recession and deflation risk, as well as of worsening overall financial conditions.

5 Conclusions

The paper yields new insights on the macroeconomic and financial effects of oil price shocks for the euro area, using an updated sample through 2015, with a special focus on the recent oil price slump. Extrapolation of available results, as for instance Peersman and van Robays (2009), would imply that the current oil price slump, while being deflationary, would not be recessionary in the short-term, actually being expansionary in the medium-term. However, the latter implications might not be reliable, neglecting that ECB monetary policy is currently managed at the zero lower-bound. A further assessment is therefore required in order to gauge insights on the expected consequences of persisting soft oil prices for the euro area.

In the light of the above issue, our analysis is carried out by means of a new time-varying parameter model of the euro area economy, cast within the framework of the semiparametric dynamic conditional correlation model (SP-DCC) of Morana (2015). Within the latter framework, semiparametric estimation of the dynamic multipliers describing the transmission of oil price changes to macroeconomic and financial variables is performed, without imposing a priori restrictions or attempting to categorize the original

[8] See Park and Ratti (2008) for previous evidence on the response of stock prices to oil price shocks for various European economies.

source of oil price shocks; yet, by assessing their effects episode-by-episode, our study is fully consistent with the view that "not all the oil price shocks are alike".

The research is also innovative for the information set employed, including, in addition to the real oil price, industrial production, CPI inflation, the real short-term rate, the real effective exchange rate, real money balances and non energy-commodity prices, a new financial conditions index (Morana, 2016) and four risk factors, i.e., the Fama and French (1993) market, size and value factors and the Charart (1999) momentum factor, computed using European stock market return data. This allows for a thorough understanding of the consequences of oil price shocks, not only concerning their effects on real activity fluctuations and price stability, but also concerning their potential contribution to financial distress and to changes in market expectations about the future economic outlook. We believe the latter concern is well justified in light of the progressive "financialization" of commodity markets since the early 2000s and the sizable contribution of oil market shocks to the determination of risk factor fluctuations (Morana, 2014).

We find strong evidence of asymmetric real effects of oil price shocks for the euro area, as net oil price increases have led to a contraction in industrial production over the whole sample, while net price decreases have yield some expansionary, yet more subdued effects only in the early and mid-2000s. Evidence of recessionary effects of oil price slumps are also detected, likewise for the most recent episode.

Moreover, the real effects of oil price shocks appear to increase with their magnitude and the level achieved by the oil price itself. In this respect, the 2008 boom was surely peculiar for the size of its effects, twice as large than for any other episode since the 2000s; it was also peculiar for its larger inflationary impact, as deflationary rather than inflationary dynamics can be observed in the other cases. In light of its recessionary and deflationary effects, it is then likely that the post-2009 oil price run-up might have contributed to slowing down recovery in the euro area.

Deflationary effects also follow from net oil price contractions. In this respect, the current oil price slump would have imparted both a recessionary and deflationary bias, through higher real interest rates and macroeconomic uncertainty. Ensuing financial distress is also signalled by the financial condition index and momentum risk factor. An increase in real money balances can finally be noted, consistent with the deepening of the expansionary monetary stance during 2014, eventually culminated with the introduction of $Q.E.$ by the ECB. Overall, our findings have a key policy implication. In so far as $Q.E.$ failed to generate recovery and inflationary effects, within the expected environment of soft oil prices, the case for a more expansionary use of fiscal policy than in the past might become compelling, in order to counteract the deflationary and recessionary threats to the euro area.

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6 Appendix: Ex-post correction

The ex-post correction to ensure well behaved conditional covariance and correlation matrices is implemented in two steps. Firstly, the estimated conditional correlations in $\hat{R}_t$, $\hat{\rho}_{ij,t}$, $i \neq j$, are bounded to lie within the range $-1 \leq \hat{\rho}_{ij,t} \leq 1$ by applying the sign-preserving bounding transformation

$$\hat{\rho}^*_{ij,t} = \hat{\rho}_{ij,t}(1 + \hat{\rho}^k_{ij,t})^{-1/k} \tag{18}$$

where $k > 0$ and even, is selected optimally by minimizing the sum of Frobenious norms over the temporal sample

$$\arg\min_k \sum_{t=1}^{T} \left\| \hat{R}_t - \hat{R}^*_t \right\|_F = \arg\min_k \sum_{t=1}^{T} \sqrt{\sum_{i=1}^{N} \sum_{j=1}^{N} \left| \hat{\rho}_{ij,t} - \hat{\rho}^*_{ij,t} \right|^2}. \tag{19}$$

This yields $\hat{R}^*_t$, the transformed correlation matrix, which satisfies, by construction, the Cauchy-Schwarz inequality.

Secondly, positive definiteness is enforced by means of nonlinear shrinkage of the negative eigenvalues of the $\hat{R}^*_t$ matrix toward their corresponding positive average values over the temporal sequence in which they are positive. In practice, the eigenvalue-eigenvector decomposition of the transformed conditional correlation matrix $\hat{R}^*_t$ is performed, yielding

$$\hat{R}^*_t = \hat{E}_t \hat{V}_t \hat{E}'_t$$

where $\hat{V}_t$ is the diagonal matrix containing the ordered original (positive and negative) eigenvalues along the main diagonal and $\hat{E}_t$ is the matrix containing the associated orthogonal eigenvectors. By denoting $\hat{V}^*_t$ the diagonal matrix containing the ordered original and shrank positive eigenvalues, the new estimate of the conditional correlation matrix can be computed as

$$\hat{R}^{**}_t = \hat{E}_t \hat{V}^*_t \hat{E}'_t \tag{20}$$

which, by construction, is well-behaved at each point in time. The implied, well-behaved conditional covariance process at time period t is then

$$\hat{H}^{**}_t = \hat{D}_t \hat{R}^{**}_t \hat{D}_t$$

which obeys the Cauchy-Schwarz inequality and the positive definiteness condition, at each point in time, by construction.

Table 1: Conditional mean and variance models for original series

Panel A: macroeconomic variables								
Variables	o	g	π	e	c	m	s	ca
δ_0	0.5053	0.0199	0.1171	-0.0495	0.1096	0.1388	0.0063	0.1862
	(0.6161)	(0.0699)	(0.0179)	(0.1101)	(0.2061)	(0.0389)	(0.0123)	(0.2539)
δ_1	0.2395		0.2471	0.2244	0.2551		0.3262	-0.5801
	(0.1004)		(0.0821)	(0.0770)	(0.0978)		(0.0786)	(0.0760)
δ_2		0.2376					0.1671	-0.2314
		(0.0825)					(0.0643)	(0.0834)
δ_3		0.3180				0.2294		
		(0.0694)				(0.0615)		
δ_6						0.2664		
						(0.0639)		
α	0.3000	0.1500	0.1000	0.0500	0.1000	0.0500	0.1000	0.0500
β	0.7000	0.8500	0.9000	0.9500	0.9000	0.9500	0.9000	0.9500
	(0.0850)	(0.0590)	(0.0389)	(0.0244)	(0.0328)	(0.0108)	(0.0345)	(0.0289)
Q	0.7536	0.1410	0.0138	0.8374	0.9463	0.0742	0.0139	0.7897
Q_2	0.2597	0.2282	0.5592	0.5618	0.3411	0.9818	0.7015	0.0028
$S\&B$	0.0404	0.2537	0.9745	0.1231	0.7310	0.2751	0.7380	0.2674
BJ	0.0374	0.0018	0.0719	0.0000	0.0001	0.0000	0.3386	0.0000
Panel B: financial variables								
Variables	mkt	smb	hml	mom	fc			
δ_0	0.4451	0.2218	0.2992	0.6619	0.0040			
	(0.3588)	(0.2217)	(0.0888)	(0.3686)	(0.0158)			
δ_1			0.3188	0.3243	0.2153			
			(0.2176)	(0.1058)	(0.0804)			
δ_2				-0.2133				
				(0.0851)				
δ_3				0.1733				
				(0.0774)				
α	0.1000	0.1000	0.0500	0.1000	0.1500			
β	0.9000	0.9000	0.9500	0.9000	0.8500			
	(0.0419)	(0.0326)	(0.0439)	(0.0470)	(0.0257)			
Q	0.9081	0.9100	0.7937	0.5511	0.4597			
Q_2	0.5123	0.1708	0.0760	0.0806	0.9320			
$S\&B$	0.2446	0.8630	0.7960	0.1884	0.0677			
BJ	0.0019	0.5928	0.6269	0.0000	0.0000			

In the Table we report the estimated parameters for the AR-IGARCH models, with standard error in round brackets. We also report the p-value for the Bera-Jarque normality test (BJ), the Box-Ljung test for serial correlation in standardized (Q) and squared standardized (Q_2) residuals up to the 20th order, the joint Engle-Ng sign and size bias test ($S\&B$). The variables are the real oil price return (o), the industrial production growth rate (g), the inflation rate (π), the real effective exchange rate return (e), the non-energy commodity price index return (c), the real money growth rate (m), the real Eonia overnight interest rate (s), the EA current account in changes (ca), the Fama-French Euopean market (mkt), size (smb) and value (hml) factors, Charart momentum factor (mom) for Europe, and the EA financial condition index (fc).

Table 2: Descriptive statistics for contemporaneous conditional correlations.

	g/o	π/o	e/o	ca/o	m/o	s/o	c/o	fc/o	smb/o	mkt/o	hml/o	mom/o
$mean$	0.07	0.34	-0.13	-0.31	-0.11	-0.31	0.22	0.11	0.18	0.05	0.02	0.27
$stdc$	0.34	0.28	0.37	0.37	0.41	0.30	0.30	0.54	0.37	0.31	0.31	0.35

The table reports sample means (*mean*) and standard deviations (*stdc*) for the estimated conditional correlations of the various macroeconomic and financial variables relatively to the real oil price return (*o*). The variables are the industrial production growth rate (g), the inflation rate (π), the real effective exchange rate return (e), the non-energy commodity price index return (c), the real money growth rate (m), the real Eonia overnight interest rate (s), the EA current account in changes (ca), the Fama-French Euopean market (mkt), size (smb) and value (hml) factors, Charart momentum factor (mom) for Europe, and the EA financial condition index (fc).

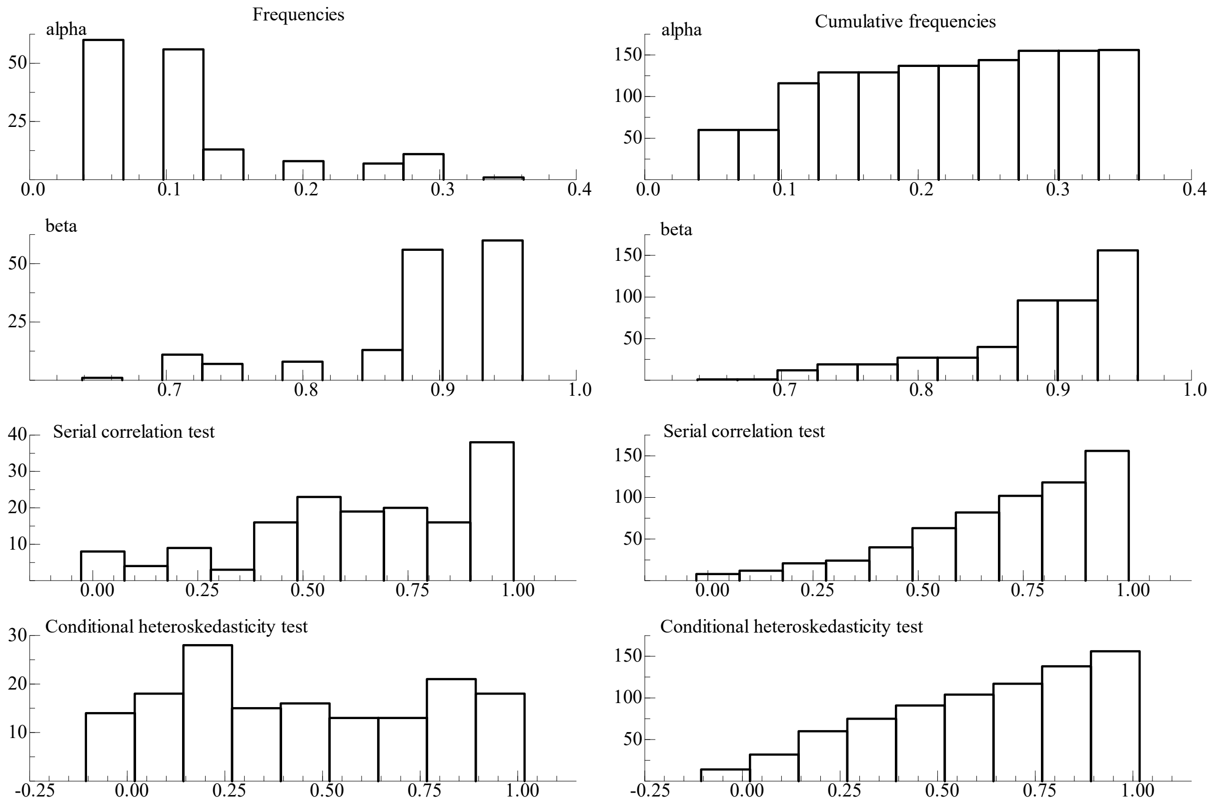

Figure 1: In the figure the cross-sectional distribution for the IGARCH (1,1) estimated parameters for the composite variables is reported. In particular, α is the squared lagged innovation parameter and β is the lagged conditional variance parameter. In the plot, also the cross-sectional distribution of the p-value of the Box-Ljung test for serial correlation in the standardized and squared standardized residuals (up to the 20th order) is reported. For all the statistics right-hand side plots refer to frequencies and left-hand side plots to cumulative frequencies.

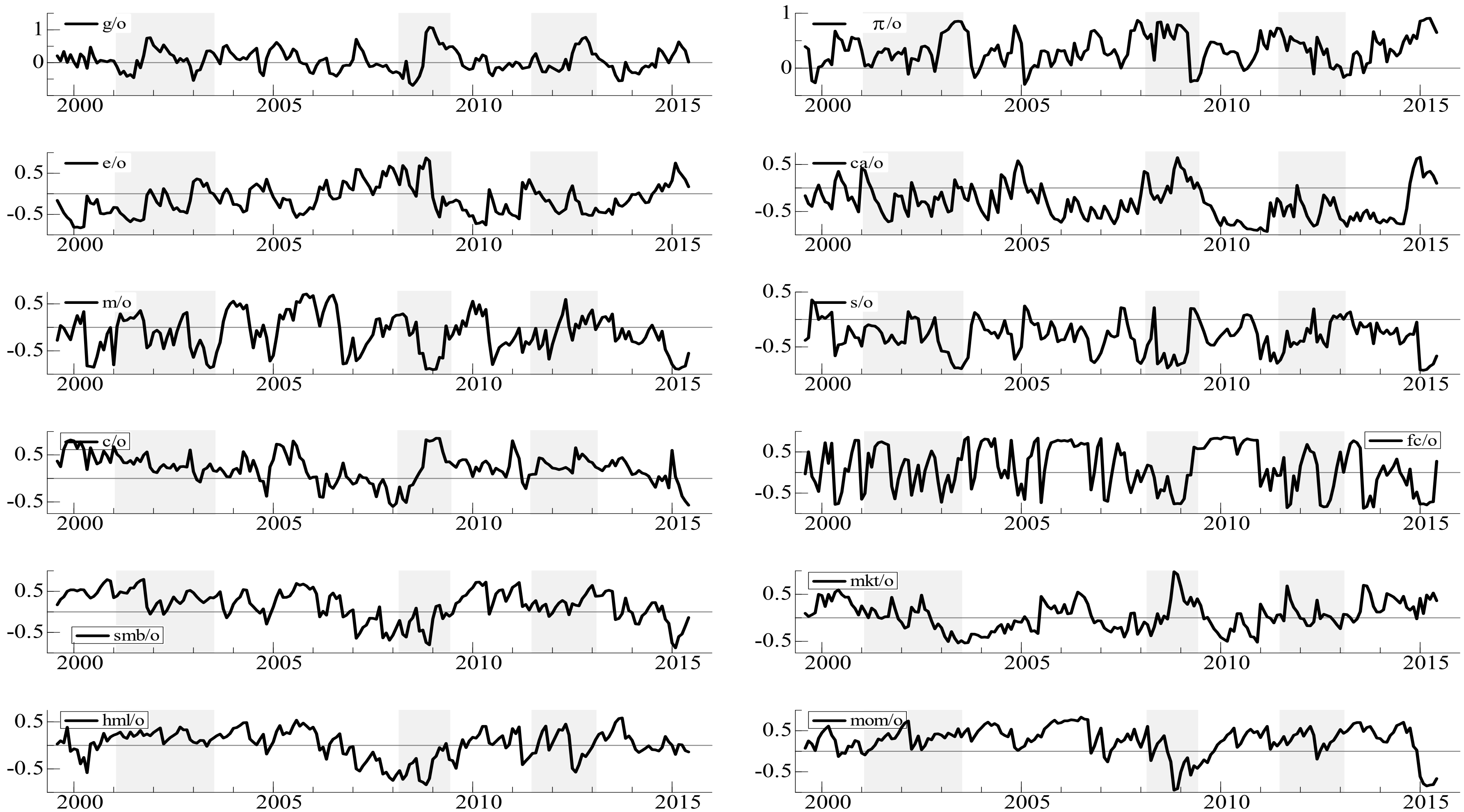

Figure 2: The figure shows the estimated conditional correlations for the various macro-financial variables relatively to the real oil price (o). The variables are: industrial production growth (g), CPI inflation (π), real Eonia interest rate (s), real M3 growth (m), real effective € exchange rate return (e), EA current account balance (ca) in changes, IMF real non-energy commodities price index return (c) in €, the Morana (2015a) EA financial condition index (fc), the Fama-French (1993, 2015) size (smb), value (hml) and market (mkt) factor returns for Europe, plus Charart (1997) momentum factor return (mom) for Europe. Shaded areas refer to recession periods for the EA . The timing is: February 2001 (start) and July 2003 (end); January 2008 (start) and June 2009 (end); July 2011 (start) and February 2013 (end).

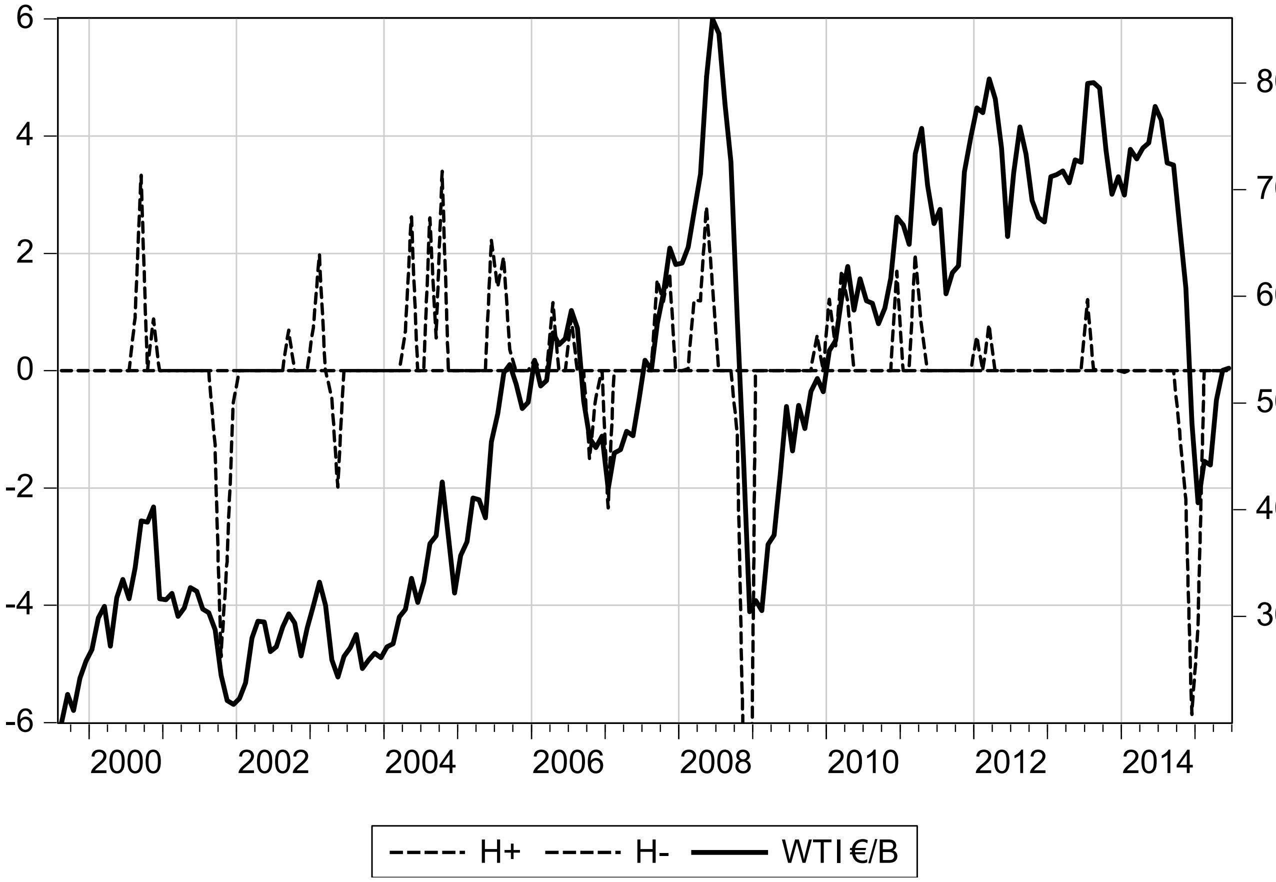

Figure 3: Real oil price (€/barrel) and Hamilton shock (net price increase H+ and net price decrease H-).

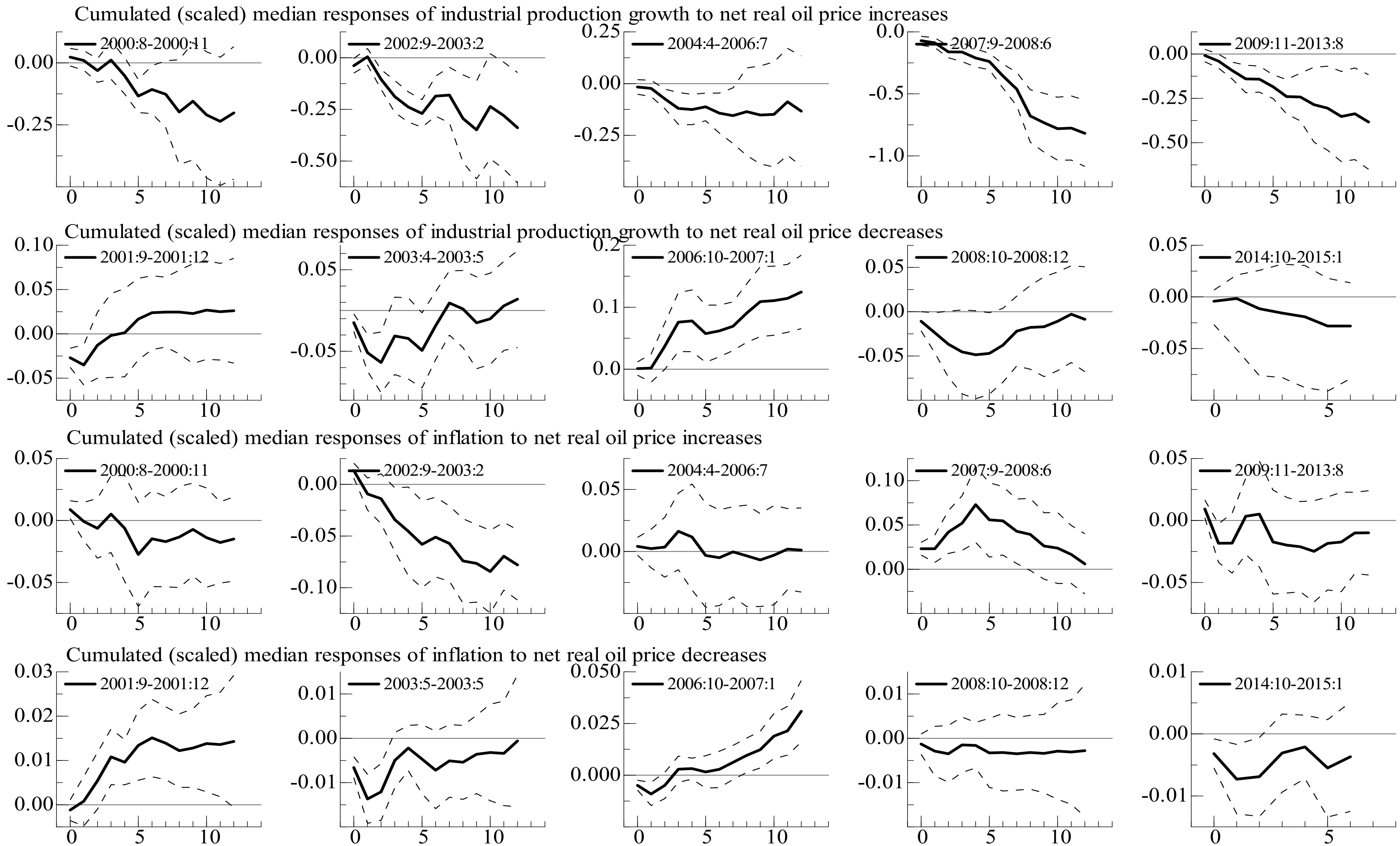

Figure 4: The Figure shows the cumulated (scaled) median response of industrial production growth and inflation to median net real oil price changes for the various episodes of interest.

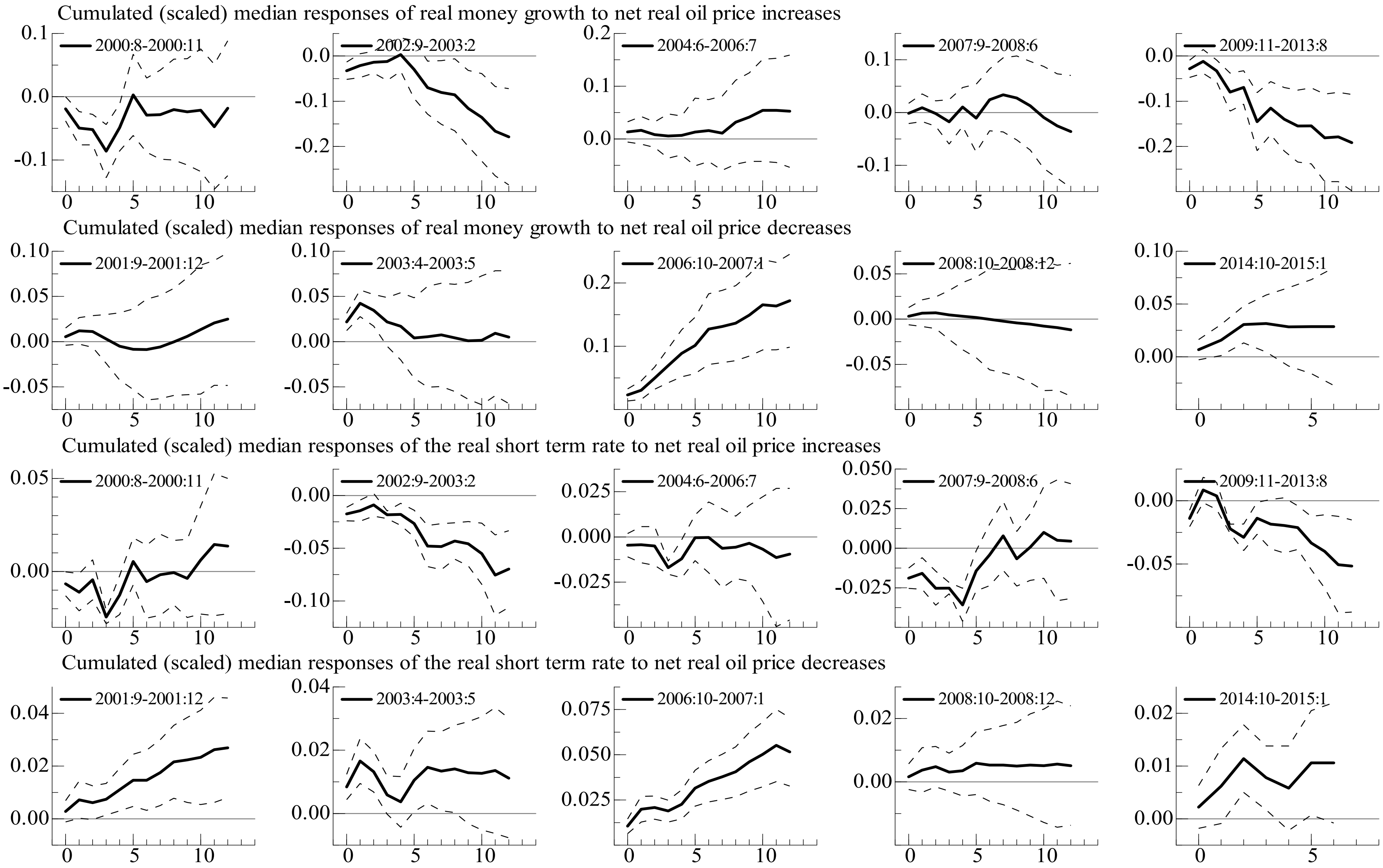

Figure 5: The Figure shows the cumulated (scaled) median response of real money growth and the real short term rate to median net real oil price changes for the various episodes of interest.

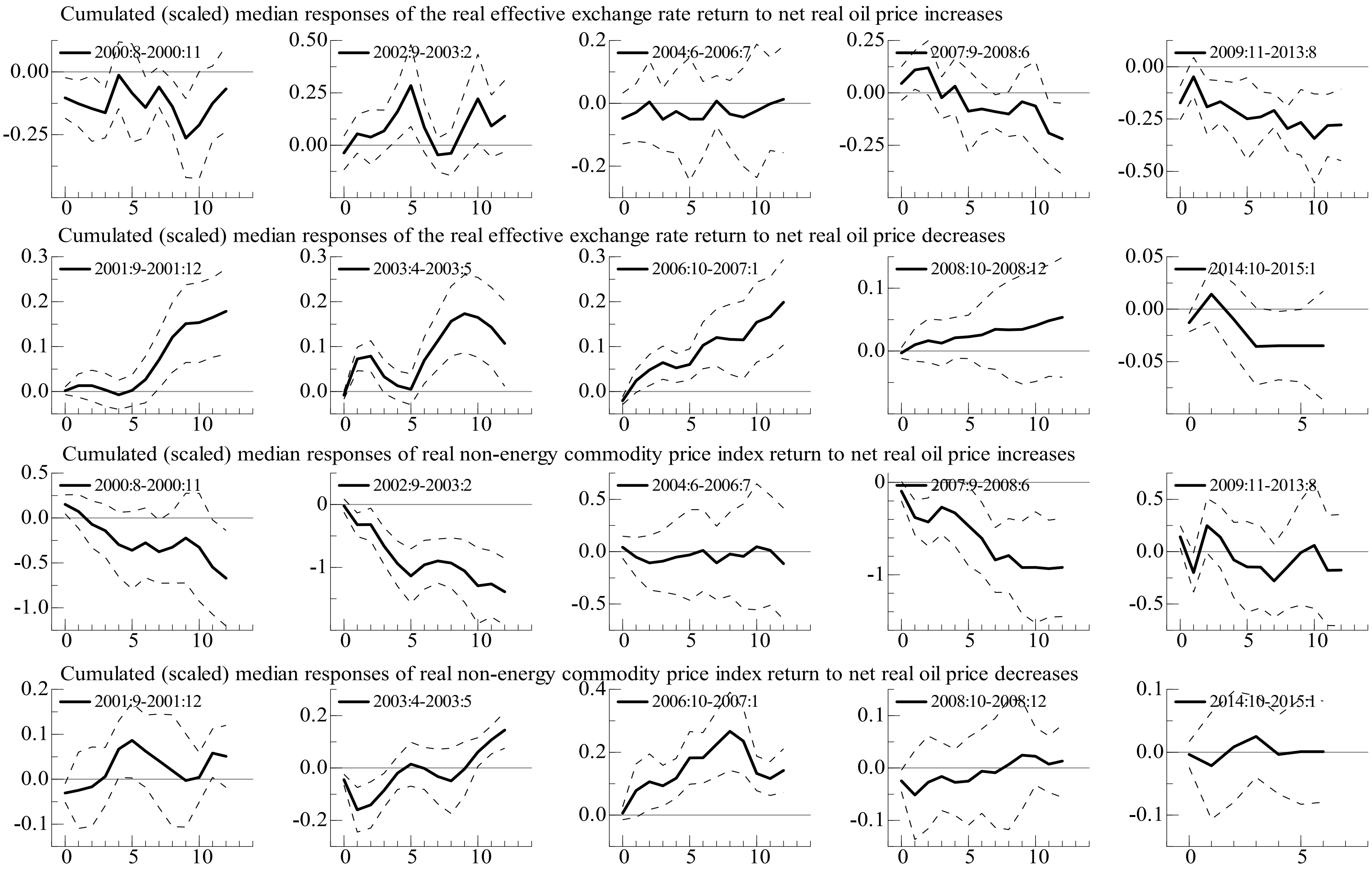

Figure 6: The Figure shows the cumulated (scaled) median response of the real effective exchange rate return and the real non-energy commodity price index return to median net real oil price changes for the various episodes of interest.

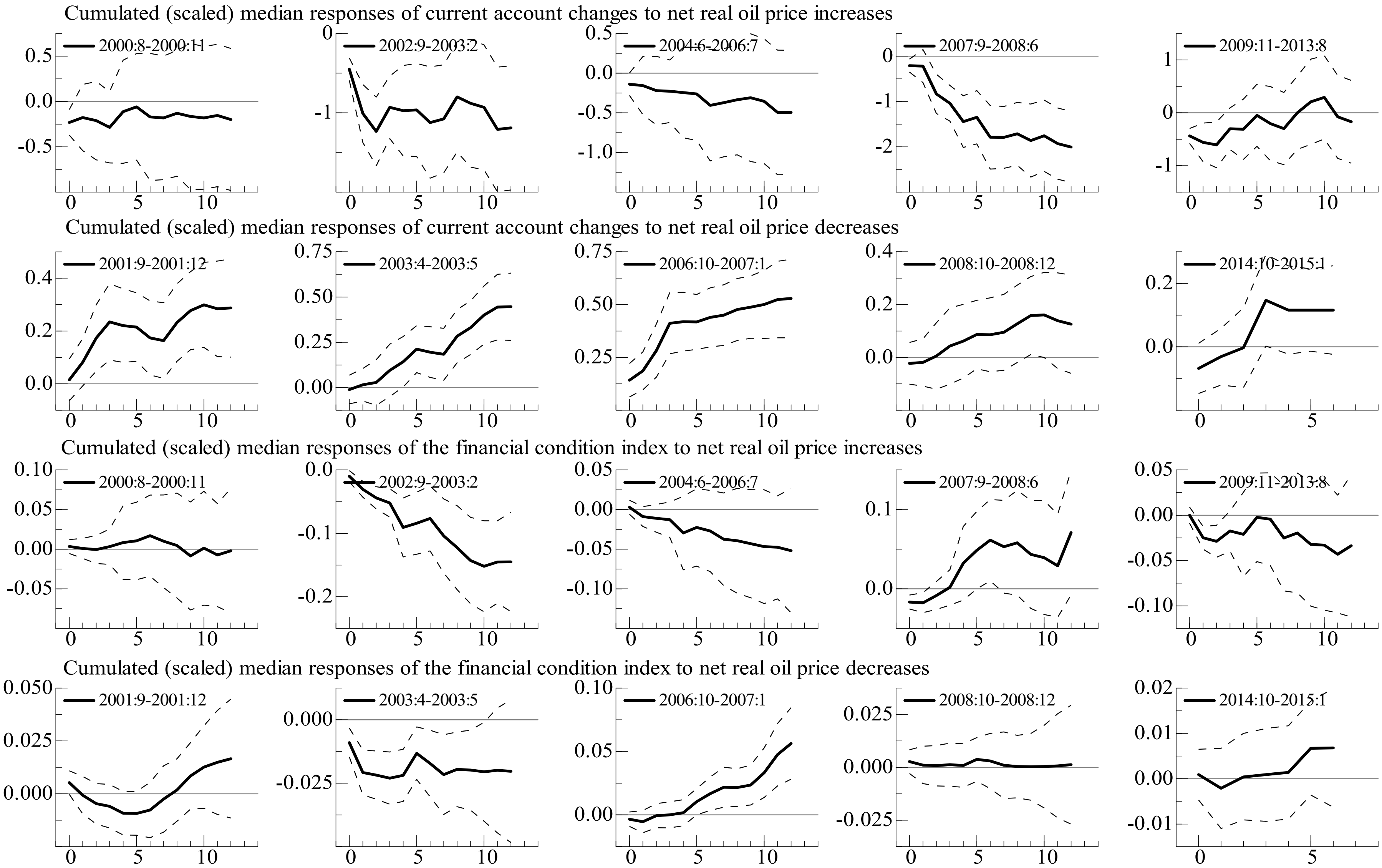

Figure 7: The Figure shows the cumulated (scaled) median response of the current account and the financial condition index to median net real oil price changes for the various episodes of interest.

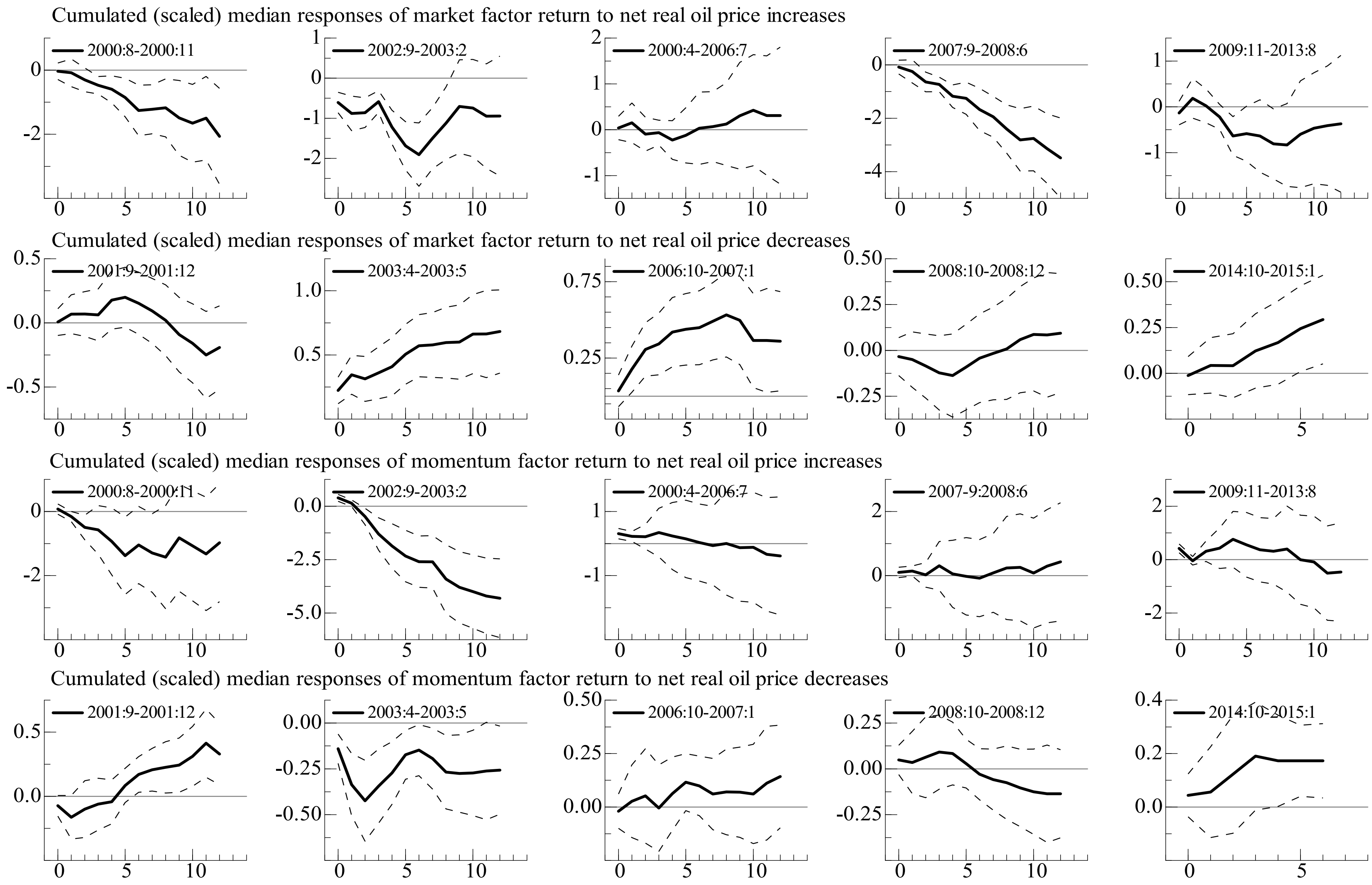

Figure 8: The Figure shows the cumulated (scaled) median response of the market and momentum factor return to median net real oil price changes for the various episodes of interest.

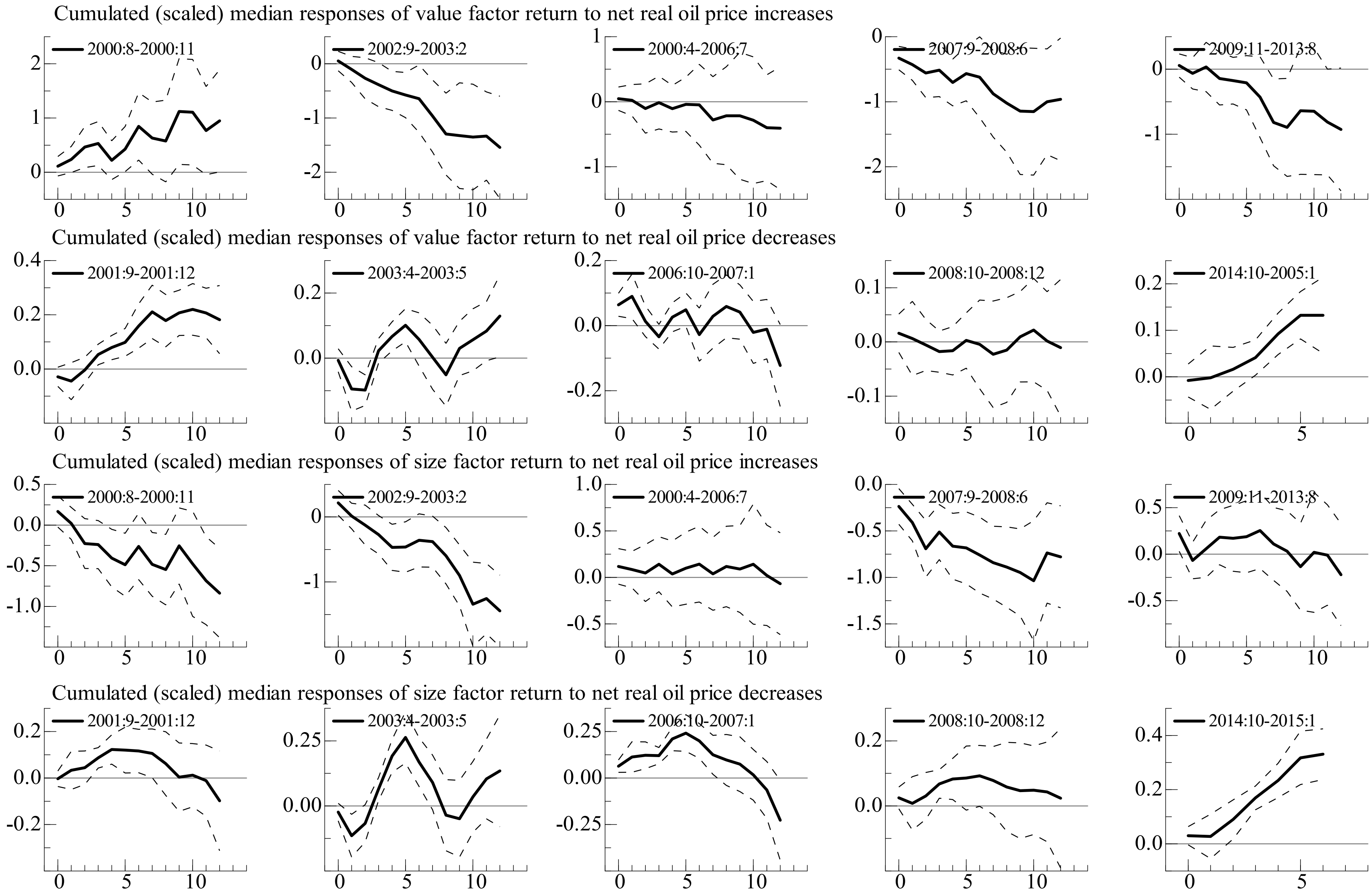

Figure 9: The Figure shows the cumulated (scaled) median response of the value and size factor return to median net real oil price changes for the various episodes of interest.