BANCA D'ITALIA
EUROSISTEMA

Temi di Discussione

(Working Papers)

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March 2010

Number 751

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DOWN THE NON-LINEAR ROAD FROM OIL TO CONSUMER ENERGY PRICES: NO MUCH ASYMMETRY ALONG THE WAY

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Abstract

In the past decade changes in oil prices have played a significant role in shaping inflation dynamics in the US and in the euro area, largely through their direct effect on fuel prices. This has revived the controversy over whether the prices of petroleum products respond more promptly to positive oil price shocks than to negative ones.

This paper provides fresh evidence on this issue for the US, the euro area and the four largest euro-area countries (Germany, France, Italy and Spain), for both petrol and diesel prices. Inference is based on the dynamic response of downstream prices to upstream shocks – rather than on tests on the regression slopes as in the majority of existing studies – and takes into account the non-linearity of the impulse response function in models with asymmetric adjustment, so far ignored in the literature. The empirical analysis shows that fuel prices respond very promptly to oil price shocks, with some heterogeneity across countries, and that no systematic evidence of asymmetries emerges. This result is robust across periods of high and low oil price volatility and holds both for standard and large shocks.

JEL Classification : C52, Q43, E31.
Keywords: energy, oil prices, asymmetry, inflation.

Contents

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1 Introduction

The debate over whether fuels consumer prices adjust asymmetrically to oil price shocks started in the 90s when the first Gulf war boosted crude prices volatility and consumer groups in the US complained of gasoline price downward stickiness (Karrenbrock, 1991). A rich literature followed (labeled “rockets and feathers”, after the definition given by Bacon, 1991, to the asymmetric response of gasoline prices), including applied work and theoretical models aiming at rationalizing the optimality of asymmetric price adjustments. The rapid increase of the price of oil since 2005 has revitalized this controversy both in the public opinion and in policy environments.[1]

The relationship between motor fuels and oil prices has always attracted a lot of attention from market regulators as asymmetries in the pass-through from upstream to downstream prices could result from collusive behavior. The US Congress, for example, requested the General Accounting Office (henceforth GAO) to report on this issue after the early 90s Persian Gulf Crisis, partly to respond to concerns from independent gasoline dealers complaining of predatory pricing by major refineries. In 2007 an investigation by the Italian antitrust authority (the *Autorità Garante della Concorrenza e del Mercato*, henceforth AGCM) verified the existence of collusive price setting behavior by the eight large oil companies that supply liquid fuels in Italy.[2] More recently, in March 2009, the European Commission began an inquiry on the relationship between wholesale and pump fuel prices following pressure from the Federation Internationale de l’Automobile (FIA) concerned that, in some European countries, petrol prices were not responding quickly enough to falling oil prices.[3] The issue is important for consumer welfare as these items constitute around 5% of household expenditure, both in the US and in the euro area.[4]

The behavior of petrol and diesel prices is of great interest also for monetary authorities for a number of reasons. First, according to recent estimates for the euro area (European Central Bank, henceforth ECB, forthcoming) fuels prices account (i) for most of the direct

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[2] See *Provvedimento n. 17754* at *http://www.agcm.it*.

[3] See *http://fiabrussels.com/en/news/archive/commission_takes_fia_worries_seriously.htm* where the FIA press release is posted and also “Brussels to investigate overpricing at petrol pumps”, TimesOnline, March, 5^{th} 2009.

[4] The theoretical literature, however, has long recognized that asymmetries might originate from other factors than collusion among large players. If markups are procyclical, for example, the difference between crude and retail prices might increase in periods of economic expansion and fall in periods of recession (Reagan and Weitzman, 1982). The response of consumers might also cause asymmetries in the adjustment of downstream prices. A price increase might trigger expectations of higher future prices, therefore leading consumers to anticipate their demand and strengthen the price momentum. In the face of a price decrease, on the other hand, consumers might want to wait for future lower prices without further depressing current prices (Brown and Yucel, 2003). Finally, asymmetries in inventory management costs might spill over to retail prices (GAO, 1993, and Borenstein et al., 1997).

effects of oil price shocks on consumer inflation (ii) for around half of the final impact on overall inflation, which includes not only direct but also indirect and second-round effects. Second, energy price fluctuations have a large effect on inflation perceptions as consumers attach a disproportionate weight to the price changes of the items that are most frequently purchased (essentially fuels and food) in forming their perceptions on overall inflation (Ranyard et al., 2008, Del Giovane and Sabbatini, 2008). Finally, fuels prices are ten times as volatile as overall inflation. Fuels price shocks are therefore potentially able to disanchor inflation expectations and set off inflation scares or deflationary traps (Orphanides and Williams, 2003, Evans and Honkapohja, 2009).[5] In this respect any asymmetry in the behavior of fuels prices could potentially impact on inflation perceptions and expectations, with direct consequences for monetary policy.

The "rockets and feathers" issue has received considerable attention in the literature. However, a recent contribution by Kilian and Vigfusson (2009) has shown that all the existing papers on asymmetric effects of energy prices shocks suffer from important methodological limitations which make their results questionable. Their argument can be summarized as follows. In this literature the issue of asymmetries is usually tackled on the basis of regression models in which fuel price changes are regressed against past fuels and oil price increases and decreases. In the majority of the existing studies evidence in favor or against asymmetries is then provided by testing that the coefficients of rises and falls are not significantly different from each other. This approach, however, cannot provide a convincing case either in favor or against asymmetries because, given the presence of censored endogenous variables which makes the model non-linear, in this setup there is no clear correspondence between the slopes of the estimated equations and the dynamic responses to an exogenous shock: small asymmetries in the slopes could be associated with large asymmetries in the responses and vice versa. Furthermore a test on the coefficients does not give any information on the persistence of the responses: if a test on the coefficients detected some asymmetries one would be still left with the question whether these asymmetries quickly fade away or are instead long lasting.

A minority of studies in this literature has analyzed impulse responses. These studies, however, are also flawed as they use the incorrect impulse response function, the so called cumulative adjustment function (henceforth CAF, see Borenstein et al., 1997). In this paper I show that this function can be obtained from the correct non linear IRF by imposing strong restrictions on the dynamics of the models. The consequences of these restrictions are that (i) they (incorrectly) make the response invariant to the size of the shock, so that liquid fuels prices are restricted to respond in the same fashion to small and large shocks and (ii) they shut off the endogenous response of oil prices to their own shocks and therefore understate the effect of an oil price shock onto liquid fuel prices.

In this paper I take Kilian and Vigfusson (2009) critique into account and through the

[5] To substantiate this concern, which might at first sound far-fetched, I plot in Figure 1 the Google Index of intensity of searches of the word "deflation" since 2004 together with the rates of growth over the previous three months of the fuels components of the euro area HICP and of the US CPI. It is suggestive that the peak of the Google Index roughly coincides with the trough in the liquid fuels inflation rates.

lens of non linear impulse responses (Koop et al., 1996) test whether liquid fuels (petrol and diesel) prices respond more to positive than to negative oil price shocks in the US, in the euro area and its largest constituent countries (Germany, France, Italy and Spain). I improve upon the literature in two directions. First, differently from papers based on coefficient testing, I look at the dynamic response of fuels prices fully taking into account the non-linearity of the model. Second, I depart from the simplifying assumption implicit in the use of CAF and study the impact of oil price shocks in a setting in which (i) shocks of different sizes have a different effect on liquid fuels prices and (ii) oil prices are also allowed to respond to their own innovations.

The empirical analysis shows that fuels prices respond very promptly to oil price shocks, generally in less than four weeks, both in the euro area and in the US. Within the euro area the adjustment is faster than the average in Germany and, to a lesser extent, in France, while it is slower in Italy and Spain. In the large majority of cases the null hypothesis of symmetry cannot be rejected, irrespectively of the size of the shock. Some asymmetries can be found, especially in France and the US, in response to large oil shocks (twice the standard deviation of the oil price innovations) but they are confined to the contemporaneous response.

The paper is structured as follows. Section 2 stresses the differences between the present approach and the ones used in the existing literature. Section 3 provides a description of the main properties of the data. Section 4 presents the empirical model, clarifies the assumptions behind the identification of oil price shocks, describes the algorithm for computing impulse responses and discusses how the CAF relates to the correct non linear IRF. Section 5 presents the empirical results. Section 6 concludes.

2 Relation with the existing literature

An exhaustive critical appraisal of the existing evidence on the fuels/oil asymmetry would require a separate paper. Here I draw on two excellent surveys: the former is the report to the US Federal Trade Commission by Geweke (2004), which focuses on the US; the latter is a paper by Manera and Frey (2007), which deals more generally with econometric models of asymmetric price adjustments.

Most of the studies reviewed are based on some variation of an equilibrium correction model in which downstream prices, following an upstream shock, are allowed to adjust asymmetrically towards their long run value, which is usually pinned down by a cointegration relationship. These studies are however, characterized by a large heterogeneity not only in the datasets used, but also in the definition of asymmetry considered. In their review of the literature Manera and Frey (2007) categorize up to six forms of asymmetries in linear models depending on whether the asymmetry is confined to some specific coefficients (the impact response, rather than the adjustment to the equilibrium correction terms or the sum of the dynamic coefficients, etc.). Most of these definitions of asymmetries, however, are based on coefficients testing, and are therefore unsuitable, for the reasons highlighted in the Introduction, to properly address the "the question of whether there is a systematic tendency for downstream prices in the oil well-to-service station gasoline industry to respond to increases in upstream prices more rapidly than downstream prices respond to decreases in upstream prices" (Geweke 2004), a question that can only be answered by looking at impulse response functions. Interestingly, however, a large part of the empirical literature is based on coefficient testing (Bacon, 1991, Karrenbrock, 1991, Duffy-Deno, 1996, Galeotti et al., 2003, Godby et al., 2000, Grasso and Manera, 2007, Meyler, 2009).

Some authors have explicitly looked at impulse response functions (Balke et al., 1998, Bettendorf et al., 2003, Borenstein et al., 1997, Bachmeier and Griffin, 2003, Rodrigues, 2009). They all, however, perform a very restrictive dynamic exercise in which oil prices are changed by one unit and the dynamic response of downstream prices is assessed assuming that (i) there is no further response of the oil price, (ii) there are no further shocks, (iii) the system prior to the shock is in equilibrium. The limitations of this type of exercise are first stressed in Geweke (2004) who notices that "The relevant real world experiment consists of [...] an unending succession of shocks to upstream price. So long as there is no asymmetry in the response of (the endogenous variable) to lagged values of itself this [...] is unimportant. [...] Given asymmetry in response to lagged values of (the endogenous variable) [...] the only practical way to assess any response of downstream to upstream price changes or volatility is through a full dynamic simulation of the model." Kilian and Vigfusson (2009), building on Koop et al. (1996), show how to perform such a full dynamic simulation and correctly compute impulse responses in a model that accommodates asymmetric responses.

An interesting implication of Kilian and Vigfusson (2009) setup is that, since the endogenous variable responds asymmetrically to its own lags, the effect of an oil price shock is non-linear. I exploit this property of the model and investigate the different effects of large oil price surprises compared to standard ones. This issue is not novel in the literature and is already addressed in Godby et al. (2000), Al Gudhea et al. (2007) and Radchenko (2005a). These papers, however, are not immune from the limitations discussed above. In particular Godby et al. (2000) apply a threshold error correction model to weekly data in the Canadian retail markets but base their inference on tests on the regression slopes. Al Gudhea et al. (2007) use impulse responses from a Threshold Vector Error Correction Model, which Kilian and Vigfusson (2009) show to be misspecified. Finally Radchenko (2005a) uses a Markov Switching model to separate permanent from transitory shocks but bases his inference on the cumulative adjustment function.

In terms of findings, the balance seems to be in favor of asymmetries in the US, for which some form of asymmetric adjustment is reported by Karrenbrock (1991), Duffy-Deno (1996), Borenstein et al. (1997), Balke et al. (1998) and, more recently, Radchenko (2005a). Recent evidence for European countries, on the other hand, finds little support for asymmetric adjustment (Meyler, 2009 and Rodrigues, 2009).

3 Data

The information set used in the analysis consists of weekly data from January 1999 to September 2009 of brent, WTI, spot gasoil and gasoline, and retail petrol and diesel prices for the euro area, its largest four countries, Germany, France, Italy and Spain, and the US. I follow the European terminology and use the word petrol for retail gasoline and diesel for retail gasoil throughout the rest of the paper. The terms gasoline and gasoil will instead be used for the refined products at the wholesale (spot) level. Retail prices for the euro area, Germany, France, Italy and Spain are daily data collected once a week from the Eurostat Weekly Oil Bulletin (WOB henceforth). Since they are collected on Monday I match them with the average oil price of the previous week. Brent, WTI, spot gasoil and gasoline, as well as US retail prices, are weekly averages from the Energy Information Administration.[6] I use brent prices for the euro area and WTI prices for the US as a measure of crude oil price. Regarding spot gasoil and gasoline prices I follow Meyler (2009) and use Rotterdam prices for the euro area, while for the US I follow Al Gudhea et al. (2007) and use an average of the prices of the New York, Los Angeles and Gulf prices.

When looking at the pass-trough from crude to retail prices it is customary to break down the pricing chain in various steps. Here I look at two specific relationships: the pass-through from crude to retail prices and the one from spot to retail prices. The wedge between crude and spot prices accounts for refining and shipping costs and margins, while the difference between spot and retail prices includes distribution costs and margins, see Figure 2 (adapted from Meyler, 2009, and ECB, forthcoming).[7] I focus on these two channels specifically as they seem the most relevant for policy makers. The direct relationship between crude and retail prices might be of particular interest for central banks for two reasons. First, the large press coverage that this link receives is likely to influence consumers inflation perceptions and expectations. Second, inflation forecasts, which play an important role for central banks, are usually conditional on a given future path of the price of crude, which is traded in futures markets. Changes in the crude price curve are behind high frequency revisions of inflation forecasts, largely through their direct effect on fuels prices. The pass-through from spot to retail prices, on the other hand, might be of greater concern for market regulators as it reveals more clearly the influence of the distribution market structure on final prices.

Crude, spot and retail prices are plotted in Figure 3 and 4.[8] Each Figure contains two plots. The top one shows data for the euro area and for the four euro area countries together with the price of brent and of the relevant wholesale refined product (gasoline for petrol and gasoil for diesel). The bottom one shows US retail prices together with the price of WTI and of wholesale refined products. Energy prices increased significantly

[6] All data are considered excluding taxes.

[7] This obviously represents a simplification of a process that involves many more steps.

[8] US data are converted in euros for illustrative purposes. In the empirical analysis, estimates for the US are based on data in US$.

in the past ten years, driven by a rising trend in the price of oil. Crude prices peaked in the summer of 2008 and precipitated thereafter in the wake of the recent economic slowdown. Following the abrupt fall in the second half of 2008 they started rising again: at the end of the sample they were around three times higher than in 1999. The existence of a common trend through the various stages of the pricing chain is pretty clear. In the empirical part I test (and can not reject) the hypothesis that upstream and downstream prices are cointegrated. Figure 5 shows the refining and distribution costs and margins, computed, respectively, as the difference between crude and spot and spot and retail prices. Only data for the US and for the euro area are reported, not to fill the plots with too many lines. For petrol prices both wedges have remained broadly stable over the past ten years, both in the US and in the euro area. In the case of diesel, on the other hand, refining margins show a smooth upward trend in the euro area, while they are much more volatile in the US. The behavior of distribution margins is much more homogeneous across geographic areas. They also have somewhat increased since 2005 to return to their historical levels in the most recent months.[9] To get a first idea of the cross country heterogeneity of fuels price developments I compare the unconditional distributions of upstream and downstream price changes across products and countries. These are shown in the histograms in Figures 6 and 7 for crude and spot prices and in Figures 8 to 11 for retail prices. The distribution of crude and spot prices look very similar to each other. Given the positive drift in these prices over the sample period it is not surprising that they are slightly right skewed. Turning to retail prices, one can notice that the shape of the distributions for the US and the euro area are very similar to those of upstream prices, both for petrol and for diesel. Italy and Spain, on the other hand, present a large spike at zero, indicating considerably higher stickiness in the adjustment of liquid fuels prices in these two countries than in the euro area average. Germany, and to a lesser extent France, are the countries where prices appear to be most flexible. Also notice that visual inspection of these histograms does not suggest the existence of stark asymmetries in any of the countries considered.

Finally I perform an analysis of the level of integration of the variables. Standard unit root tests (Table 1) cannot reject the presence of a stochastic trend in the levels, while the presence of a unit root in the first differences is strongly rejected. I therefore proceed by modeling the variables as being I(1) in the levels.

[9]The different behavior of diesel prices across the Atlantic has been the center of a heated debate in the past few years. According to most commentators the temporary drift in euro area margins is due to a number of factors, among which the increasing refining costs imposed by tighter regulation in the euro area on the sulphur content of diesel together with higher demand due to the rising penetration of diesel cars.

4 A structural model of (possibly asymmetric) dynamic adjustment

The pass-through from oil to liquid fuels prices is modeled in a two equations structural model, in which cointegration between oil and liquid fuels prices is entertained and downstream prices are allowed to adjust asymmetrically to upstream shocks. The model is the following:[10]

$$
\begin{aligned}
\Delta p_t^{oil} &= a_{oil} + \sum_{i=1}^{p}(b_{11,i}\Delta p_{t-i}^{oil}) + \sum_{i=1}^{q}(b_{12,i}\Delta p_{t-i}^{fuel}) + \epsilon_{1t} && (1)\\
\Delta p_t^{fuel} &= a_{fuel} + d_{fuel}ecm_{t-1} + \sum_{i=0}^{\bar{p}}(b_{21,i}\Delta p_{t-i}^{oil}) + \sum_{i=1}^{\bar{q}}(b_{22,i}\Delta p_{t-i}^{fuel}) + && (2)\\
&\quad d_{fuel}^{+}ecm_{t-1}^{+} + \sum_{i=0}^{\bar{p}}(b_{21,i}^{+}\Delta p_{t-i}^{oil+}) + \sum_{i=1}^{\bar{q}}(b_{22,i}^{+}\Delta p_{t-i}^{fuel+}) + \epsilon_{2t}
\end{aligned}
$$

where:

1. Variables with an upper + appearing in equation (2) are defined as follows:

$$y_t^{+} = \left\{ \begin{array}{l} y_t \text{ if } y_t > 0 \\ 0 \text{ if } y_t <= 0 \end{array} \right\} \quad (3)$$

   Asymmetric adjustment works through three channels: first, a higher response to a positive, rather than negative, disequilibrium from the long run relationship ecm_t (to be defined below), second, a higher direct impact of past positive oil price changes, third, a stronger correlation with own positive lags. Most of the papers on asymmetric price adjustment have based their analysis on an exclusion test of these regressors from equation (2).

2. The long run term ecm_t is minus the residual of a long run price equation estimated in levels:

$$p_t^{fuel} = \alpha + \phi p_t^{oil} + u_t \quad (4)$$

   The residual u_t represents an equilibrium correction term if it is stationary. The parameters of the cointegration relationship α and ϕ are estimated in a first stage (see Engle and Granger, 1987).[11]

3. ϵ_{1t} and ϵ_{2t} are assumed to be uncorrelated.

[10] The model is estimated in absolute changes, rather than log changes. See Meyler (2009) for a discussion on why it is desirable to do so in periods of volatile oil prices but relatively stable margins.

[11] The error correction term is specified with the minus sign to keep the definition of its censored counterpart ecm_t^{+} consistent with (3). An error correction term in the oil price equation was generally found to have a coefficient not significantly different from zero and therefore excluded.

The crucial aspect of model (1)-(2), which has direct relevance for impulse response analysis, is that an unexpected shock to the change in oil prices (ϵ_{1t}) affects contemporaneously gasoline prices (its impact is $b_{12,0}$ times ϵ_{1t}, $b_{12,0} + b^{+}_{12,0}$ times ϵ_{1t} if $\epsilon_{1t} > 0$), while the gasoline price shock (ϵ_{2t}) feeds back to oil prices only with a lag. This exclusion restriction ($b_{12,0} = 0$), together with the hypothesis that ϵ_{1t} and ϵ_{2t} are uncorrelated, makes the system exactly identified and ordinary least squares a consistent estimator.[12]

4.1 Computing impulse-responses

In a model like (1)-(2) the impulse response function (IRF) at horizon h depends on:

1. the slope parameters of the model (β).
2. the size of the shock (δ).
3. past observations (ω_{t-1}).
4. future shocks ($\epsilon_{1t+1}, \epsilon_{1t+2}, \ldots, \epsilon_{1t+h}$ and $\epsilon_{2t+1}, \epsilon_{2t+2}, \ldots, \epsilon_{2t+h}$)

Since the model does not have a moving average representation, IRF need to be obtained by simulation methods. For a specific value of the slopes β the h steps ahead response to an oil price shock of size δ can be estimated as the difference between two conditional forecasts in the following steps:

1. Define a history of the data ω_{t-1} by picking p and $\bar{p}$ consecutive values of Δp^{oil} and q and $\bar{q}$ consecutive values of Δp^{fuel}.
2. Using model (1)-(2) simulate two paths of the endogenous variables. The first is obtained by hitting the oil equation with a sequence of shocks ($\epsilon^{\mathbf{s}}_{\mathbf{1}} = [\delta, \epsilon_{1t+2}, \ldots, \epsilon_{1t+h}]$) where the sequence ($[\epsilon_{1t+2}, \ldots, \epsilon_{1t+h}]$) is drawn from the empirical distribution and the second equation with a sequence ($\epsilon^{\mathbf{s}}_{\mathbf{2}} = [\epsilon_{2t+1}, \ldots, \epsilon_{2t+h}]$), entirely drawn from their empirical distribution. In the second path the shocks to both equations ($\epsilon^{\mathbf{b}}_{\mathbf{1}}$ and $\epsilon^{\mathbf{b}}_{\mathbf{2}}$) are drawn from their respective empirical distributions. Label the two simulated paths for liquid fuels prices $\Delta p^{fuel}_{h,s}$ and $\Delta p^{fuel}_{h,b}$.
3. Take the difference between the two simulated paths $\Delta p^{fuel}_{h,s}$ and $\Delta p^{fuel}_{h,b}$. This is the response to an oil price shock conditional on a specific history ω_{t-1} and on a specific sequence of future shocks:
$$I^h_y(\beta, \delta, \omega_{t-1}, \epsilon^{\mathbf{s}}_{\mathbf{1}}, \epsilon^{\mathbf{b}}_{\mathbf{1}}, \epsilon^{\mathbf{s}}_{\mathbf{2}}, \epsilon^{\mathbf{b}}_{\mathbf{2}}) \tag{5}$$
4. Repeat steps from 2 to 3 m times and average the responses across m. This amounts to averaging out future shocks and therefore obtaining a response that is conditional only on the parameters (β), the size of the shocks (δ) and past history (ω_{t-1}): $I^h_y(\beta, \omega_{t-1}, \delta)$.

[12]Notice that these identifying assumptions are akin to causal ordering restrictions in a structural VAR.

5. Further average across all histories to obtain the unconditional response: $I_y^h(\beta, \delta)$.

Unconditional IRF in a non-linear setting are theoretically neat but in practice computationally very demanding. To give an idea of the problem consider that researchers usually set $m = 250$ in step 4 and that the number of iterations in step 5 equals the sample size, T. The total number of replications required to obtain one realization of the IRF is then mT. Since in my case T is greater than 500 I would need more than 125000 simulations to obtain a single IRF. If one wanted to get some measure of uncertainty around the IRF by bootstrapping the regression coefficients and computing an IRF for each bootstrap replication the problem would become intractable. To avoid these complications the existing literature has relied on a very restricted form of the correct IRF, the cumulative adjustment function (CAF). The CAF can be obtained directly from the parameters of equation (2) and is defined, in the case of a positive shock as:

$$\begin{aligned} B_h^+ &= B_{h-1}^+ + (d_{fuel} + d_{fuel}^+)(B_{h-1}^+ - \phi) + (b_{21,h} + b_{21,h}^+) + \\ & \sum_{i=1}^{h} b_{22,i}(B_h^+ - B_{h-1}^+) + \sum_{i=1}^{h} b_{22,i}^+ max(0, B_h^+ - B_{h-1}^+) \end{aligned} \tag{6}$$

and the initial condition is $B_0^+ = b_{21,0} + b_{21,0}^+$. In the case of a negative shock the CAF is computed by setting to zero the parameters d_{fuel}^+, $b_{21,i}^+$ and $b_{22,i}^+$ in (6).

It can be seen from (6) that the CAF differs from the true impulse response function in four respects. The CAF is independent from:

1. the size of the shock δ
2. the history of the data
3. future shocks
4. the parameters of equation (1)

To better clarify the relationship between the CAF and the correct IRF it is possible to modify the simulation steps described above and obtain the CAF in three, instead of five, steps:

1. Set the history of the data ω_{t-1} to 0.
2. Using model (1)-(2) simulate two paths of the endogenous variables. The first is obtained by hitting the oil equation with a sequence of shocks ($\epsilon_{\mathbf{1}}^{\mathbf{s}} = [1, \epsilon_{1t+2}, \ldots, \epsilon_{1t+h}]$) where the sequence ($[\epsilon_{1t+2}, \ldots, \epsilon_{1t+h}]$) is such as to make the endogenous responses of the oil price from the second step forward exactly equal to zero. The CAF therefore measures the response to a one off shock to the oil price level of unit size. The shocks to the second equation ($\epsilon_{\mathbf{2}}^{\mathbf{s}} = [\epsilon_{2t+1}, \ldots, \epsilon_{2t+h}]$), are all set to zero. In the second path the shocks to both equations ($\epsilon_{\mathbf{1}}^{\mathbf{b}}$ and $\epsilon_{\mathbf{2}}^{\mathbf{b}}$) are set to zero. Label the two simulated paths for liquid fuels $\Delta p_{h,s}^{fuel}$ and $\Delta p_{h,b}^{fuel}$.

3. Take the difference between the two simulated paths $\Delta p_{h,s}^{fuel}$ and $\Delta p_{h,b}^{fuel}$.

The CAF therefore stands on the other end of the spectrum of tractability with respect to the true IRF. The former can be obtained without any simulation but at the cost of neglecting important sources of uncertainty and the non linear nature of the model, the latter is computationally costly yet it correctly portrays the system dynamics.

In the empirical analysis that follows I work with the non linear IRF but, to make it less computationally demanding, I restrict it along one particular dimension: the dependence on the history of the data. More specifically I modify the first step of the IRF simulation by setting the history of the data ω_{t-1} equal to the sample average of Δp^{oil} and Δp^{fuel} (which I define $\bar{\omega}_{t-1}$). This simplification makes step 5 in the IRF simulation unnecessary, while all the remaining steps are exactly the same. This approximate IRF is therefore conditional on the parameters (β), the size of the shocks (δ) and the average history $I_y^h(\beta, \bar{\omega}_{t-1}, \delta)$. The computational benefit of such an approximation is sizable as the number of simulations needed to get a single IRF drops from around 125000 to 250. The cost in terms of precision, however, needs to be assessed. In the Appendix I present the results of a simulation exercise in which I compare the CAF with the correct and the approximate IRF and show that, while the CAF gives a distorted picture of the response of the endogenous variables to an oil price shock, the approximate IRF that conditions on the mean of the data tracks the true one very accurately.

4.2 IRF asymmetry tests and bootstrap

Once IRF to positive and negative shocks have been computed, a formal test on the presence of asymmetries can be carried out with a Wald test on the cumulated responses up to a specific horizon H:

$$I_y^h(\delta) - I_y^h(-\delta) = 0 \quad for \;\; h = 0, 1, 2, ..., H \tag{7}$$

To carry out this test an estimate of the variance of $I_y^h(\delta) - I_y^h(-\delta)$ is needed. Uncertainty is measured via bootstrap simulation, that is, artificial samples are generated using model (1)-(2), and for each sample the IRF to a positive and negative shock are estimated. The variance of $I_y^h(\delta) - I_y^h(-\delta)$ is then computed using the bootstrapped IRF estimates. Since, as will be discussed in the empirical results sections, the OLS residuals present significant heteroskedasticity I employ a recursive-design wild bootstrap, which Gonçalves and Kilian (2004) show to give consistent parameter estimates in dynamic models in the presence of conditional heteroskedasticity of unknown form. To illustrate the method, in the case of an AR(1) model a pseudo time series $y_t^\star$ is generated as follows:

$$y_t^\star = \hat{\phi} y_{t-1}^\star + \epsilon_t^\star \tag{8}$$

where $\epsilon_t^\star = \hat{\epsilon}_t \eta_t$, η_t is a zero mean, unit variance, i.i.d. sequence, $\hat{\phi}$ and $\hat{\epsilon}_t$ are estimates of the autoregressive parameter and of the regression errors. I set the initial conditions to the true values and draw the sequence η_t from a zero mean, unit variance, normal distribution.[13]

5 Empirical results

In this section I present the main empirical results. Since the number of estimated equations is rather high (6 countries, 2 products and two channels, from crude to retail and from refined to retail, for a total of 48 equations), only a summary of the regression results is reported, together with impulse responses and asymmetry tests.[14]

5.1 From crude to retail prices

I start by using model (1)-(2) to explore the direct relationship from crude to retail (petrol and diesel) prices. Summary results are presented in Tables 2 and 3. These tables are organized in three panels. The top panel shows the parameter estimates of the long run regressions (4) and a unit root test on their residuals that is used to check for cointegration. The middle (bottom) panel reports the salient results of the estimates of the dynamic retail (upstream) price equation. For the dynamic retail price equations, besides the number of lags, the results of a test on the coefficient of adjustment to the long-run equilibrium (ECM drop test in the table, corresponding to the test $d_{fuel} = 0$), of standard tests on the heteroskedasticity and autocorrelation of the residuals and a measure of fit are shown. For the dynamic upstream price equations I report the number of lags, the results of a test on the joint significance of autoregressive terms, and of a test on the significance of feedback terms from downstream prices, residuals diagnostics tests (heteroskedasticity and autocorrelation) and a measure of fit.

Starting from the long run equations in Tables 2 and 3, the pass-through from crude to petrol prices is estimated to be very close to 1 in the euro area, slightly larger for the US. The long-run response of diesel to crude oil prices is larger, of the order of 1.3, most likely owing to the fact that the refining process of diesel is itself very oil intensive. Average equilibrium prices are generally much higher in Italy than elsewhere, both for petrol and for diesel prices, due to a higher long-run constant α, which captures all other margins and costs that do not depend on the price of oil, in the face of a pass-through from oil to retail prices similar to that found in other countries. Unit root tests on the residuals of these equations confirm that there is a cointegration relationship between crude and retail prices as the presence of a stochastic trend is strongly rejected in all cases.

Turning to the dynamic petrol and diesel equations, three results stand out. First, in all the retail price equations the error correction term plays a significant role as the d_{fuel} parameter is always found to be significantly different from zero. Second, retail price

[13]Gonçalves and Kilian(2004) find that this method is robust to the choice of alternative zero mean unit variance distributions from which η_t can be drawn.

[14]Full results are available upon request.

equations residuals show significant heteroskedasticity but no sign of autocorrelation. Third, these simple dynamic equations capture more than 50% of the overall volatility of liquid fuels prices in all countries but Germany where the fit is lower (below 40%). The worse performance of the model in Germany is very likely to be related to the fact that fuels prices are *daily* while oil prices are weekly averages (referring to the previous week). The former are by construction more volatile than the latter as the weekly average washes away transitory shocks. This could lower the model fit in countries where prices change very frequently (like Germany) and provide a better fit in countries where firms adjust their prices less often and are therefore likely to base their decisions on longer lasting variations of costs (like Italy and Spain). Also this might explain the fact that models for the US generally have a better fit since for the US weekly averages of retail prices are used.

The results for the oil price equations show that (i) oil prices are significantly autocorrelated and (ii) in most cases they respond to downstream prices. This result alone would be sufficient to cast doubts on all the empirical experiments on the asymmetry of fuels prices based on tests on the parameters of equation (2) or on the CAF, as they both ignore the endogenous response of upstream shocks to own shocks and the interaction between upstream and downstream prices. Oil price changes, however, are barely predictable on the basis of past own and retail price changes, as generally less than 10% of their variance can be explained by this model.[15] Oil price equation residuals also show significant heteroskedasticity but no autocorrelation.

5.1.1 IRF and Asymmetry tests

The estimated responses of petrol and diesel prices to a one standard deviation oil price shock are presented in Figure 12 and 13, where the median response together with the 15th and the 85th percentiles are plotted. The dash-dotted lines represent the response to a positive shock while the continuous lines are the responses to a negative shock.[16] An oil price shock is completely passed through to petrol and diesel prices in the euro area in 3 to 4 weeks. Within the euro area the fastest adjustment occurs in Germany, where after the first week retail prices have already reached their equilibrium level, and the slowest in Italy and Spain.[17] This cross-country heterogeneity in the IRF is not surprising, given the presence of a large spike at zero in the distribution of fuels price changes in the latter

[15] The low fit of this equation is not by itself worrysome, as it simply indicates that most oil price changes can actually be interpreted as shocks.

[16] The IRF have been rescaled (ex-post) with the new oil price equilibrium after the shock, so as to provide a measure of the effect of a shock on fuels prices relative to the long run effect of the same shock on the price of oil.

[17] Notice that I report quite tight confidence bands in the charts, so the exact timing of completion of the pass-through, that is the week in which the long-run pass-through ϕ is contained within a 95% IRF confidence interval, is shorter. Formal tests for the timing of the complete pass through are not presented because the results of the empirical analysis show that there is very little asymmetry so that the timing of the pass-through would be better estimated using a simpler linear model, as in Meyler (2009).

countries (see Figure 9). The response in the US is slower (yet this could be due to the fact that for the US weekly averages, rather than a single daily price for each week, are used) and uncertainty around the median response is generally higher than elsewhere. Visual inspection does not suggest any asymmetry in the response of petrol and diesel prices in any countries, as the responses to a positive and a negative shock are never disjoint.

A crucial feature of model (1)-(2) is that the response depends non linearly on the size of the shocks. Fuels prices could therefore respond in a different fashion to larger oil price shocks. Figures 14 and 15 show the response of petrol and diesel prices to a two standard deviation oil price shock. In the case of petrol prices the results are very similar to those obtained for a standard sized shock: the persistence of the responses is unchanged and there is no clear separation between the effects of positive and negative shocks. The size of the underlying shock seems to matter more for diesel prices as the difference between the response to positive and negative shocks becomes much clearer especially in France, US and in the euro area.

The statistical significance of such differences is next assessed through the Wald asymmetry tests discussed above. In Table 4 to 7 I report the results of the tests on the IRF in both cases of a standard and of a large shock. Only the results for the contemporaneous, the first and the third week, are reported as the responses quickly approach their common long run values and after the third week the p-values of the tests get very close to 1. Based on these Wald tests the null hypothesis of symmetry is basically never rejected as all test statistics have p-values well above 5%. The only exception, given by the impact (lag 0) response of diesel prices in the euro area in the case of large shocks, should not be overstated since, given how data are constructed for the euro area and for the four euro area countries considered, the contemporaneous response depends on the correlation between retail prices on Monday and the average oil price in the previous week. As retailers presumably need some days to filter out the persistence of oil price changes before setting a new price, the contemporaneous correlation (and therefore the contemporaneous response to shocks) is likely to be noisier than the correlations at longer lags.

5.2 From refined to retail prices

I now turn to a lower stage of the pricing chain, the relationship between spot and retail prices. A summary of regression estimates is presented in Tables 8 and 9, which are organized exactly like Tables 2 and 3. The results are qualitatively very similar to those obtained when using crude prices. First, looking at the long-run equation results (top panel) it can be seen that retail and spot prices share a common trend given the stationarity of the residuals of the long run equations. The long run pass-through of both petrol and diesel prices is now close to unity, in line with findings in Meyler (2009). Second, results on the dynamic retail price equations (middle panels) show that spot prices are strong attractors for downstream prices as the error correction term enters these with a coefficient significantly different from zero. Third, the residuals of both spot and retail price equations are heteroskedastic but do not show any sign of autocorrelation. Fourth, spot gasoline and gasoil prices are autocorrelated and, in most cases, also respond to retail prices. Notice, also, that the goodness of fit of the retail price equations is now much higher, due to the higher proximity of refined and retail products in the production chain, albeit still lower in Germany than elsewhere. Spot prices are nonetheless very noisy, so that most of their weekly variance is left unexplained by these simple equations.

5.2.1 IRF and Asymmetry tests

The dynamic responses of petrol and diesel prices to a standard spot price shock are shown in Figures 16 and 17. The pass-through from refined to retail products is not any faster than that from crude to retail products: the number of weeks needed to reach the new long-run equilibrium following a shock is again around two weeks in the case of Germany and 4/5 weeks in all the other countries. This is not surprising since oil is a storable commodity and inventories of crude and refined oil are quickly and simultaneously priced at the level induced by underlying shocks (GAO, 1993). The stickiness of consumer prices on the other hand, may reflect rigidities specific to the retail sector. Retailers, for example, might invest in market shares by buffering temporary oil price increases through their margins and passing onto consumers only lasting changes in spot prices. Again, based on the visual inspection of the IRF, no stark difference between the responses to positive and negative shocks can be spotted. The response to large shocks (Figures 18 and 19), differs only marginally from those to standard surprises. The results of the Wald asymmetry tests reported in Tables 10 to 13 suggest that, also at this lower stage of the pricing chain, no systematic asymmetric pattern can be found. In a few cases (petrol prices in the US and in France, and diesel prices in the US) the tests reject the null hypothesis of symmetric adjustment at the 5% confidence level. In these three cases, however, asymmetries are extremely short lived and, in two cases out of three, confined to the case of relatively large shocks, which, by definition, have a low probability of occurring. In interpreting the significance of these asymmetries it is useful to keep in mind that previous studies that emphasized the existence of asymmetric adjustment in the prices of petroleum products found differences in the response to unexpected rises and decreases up to the tenth (Borenstein et al., 1994) or even the twentieth (Radchenko, 2005a) week after the shock.

5.3 Robustness checks

The results presented so far constitute a compelling evidence against the “rockets and feathers” effect. I check their robustness along two dimensions. First, I modify the algorithm for computing IRF to accommodate serial correlation in the variance of the shocks. This robustness check is motivated by the presence of significant heteroskedasticity in the estimated errors, as signaled by the results of the tests shown in Tables 2, 3, 8 and 9 which warned of significant ARCH effects in the residuals. While the wild bootstrap procedure that I employ makes parameter estimates robust to residuals heteroskedasticity there could still be a problem with the way future random shocks are drawn when simulating the IRF. As explained in subsection 4.1, the IRF are obtained by hitting each equation with a sequence of shocks, where the first shock to the oil price equation is restricted to be a constant δ and the subsequent ones are drawn from the empirical distribution. Increasing the size of the initial shock δ does not require a modification of the algorithm if errors are i.i.d., the case analyzed by Kilian and Vigfusson (2009). If the variance of the errors, however, depends on the size of past shocks, a larger initial shocks δ has to be followed by a shock drawn from a distribution with higher variance. I therefore estimate model (1)-(2) specifying an ARCH(4) process for the residuals:

$$\epsilon_t = \sigma_t u_t \tag{9}$$

$$\sigma_t^2 = \mu_0 + \sum_{k=1}^{4} \mu_k \epsilon_{t-k}^2 \tag{10}$$

where u_t is a N(0,1) random variable, and then use the estimates of the parameters of equation (10) to simulate the future shocks in the IRF algorithm rather than drawing them from the empirical distribution. Results for the responses to crude oil price shocks are presented in Tables 14 to 17, while the tests on the responses to refined oil price shocks are shown in Tables 18 to 21. It is clear that using this variation of the model yields even stronger results in favor of symmetry, as the Wald tests reject the null hypothesis even more sporadically than in the benchmark cases.

The second robustness check consists of verifying whether the results hold in periods of low and high oil price volatility. The issue of oil price volatility and asymmetries is investigated by Radchenko (2005b), where asymmetries are found to be positively correlated with the standard deviation of oil price shocks. Since in our sample we have a rather clear separation between a first period of relatively low turbulence, which ended more or less in 2005, followed by a period of high volatility, we proceed by repeating the empirical analysis on two equally sized sub-samples. The former includes data between January 1999 and the second week of April 2004, the latter starts in the third week of April 2004 and ends in September 2009. In the second subsample the standard deviation of oil price changes is around three times as high as in the first one. Results, shown in Tables 22 to 29, are overall in line with those for the full sample: some asymmetries show up in response to large oil price shocks in the US and in France in both sub-periods. There are a few more cases in which the tests reject the symmetry hypothesis for Germany and the euro area average. In all these cases, however, asymmetry is found only at the contemporaneous lag.

6 Conclusions

The increased volatility of oil prices in the past few years has revived the interest both by antitrust authorities and central banks on whether petroleum prices rise faster than they fall in response to oil price fluctuations. This paper revisits this issue for the US, the euro area, Italy, Germany, France and Spain adopting a methodology that improves upon the literature in two respects. First, a structural model is used in which upstream and downstream prices are modeled jointly and an oil price shock is identified. This approach proves to be superior to the single equation one, widely used in this literature, given that oil price changes are autocorrelated and also respond to downstream prices. Second, inference is based on formal tests on the dynamic response of downstream prices to shocks to upstream costs rather than on regression coefficients, taking into proper account the non linearity of this response implied by the presence of endogenous censored variables.

The empirical analysis shows that fuels prices respond within 4/5 weeks to an oil price shock both in the US and in the euro area. Some heterogeneity emerges across euro area countries, with Italy and Spain adjusting significantly more slowly than the rest. Despite some (very mild) stickiness in retail prices, no significant evidence of systematic asymmetries in the response of retail (petrol and diesel) prices to upstream (crude and refined) shocks emerges. This result is robust across periods of high and low oil price volatility and holds both for standard and large shocks. Some asymmetries are found in the case of the US and of France. They are, however, sporadic and very short lived.

This paper leaves a number of open questions. In particular integrating the non-linear IRF in a framework that allows for the parameters to change over time or to switch across regimes could offer further insights on the oil/retail price relationship. Even more interestingly the issue of changing oil price volatility could be more carefully modeled. Both issues are left for future research.

References

[1] Al Gudhea, S., Kenc, T., and S. Dibooglu (2007), Do retail gasoline prices rise more readily than they fall? A threshold cointegration approach. Journal of Economics and Business 59, 560-574

[2] Bachmeier, L.J. and J.M. Griffin (2003), New evidence on asymmetric gasoline price responses. The Review of Economics and Statistics 85, 772-776.

[3] Bacon, R.W. (1991), Rockets and feathers: the asymmetric speed of adjustment of UK retail gasoline prices to cost changes. Energy Economics 13, 211-218.

[4] Balke, N.S., Brown, S.P. and M.K. Yucel (1998), Crude oil and gasoline prices: an asymmetric relationship? Federal Reserve Bank of Dallas Economic Review (First Quarter), 2-11.

[5] Bettendorf, L., van der Geest, S., A. and M. Varkevisser (2003), Price asymmetry in the Dutch retail gasoline market, Energy Economics 25, 669-689.

[6] Borenstein, S., Cameron, A.C. and R. Gilbert (1997), Do gasoline prices respond asymmetrically to crude oil price changes? The Quarterly Journal of Economics 112, 305-339.

[7] Brown, P.A., and M. K. Yücel (2003), Gasoline and Crude Oil Prices: Why the Asymmetry? Economic and Financial Policy Review, Federal Reserve Bank of Dallas, Q3, 23-29.

[8] Del Giovane, P. and R. Sabbatini (2008), The Euro, Inflation and Consumers' Perceptions: Lessons from Italy. Berlin- Heidelberg Springer.

[9] Duffy-Deno, K.T. (1996), Retail price asymmetries in local gasoline markets. Energy Economics 18, 81-92.

[10] Engle, R.F. and C.W.J. Granger (1987), Co-Integration and Error Correction: representation, estimation, and testing. Econometrica 55, 251-276.

[11] Evans, G. W. and S. Honkapohja (2009), Expectations, Deflation Traps and Macroeconomic Policy. CEPR Discussion Papers 7397.

[12] European Central Bank, Energy Market and the Macroeconomy, Structural Issues Report, in course of publication.

[13] Galeotti, M., Lanza A. and M. Manera (2003), Rockets and feathers revisited: An international comparison on European gasoline markets. Energy Economics 25, 175-190.

[14] General Accounting Office (1993), Energy security and policy: Analysis of the pricing of crude oil and petroleum products. GAO/RCED-93-17, March, 1993.

[15] Geweke, J. (2004), Issues in the "Rockets and feathers" Gasoline Price Literature. Report to the Federal Trade Commission.

[16] Godby, R., Lintner, A.M., Stengos, T. and B. Wandschneider (2000), Testing for asymmetric pricing in the Canadian retail gasoline market. Energy Economics 22, 251-276.

[17] Gonçalves, S. and L. Kilian (2004), Bootstrapping Autoregressions with Conditional Heteroskedasticity of Unknown Form, Journal of Econometrics, 123, 2004, 89-120.

[18] Grasso, M. and M. Manera (2007), Asymmetric error correction models for the oil-gasoline price relationship. Energy Policy 35, 156-177.

[19] Karrenbrock, J.D. (1991), The behavior of retail gasoline prices: symmetric or not? Federal Reserve Bank of St. Louis, July/August, 19-29.

[20] Kilian, L. and Vigfusson, R. J. (2009), Pitfalls in Estimating Asymmetric Effects of Energy Price Shocks, CEPR Discussion Papers 7284.

[21] Koop, G., Pesaran, M.H and S. Potter (1996), Impulse response analysis in non linear multivariate models. Journal of Econometrics 74, 119-147.

[22] Manera, M. and G. Frey (2007), Econometric models of asymmetric price transmission. Journal of Economic Surveys 21, 259-325.

[23] Meyler A. (2009), The pass through of oil prices into consumer liquid fuels prices in an environment of high and volatile oil prices. Energy Economics, 6, pp. 867-881.

[24] Orphanides, A. and J.C. Williams (2005), Inflation scares and forecast-based monetary policy. Review of Economic Dynamics, 8(2), 498-527.

[25] Radchenko, S. (2005a), Lags in the response of gasoline prices to changes in crude oil prices: the role of short and long-term shocks. Energy Economics 27(4), 573-602.

[26] Radchenko, S. (2005b), Oil Price Volatility and the Asymmetric Response of Gasoline Prices to Oil Price Increases and Decreases. Energy Economics 27(5), 708-730.

[27] Ranyard, R., Del Missier, F., Bonini, N., Duxbury, D., and B. Summers (2008), Perceptions and expectations of price changes and inflation: A review and conceptual framework. Journal of Economic Psychology, 29, 378-400.

[28] Reagan, P, and M. Weitzman (1982), Asymmetries in Price and Quantity Adjustments by the Competitive Firm. Journal of Economic Theory 27, 410-20.

[29] Rodrigues, J., (2009), Asymmetries in the adjustment of motor diesel and gasoline pump prices in Europe. Working Papers 37, Portuguese Competition Authority.

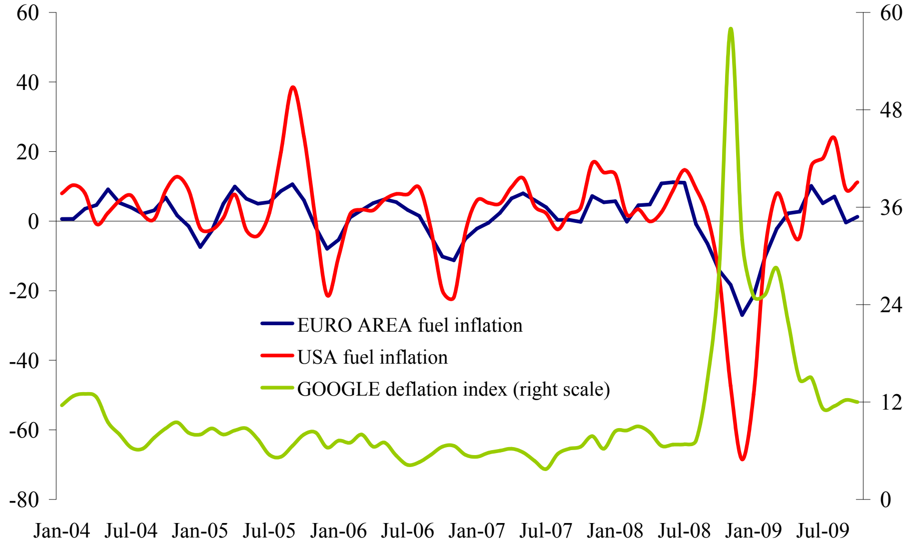

Figure 1: Motor fuels inflation and Google Index of the number of queries of the word "deflation" since 2004

Note to Figure 1. Data for fuel inflation rates are from Eurostat for the euro area and from the Bureau of Labor Statistics for the US. In the graph we report the percentage change over the previous three months of the seasonally adjusted index. The Google Index is the monthly average of the weekly index obtained from *www.google.com/insights/search/#*. The data used in this article were downloaded on December 4^{th}, 2009. The index summarizes the incidence of searches in Google of the word "deflation" on total searches each week. It is normalized to 100 in the week with the highest incidence.

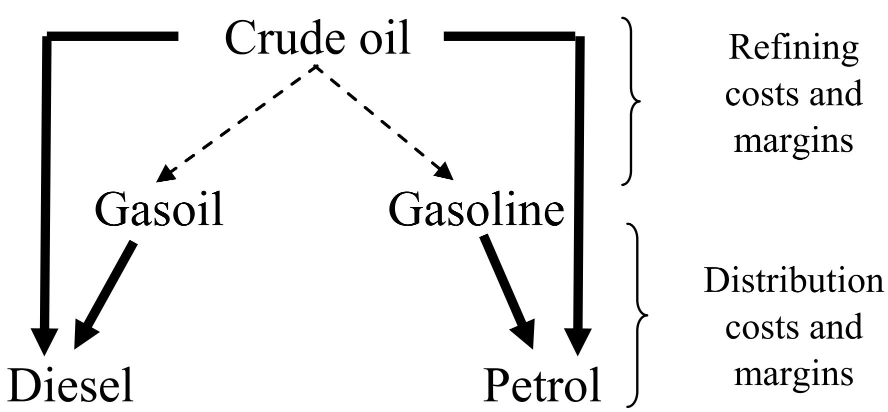

Figure 2: Stylized pricing chain from crude to retail prices

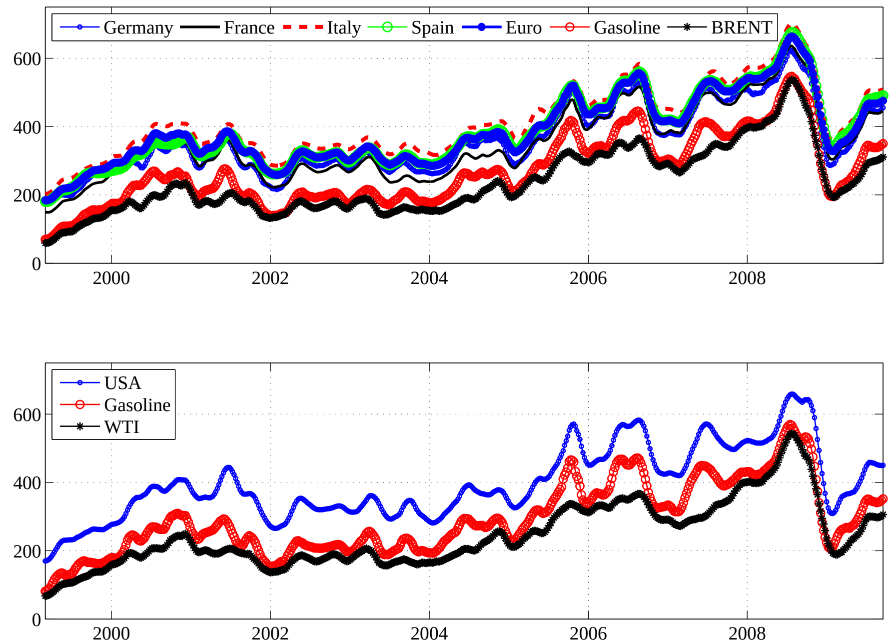

Figure 3: Petrol and oil prices (euros per thousand litres, eight weeks averages)

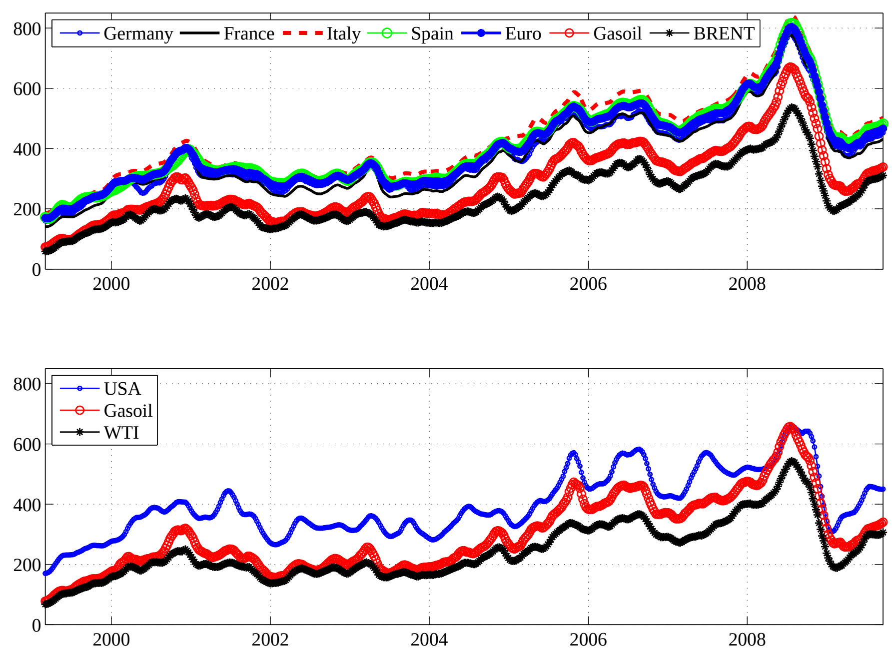

Figure 4: Diesel and oil prices (euros per thousand litres, eight weeks averages)

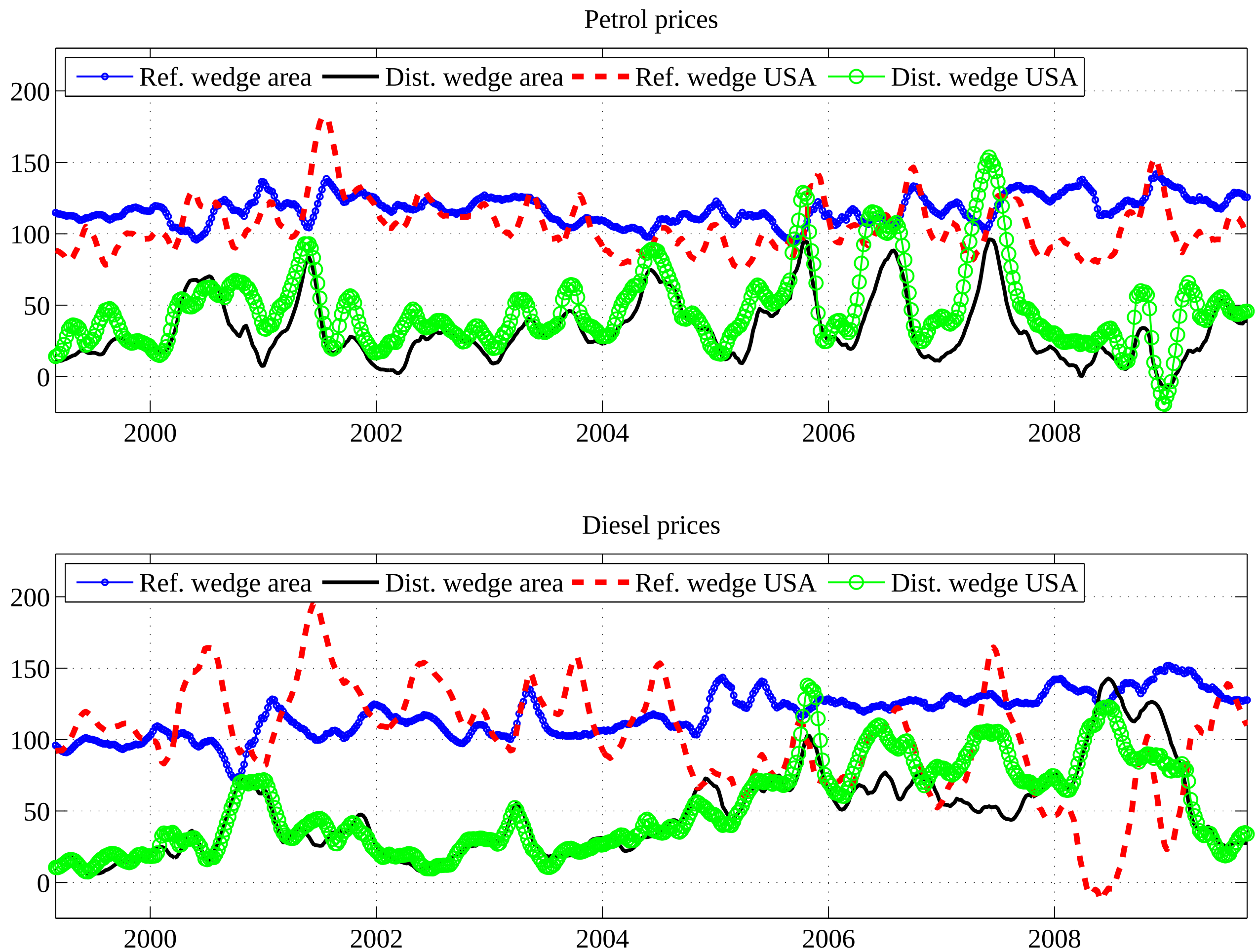

Figure 5: Refining and distribution costs and margins (euros per thousand litres, eight weeks averages)

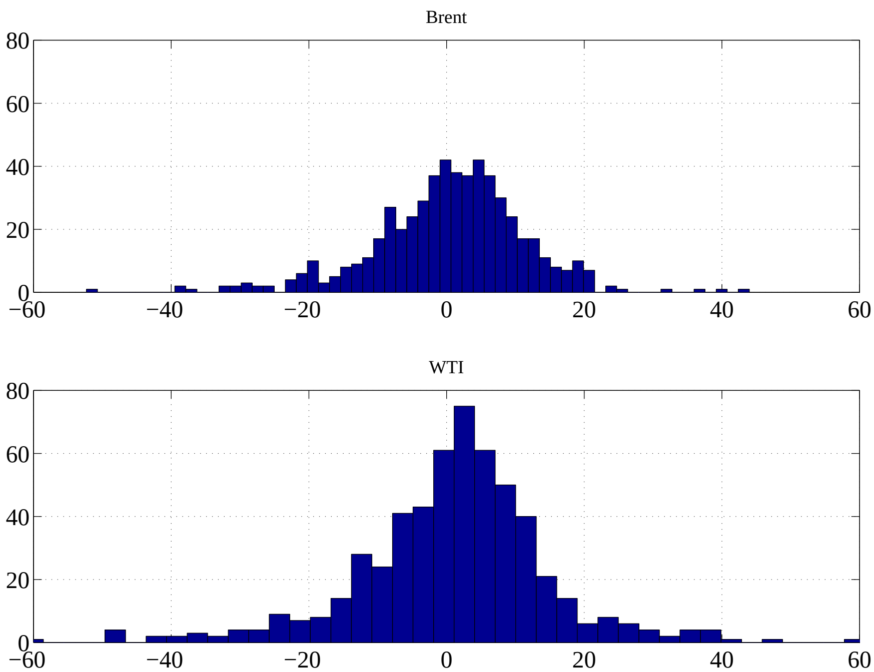

Figure 6: Unconditional distribution of oil (brent) price changes

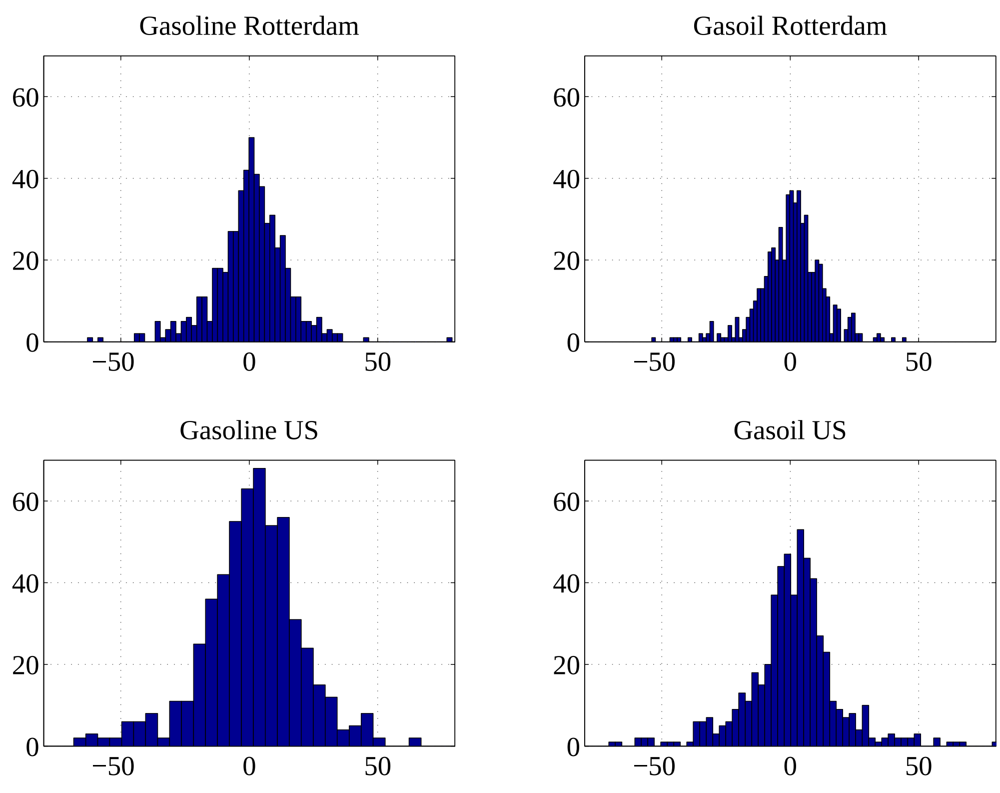

Figure 7: Spot prices changes distribution

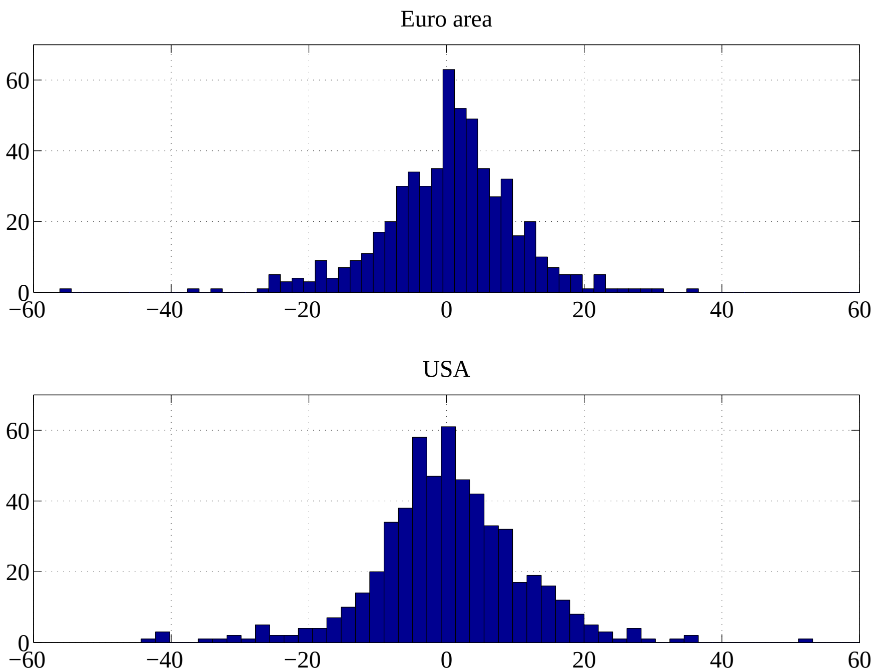

Figure 8: Petrol price changes distribution: euro area and USA

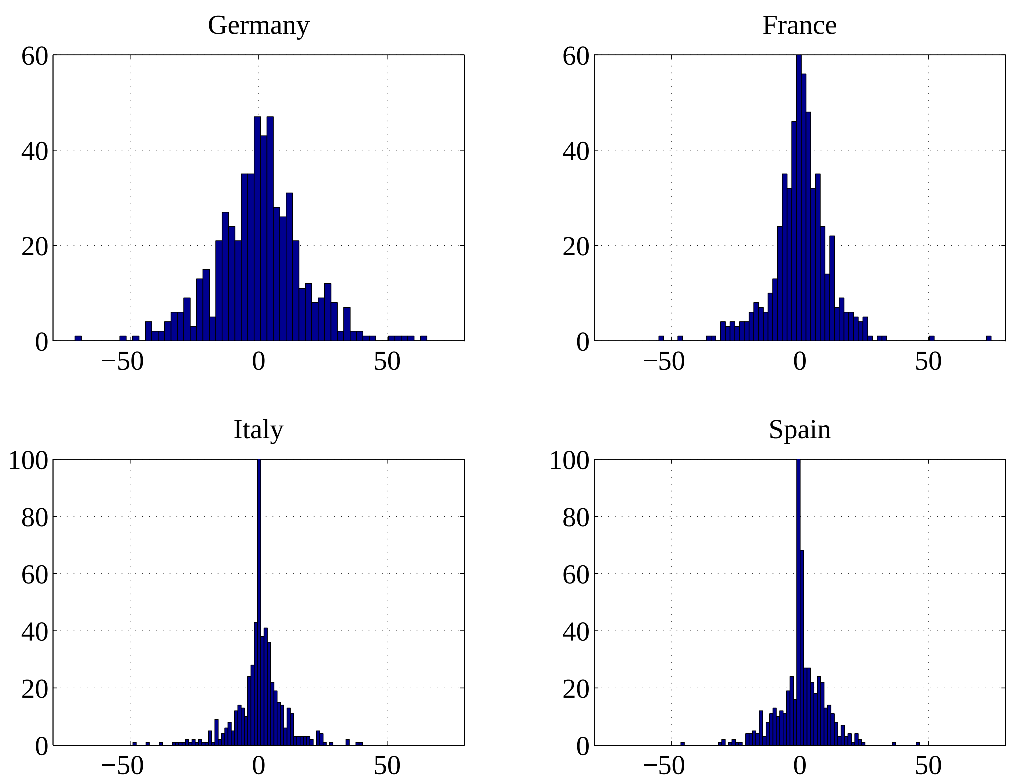

Figure 9: Petrol price changes distribution: euro area countries

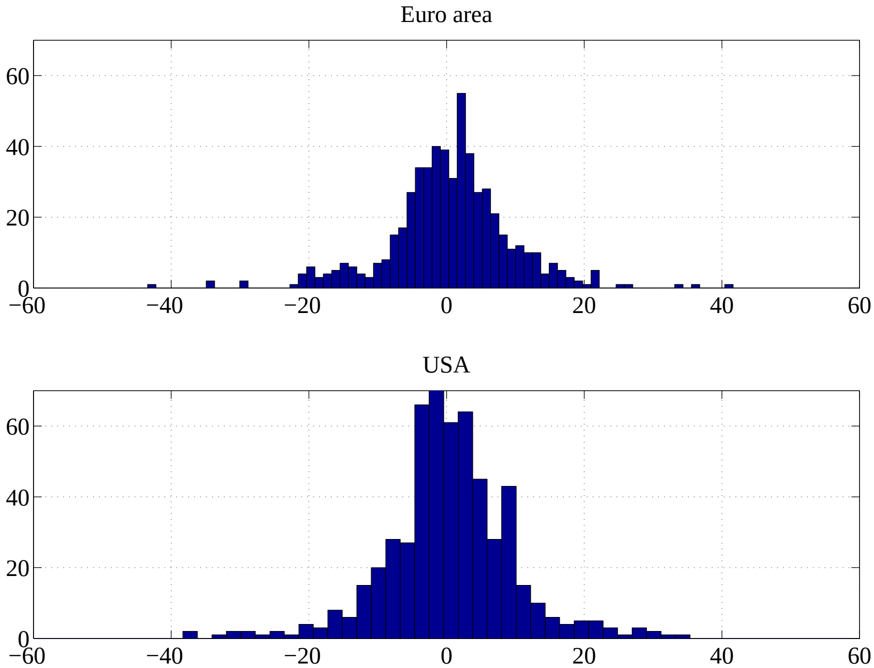

Figure 10: Diesel price changes distribution: euro area and USA

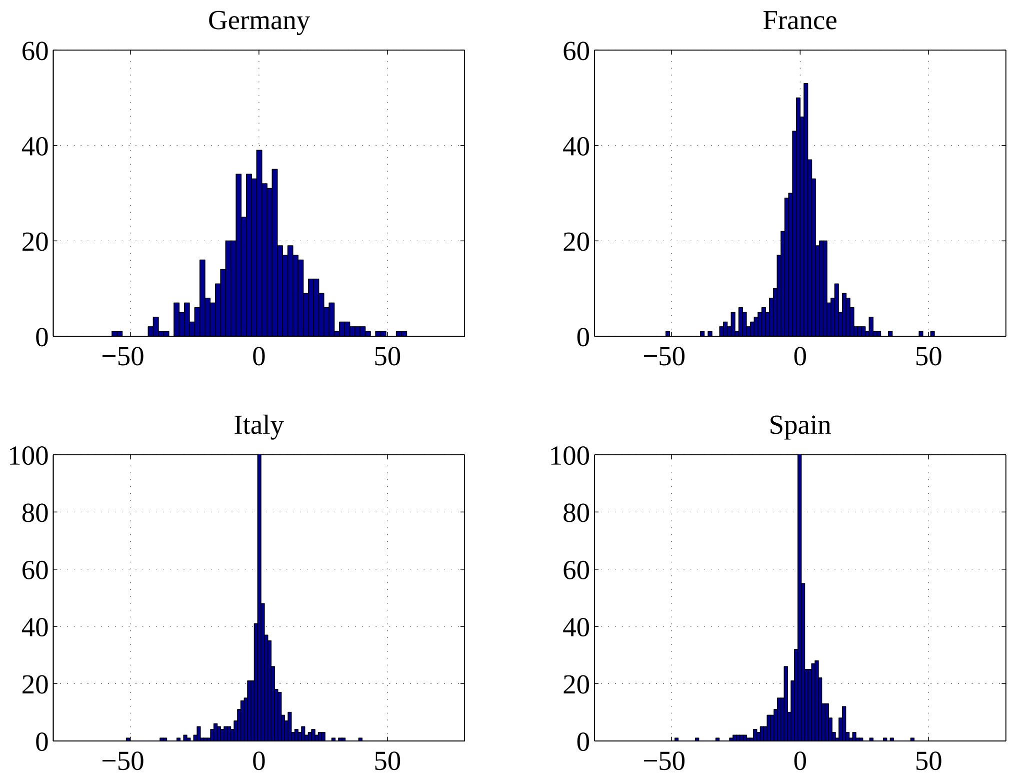

Figure 11: Diesel price changes distribution: euro area countries

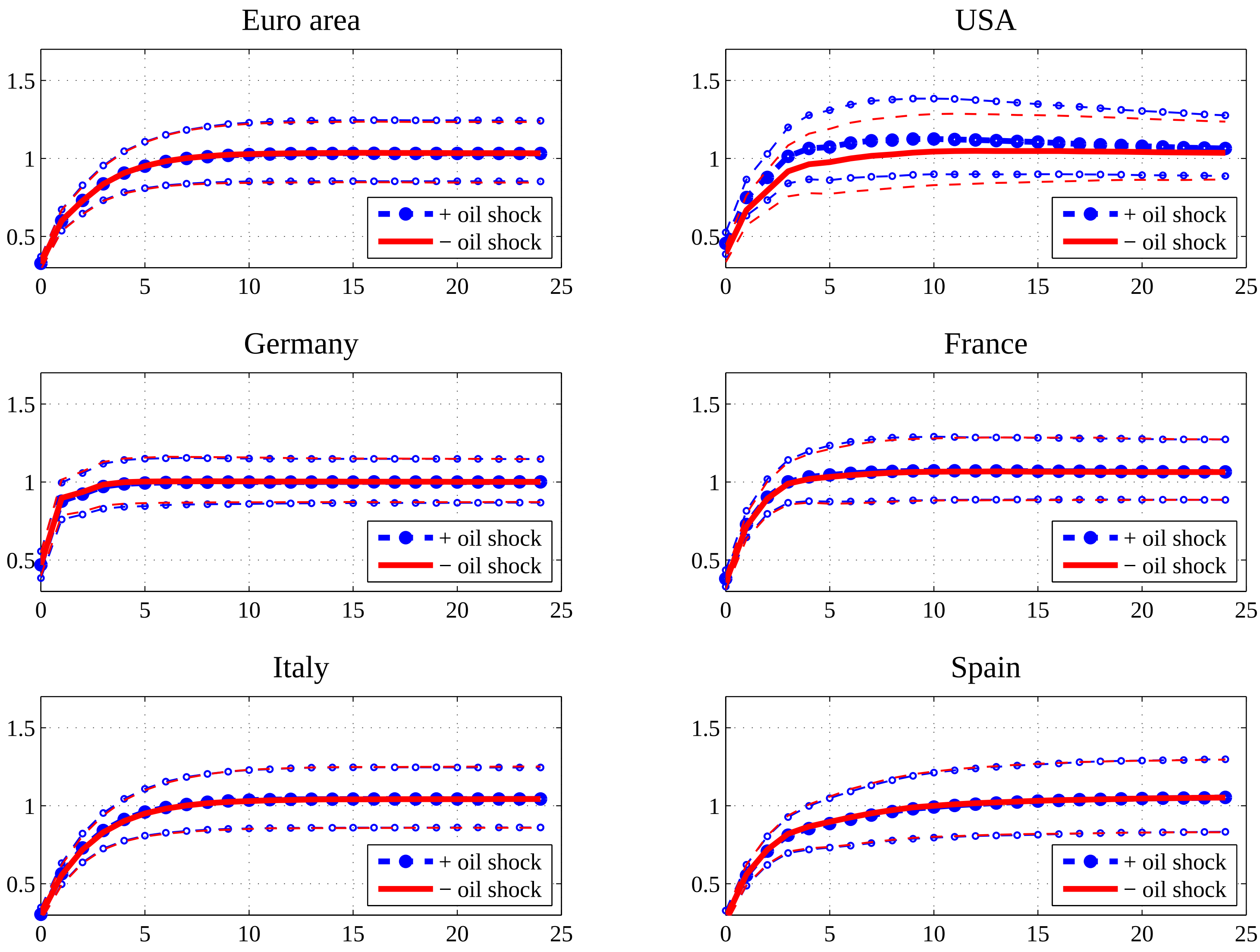

Figure 12: Response of petrol prices to a one standard deviation oil price shock

Note to Figure 12. All impulse responses are normalized by the long-run response of oil prices to the shock.

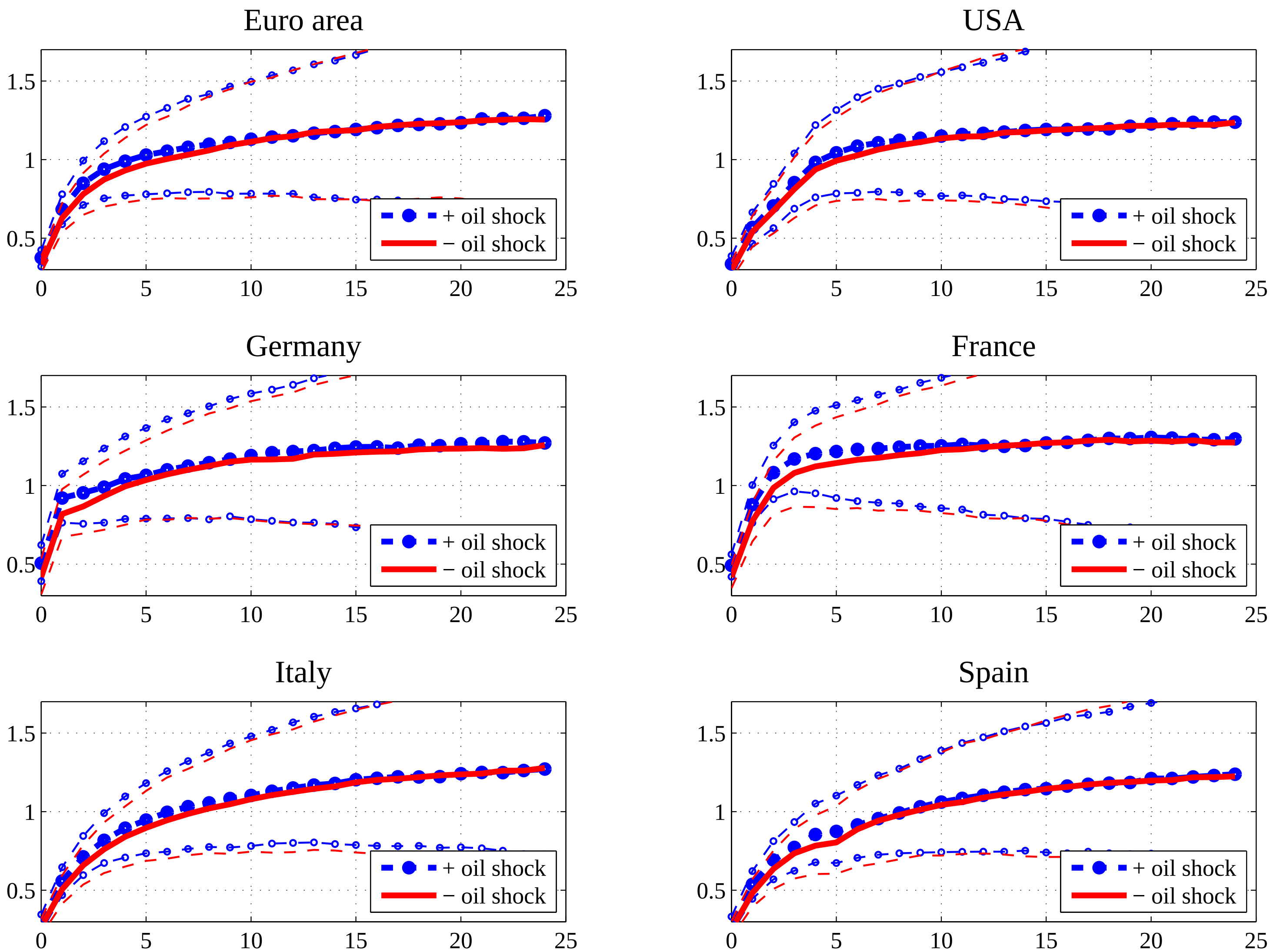

Figure 13: Response of diesel prices to a one standard deviation oil price shock

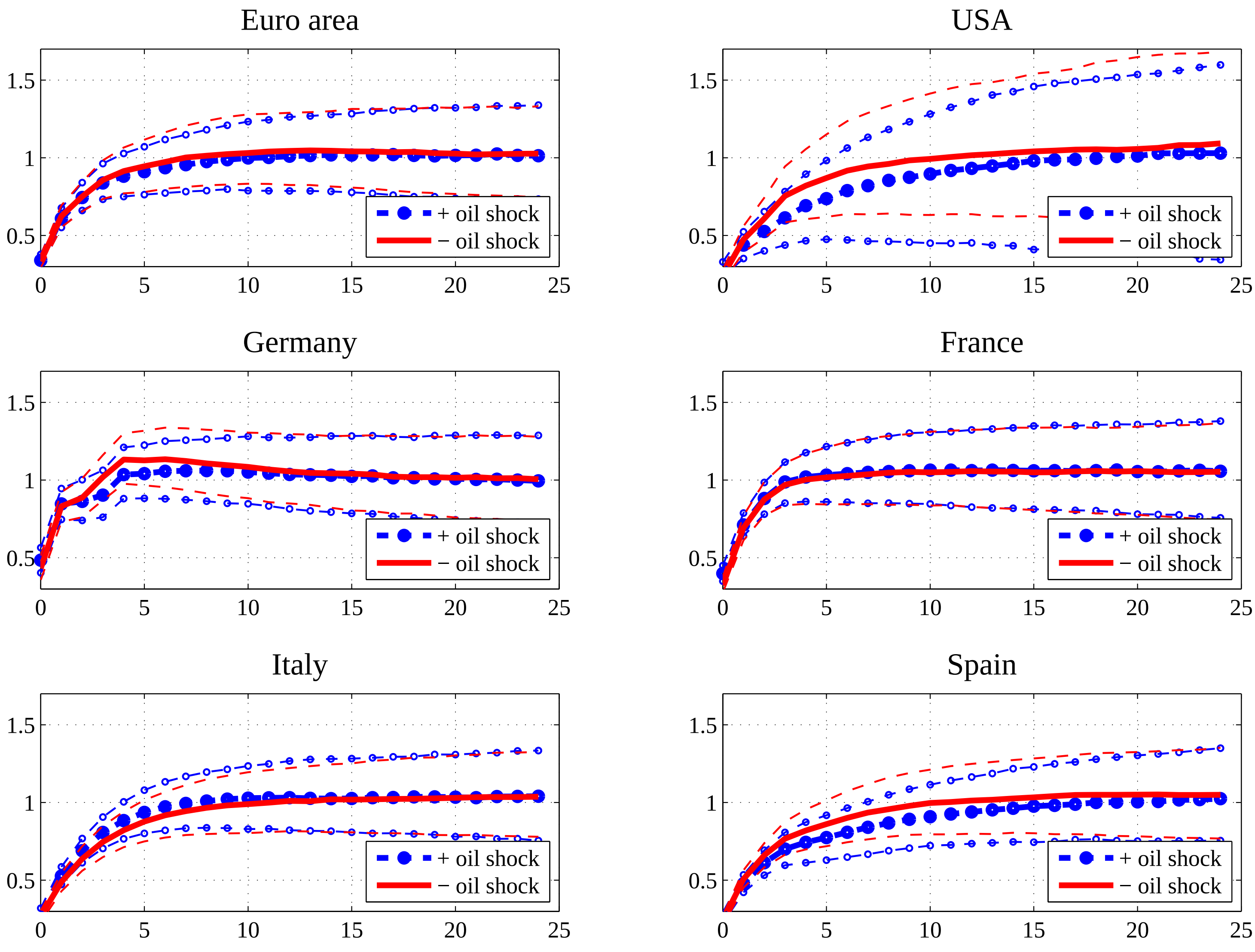

Figure 14: Response of petrol prices to a two standard deviation oil price shock

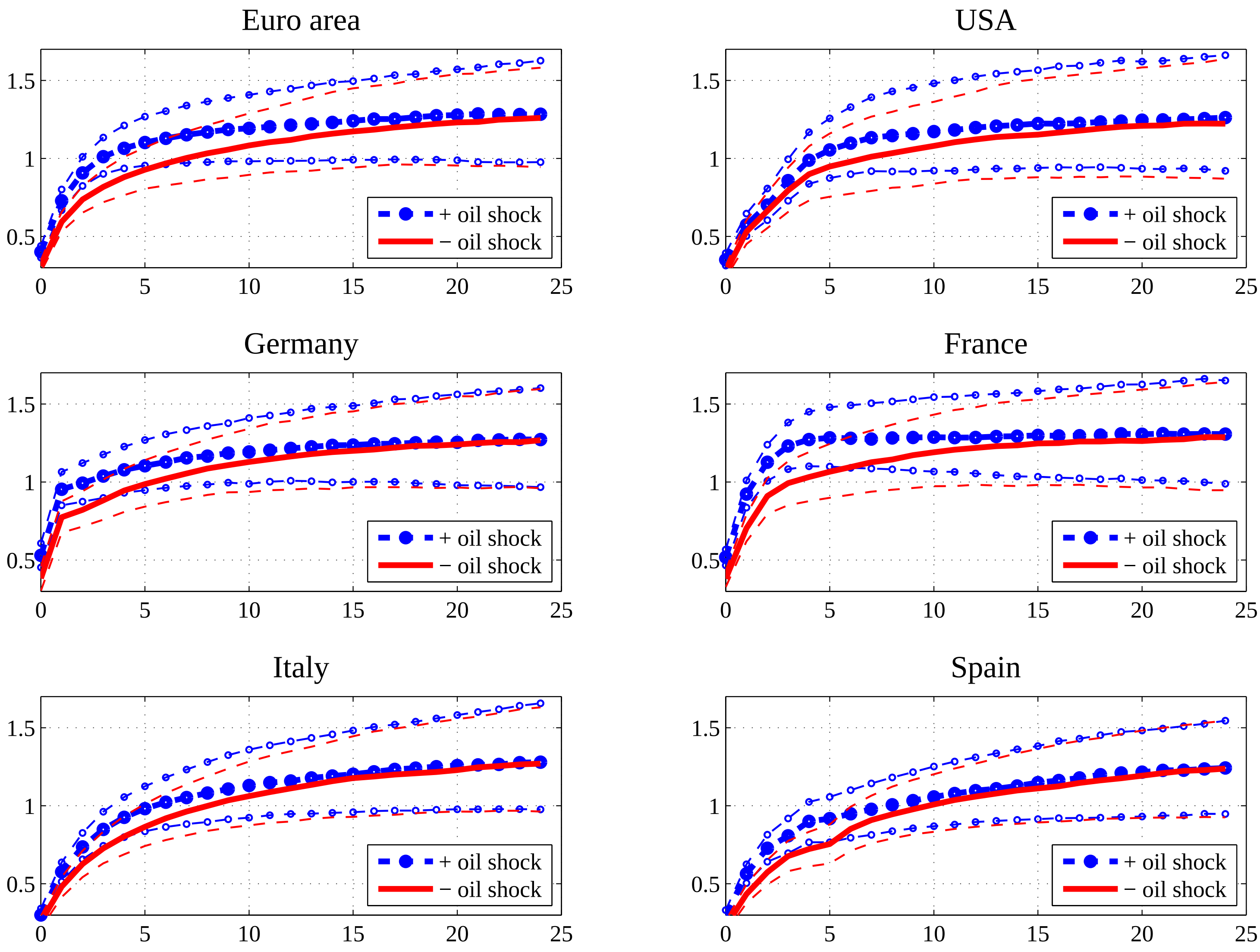

Figure 15: Response of diesel prices to a two standard deviation oil price shock

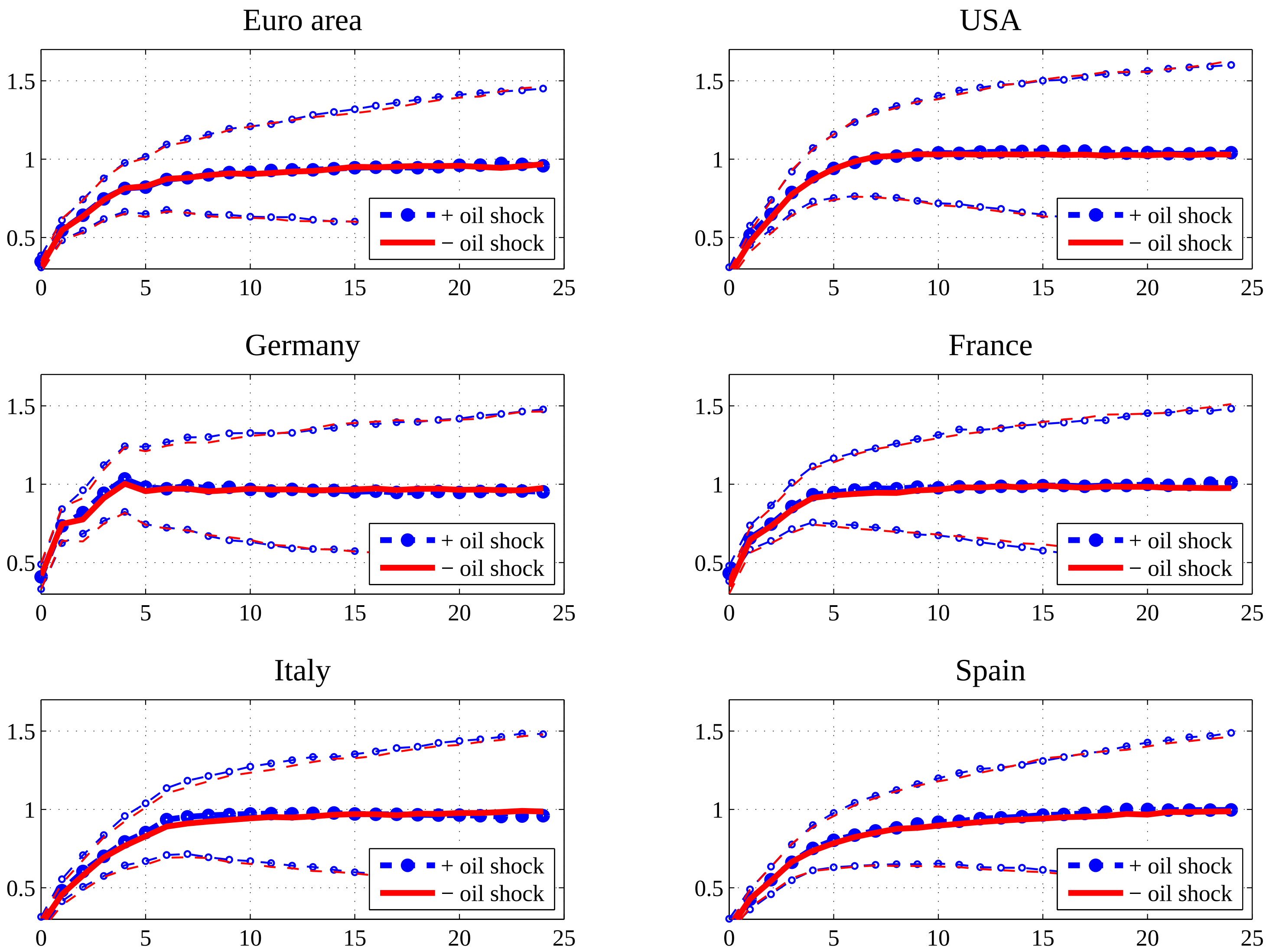

Figure 16: Response of petrol prices to a one standard deviation spot gasoline price shock

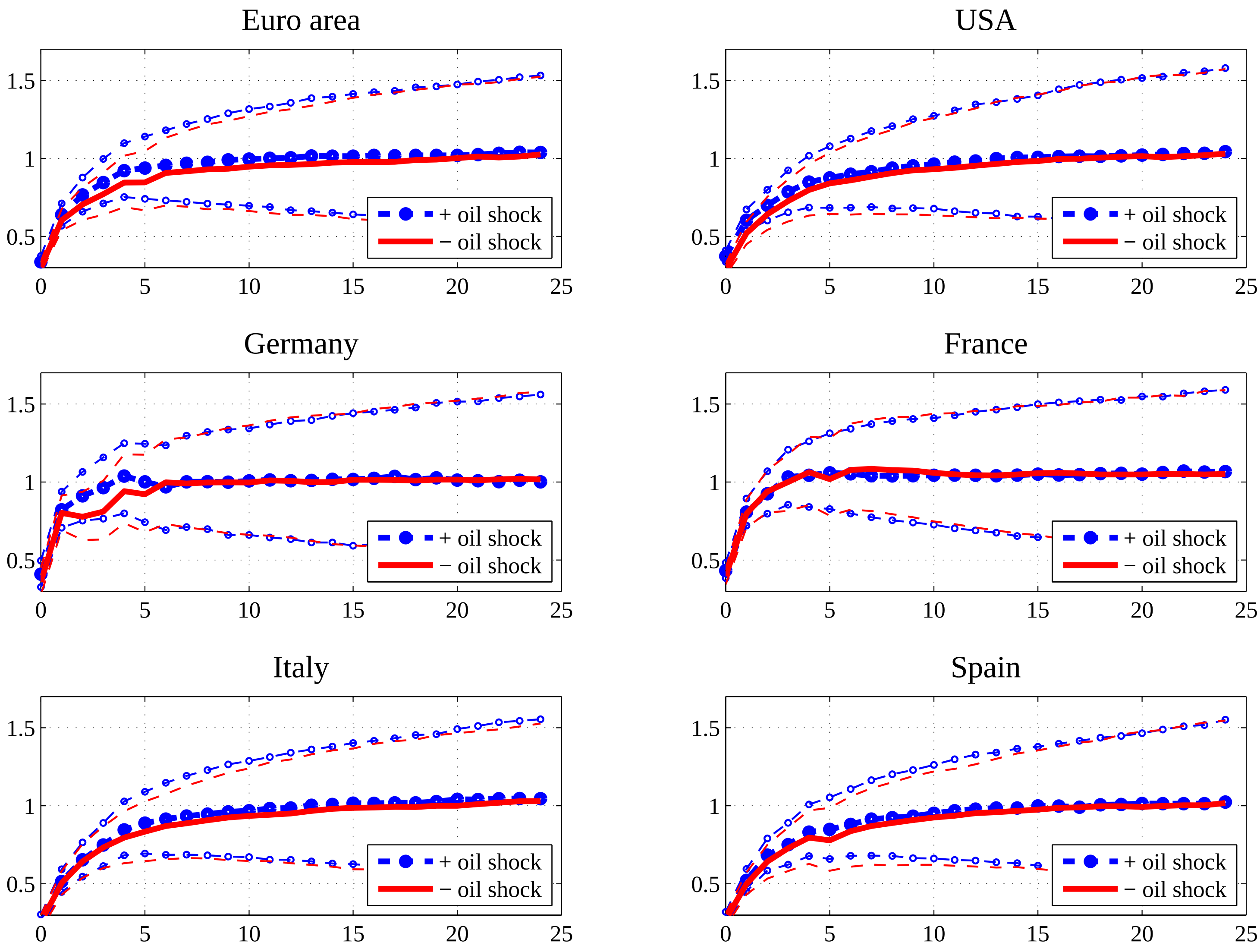

Figure 17: Response of diesel prices to a one standard deviation spot gasoil price shock

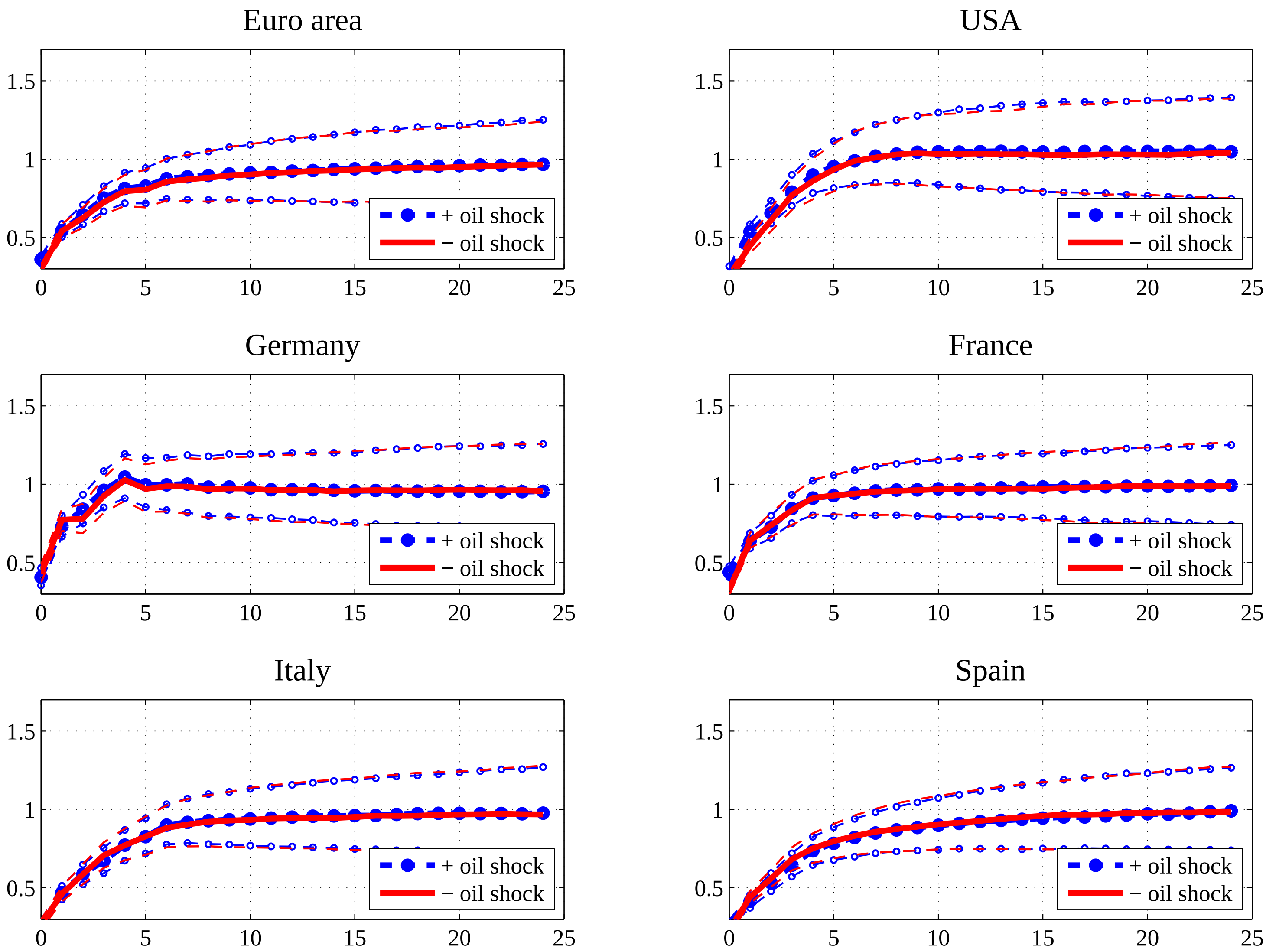

Figure 18: Response of petrol prices to a two standard deviation spot gasoline price shock

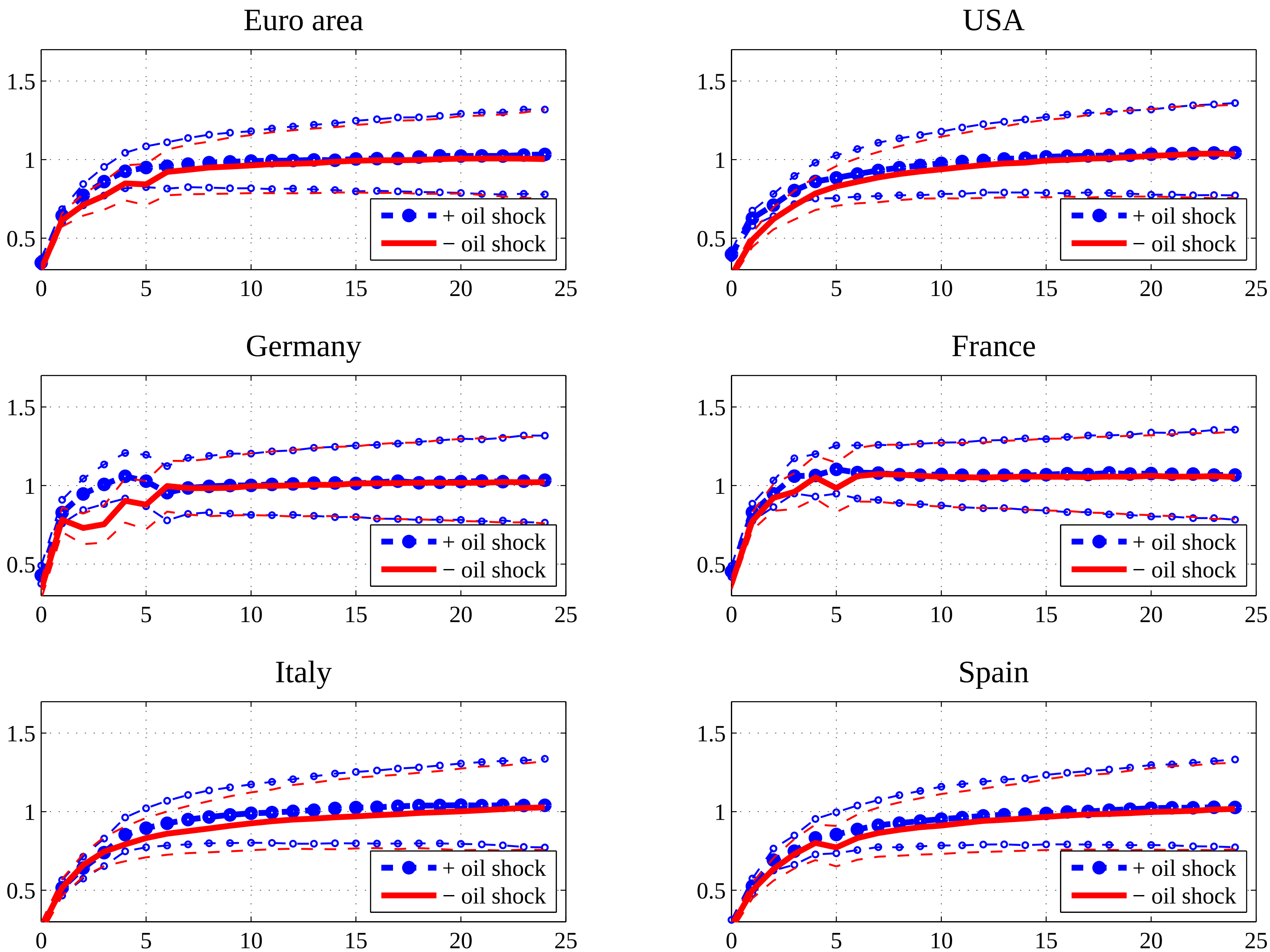

Figure 19: Response of diesel prices to a two standard deviation gasoil price shock

	Levels		First differences	
	ADF	Phillips-Perron	ADF	Phillips-Perron
US - diesel	0.57	0.62	0.00	0.00
US - petrol	0.57	0.63	0.00	0.00
US - gasoil	0.59	0.58	0.00	0.00
US - gasoline	0.47	0.56	0.00	0.00
US - crude	0.67	0.59	0.00	0.00
Germany - petrol	0.69	0.66	0.00	0.00
France - petrol	0.68	0.68	0.00	0.00
Italy - petrol	0.69	0.72	0.00	0.00
Spain - petrol	0.67	0.71	0.00	0.00
euro area - petrol	0.67	0.69	0.00	0.00
Germany - diesel	0.69	0.69	0.00	0.00
France - diesel	0.68	0.68	0.00	0.00
Italy - diesel	0.73	0.73	0.00	0.00
Spain - diesel	0.73	0.73	0.00	0.00
euro area - diesel	0.72	0.71	0.00	0.00
euro area - gasoil	0.62	0.61	0.00	0.00
euro area - gasoline	0.60	0.57	0.00	0.00
euro area - crude oil	0.63	0.64	0.00	0.00

Table 1: Unit root tests on the variables in levels and first differences

Note to Table 1. The lag length in the regressions on which the unit root tests are based is selected on the basis of the Akaike criterion.

	Germany	France	Italy	Spain	Euro area	USA
			Long-run equation			
α	121.50	98.25	166.47	133.47	140.57	144.65
ϕ	1.00	1.07	1.05	1.07	1.04	1.15
Cointegration test (p-val)	0.00	0.00	0.00	0.00	0.00	0.00
			Petrol equation			
lags petrol - petrol equation	1	2	2	2	3	4
lags oil - petrol equation	1	1	2	2	2	4
ECM drop test (p-val)	0.02	0.00	0.01	0.00	0.00	0.00
Res. heteros. (p-val)	0.02	0.00	0.01	0.00	0.00	0.00
Res. autocor (p-val)	0.95	0.55	0.82	0.70	0.89	0.59
$\bar{R}$ fuel equation	0.34	0.52	0.51	0.55	0.55	0.65
			Oil equation			
lags petrol - oil equation	2	3	2	3	2	4
lags oil - oil equation	1	4	1	4	1	3
Drop autocorr. terms	0.00	0.00	0.00	0.00	0.00	0.05
Drop feedback test (p-val)	0.01	0.00	0.00	0.00	0.00	0.00
Res. heteros. (p-val)	0.00	0.00	0.00	0.00	0.00	0.00
Res. autocor (p-val)	0.75	0.87	0.62	0.47	0.47	0.24
$\bar{R}$ oil equation	0.06	0.08	0.08	0.09	0.07	0.11
Number of obs.	557	555	557	555	557	555

Table 2: Regression results summary: from crude to petrol prices

Note to Table 2. The estimation sample goes from the first week of January 1999 to the second week of September 2009. The cointegration test is a unit root test (Augmented Dickey Fuller) on the residuals of the long run equation. The lags in the dynamic equations are selected on the basis of the Akaike criterion. The ECM drop test is a Wald test on the ECM coefficient d_{fuel} in equation (2). The residuals heteroskedasticity test is a Lagrange Multiplier test for ARCH effects up to the fourth order. The residuals autocorrelation test in the dynamic fuels equation is a Breusch-Godfrey Lagrange Multiplier test for serial correlation up to the fourth order. The drop feedback test is a joint Wald test on the $b_{12,i}$ coefficients in equation (1).

	Germany	France	Italy	Spain	Euro area	USA
			Long-run equation			
α	79.78	61.05	102.61	89.78	86.90	44.52
ϕ	1.29	1.32	1.39	1.34	1.33	1.33
Cointegration test (p-val)	0.00	0.00	0.00	0.00	0.00	0.00
			Diesel equation			
lags diesel - diesel equation	2	3	2	1	2	3
lags oil - diesel equation	1	1	1	6	1	1
ECM drop test (p-val)	0.00	0.00	0.00	0.00	0.00	0.00
Res. heteros. (p-val)	0.02	0.00	0.01	0.00	0.00	0.00
Res. autocor (p-val)	0.59	0.63	0.54	0.48	0.53	0.98
$\bar{R}$ fuel equation	0.36	0.57	0.50	0.55	0.63	0.62
			Oil equation			
lags diesel - oil equation	1	2	5	1	1	4
lags oil - oil equation	1	1	1	1	1	1
Drop autocorr. terms	0.00	0.00	0.00	0.00	0.00	0.03
Drop feedback test (p-val)	0.93	0.22	0.01	0.03	0.72	0.00
Res. heteros. (p-val)	0.00	0.00	0.00	0.00	0.00	0.04
Res. autocor (p-val)	0.63	0.17	0.85	0.57	0.50	0.46
$\bar{R}$ oil equation	0.05	0.05	0.07	0.06	0.05	0.07
Number of obs.	558	557	554	558	558	555

Table 3: Regression results summary: from crude to diesel prices

Note to Table 3. The estimation sample goes from the first week of January 1999 to the second week of September 2009. The cointegration test is a unit root test (Augmented Dickey Fuller) on the residuals of the long run equation. The lags in the dynamic equations are selected on the basis of the Akaike criterion. The ECM drop test is a Wald test on the ECM coefficient d_{fuel} in equation (2). The residuals heteroskedasticity test is a Lagrange Multiplier test for ARCH effects up to the fourth order. The residuals autocorrelation test in the dynamic fuels equation is a Breusch-Godfrey Lagrange Multiplier test for serial correlation up to the fourth order. The drop feedback test is a joint Wald test on the $b_{12,i}$ coefficients in equation (1).

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.73	0.21	0.77	0.12	0.63	0.45
1	0.85	0.60	0.94	0.47	0.94	0.86
3	0.96	0.93	0.96	0.85	0.99	0.99

Table 4: Test of symmetry of the response of petrol prices to a 1 s.d. oil price shock

Note to Tables from 4 to 7. P-values reported in these tables are based on the χ^2_{H+1} distribution.

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.49	0.18	0.73	0.84	0.47	0.74
1	0.89	0.57	0.88	0.87	0.80	0.95
3	0.99	0.89	0.99	0.93	0.68	0.99

Table 5: Test of symmetry of the response of diesel prices to a 1 s.d. oil price shock

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.79	0.58	0.43	0.69	0.20	0.26
1	0.98	0.80	0.74	0.66	0.63	0.29
3	1.00	0.95	0.93	0.97	0.88	0.88

Table 6: Test of symmetry of the response of petrol prices to a 2 s.d. oil price shock

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.16	0.17	0.19	0.46	**0.03**	0.22
1	0.60	0.62	0.66	0.84	0.23	0.51
3	0.93	0.92	0.86	0.98	0.82	0.93

Table 7: Test of symmetry of the response of diesel prices to a 2 s.d. oil price shock

	Germany	France	Italy	Spain	Euro area	USA
			Long-run equation			
α	96.77	75.21	144.91	111.02	117.33	101.62
ϕ	0.97	1.03	1.00	1.03	1.00	1.09
Cointegration test (p-val)	0.00	0.00	0.00	0.00	0.00	0.00
			Petrol equation			
lags fuel - fuel equation	4	3	3	1	4	4
lags oil - fuel equation	4	4	6	4	6	1
ECM drop test (p-val)	0.00	0.01	0.01	0.00	0.02	0.00
Res. heteros. (p-val)	0.00	0.01	0.01	0.00	0.02	0.01
Res. autocor (p-val)	0.59	0.25	0.17	0.61	0.02	0.32
$\bar{R}$ fuel equation	0.51	0.78	0.63	0.69	0.77	0.86
			Gasoline equation			
lags gas - oil equation	4	1	3	1	2	4
lags oil - oil equation	3	3	3	3	4	3
Drop autocorr. terms	0.00	0.00	0.00	0.04	0.02	0.01
Drop feedback terms	0.21	0.02	0.03	0.09	0.09	0.00
Res. heteros. (p-val)	0.02	0.26	0.07	0.24	0.05	0.00
Res. autocor (p-val)	0.99	0.71	0.99	0.97	0.76	0.56
$\bar{R}$ oil equation	0.05	0.06	0.07	0.07	0.06	0.09
Number of obs.	555	556	556	556	555	555

Table 8: Regression results summary: from refined gasoline to petrol prices

Note to Table 8. The estimation sample goes from the first week of January 1999 to the second week of September 2009. The cointegration test is a unit root test (Augmented Dickey Fuller) on the residuals of the long run equation. The lags in the dynamic equations are selected on the basis of the Akaike criterion. The ECM drop test is a Wald test on the ECM coefficient d_{fuel} in equation (2). The residuals heteroskedasticity test is a Lagrange Multiplier test for ARCH effects up to the fourth order. The residuals autocorrelation test in the dynamic fuels equation is a Breusch-Godfrey Lagrange Multiplier test for serial correlation up to the fourth order. The drop feedback test is a joint Wald test on the $b_{12,i}$ coefficients in equation (1).

	Germany	France	Italy	Spain	Euro area	USA
			Long-run equation			
α	89.58	70.08	112.95	100.07	96.49	40.85
ϕ	1.04	1.07	1.12	1.08	1.07	1.08
Cointegration test (p-val)	0.00	0.00	0.00	0.00	0.00	0.00
			Diesel equation			
lags fuel - fuel equation	4	1	2	1	1	3
lags oil - fuel equation	6	6	4	6	6	1
ECM drop test (p-val)	0.00	0.00	0.00	0.01	0.01	0.00
Res. heteros. (p-val)	0.00	0.00	0.00	0.01	0.01	0.00
Res. autocor (p-val)	0.14	0.54	0.61	0.19	0.95	0.01
$\bar{R}$ fuel equation	0.49	0.78	0.60	0.66	0.79	0.81
			Gasoil equation			
lags gas - oil equation	1	2	2	2	1	4
lags oil - oil equation	3	3	1	1	3	1
Drop autocorr. terms	0.00	0.00	0.00	0.00	0.00	0.00
Drop feedback terms	0.23	0.00	0.13	0.48	0.50	0.00
Res. heteros. (p-val)	0.09	0.05	0.06	0.03	0.09	0.00
Res. autocor (p-val)	0.49	0.93	0.30	0.17	0.37	0.58
$\bar{R}$ oil equation	0.07	0.09	0.09	0.07	0.07	0.06
Number of obs.	556	556	557	557	556	555

Table 9: Regression results summary: from refined gasoil to diesel prices

Note to Table 9. The estimation sample goes from the first week of January 1999 to the second week of September 2009. The cointegration test is a unit root test (Augmented Dickey Fuller) on the residuals of the long run equation. The lags in the dynamic equations are selected on the basis of the Akaike criterion. The ECM drop test is a Wald test on the ECM coefficient d_{fuel} in equation (2). The residuals heteroskedasticity test is a Lagrange Multiplier test for ARCH effects up to the fourth order. The residuals autocorrelation test in the dynamic fuels equation is a Breusch-Godfrey Lagrange Multiplier test for serial correlation up to the fourth order. The drop feedback test is a joint Wald test on the $b_{12,i}$ coefficients in equation (1).

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.75	0.98	0.89	0.06	0.59	**0.01**
1	0.96	0.80	0.98	0.41	0.90	0.19
3	1.00	0.94	0.99	0.90	0.97	0.66

Table 10: Test of symmetry of the response of petrol prices to a 1 s.d. spot gasoline price shock

Note to Tables from 10 to 13. P-values reported in these tables are based on the χ^2_{H+1} distribution.

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.31	0.32	0.68	0.33	0.56	0.39
1	0.76	0.24	0.92	0.79	0.39	0.53
3	0.99	0.73	0.99	0.94	0.83	0.90

Table 11: Test of symmetry of the response of diesel prices to a 1 s.d. spot gasoil price shock

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.62	**0.01**	0.60	0.26	0.40	0.63
1	0.81	**0.02**	0.87	0.14	0.44	0.94
3	0.92	0.37	0.99	0.75	0.91	0.90

Table 12: Test of symmetry of the response of petrol prices to a 2 s.d. spot gasoline price shock

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.92	0.68	0.69	0.42	0.38	**0.00**
1	1.00	0.91	0.59	0.84	0.62	**0.05**
3	0.74	0.99	0.96	0.99	0.96	0.40

Table 13: Test of symmetry of the response of diesel prices to a 2 s.d. spot gasoil price shock

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.45	0.75	0.82	0.31	0.16	0.57
1	0.65	0.87	0.88	0.76	0.52	0.91
3	0.96	0.92	0.98	0.98	0.92	1.00

Table 14: Test of symmetry of the response of petrol prices to a 1 s.d. oil price shock - shocks simulated from an ARCH(4) model

Note to Tables from 14 to 17. P-values reported in these tables are based on the χ^2_{H+1} distribution.

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.98	0.81	0.92	0.14	0.18	0.96
1	1.00	0.56	0.99	0.44	0.61	0.97
3	1.00	0.87	0.99	0.91	0.97	1.00

Table 15: Test of symmetry of the response of diesel prices to a 1 s.d. oil price shock - shocks simulated from an ARCH(4) model

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.39	0.29	0.77	0.99	0.89	0.88
1	0.83	0.62	0.98	0.98	0.73	0.92
3	0.96	0.97	1.00	1.00	0.94	0.99

Table 16: Test of symmetry of the response of petrol prices to a 2 s.d. oil price shock - shocks simulated from an ARCH(4) model

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.41	**0.05**	0.78	0.20	0.56	0.62
1	0.52	0.33	0.98	0.47	0.91	0.88
3	0.92	0.69	1.00	0.91	0.96	0.98

Table 17: Test of symmetry of the response of diesel prices to a 2 s.d. oil price shock - shocks simulated from an ARCH(4) model

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.14	0.18	0.39	0.93	0.81	0.47
1	0.52	0.47	0.73	0.88	0.91	0.84
3	0.95	0.86	0.97	0.83	0.97	0.98

Table 18: Test of symmetry of the response of petrol prices to a 1 s.d. spot gasoline price shock - shocks simulated from an ARCH(4) model

Note to Tables from 18 to 21. P-values reported in these tables are based on the χ^2_{H+1} distribution.

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.83	0.90	0.80	0.45	0.84	0.70
1	0.92	0.96	0.98	0.87	0.97	0.83
3	0.96	1.00	1.00	0.91	0.95	0.99

Table 19: Test of symmetry of the response of diesel prices to a 1 s.d. spot gasoil price shock - shocks simulated from an ARCH(4) model

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.82	**0.05**	0.79	0.08	0.67	0.76
1	0.89	0.23	0.93	0.36	0.42	0.92
3	0.92	0.79	0.98	0.87	0.89	1.00

Table 20: Test of symmetry of the response of petrol prices to a 2 s.d. spot gasoline price shock - shocks simulated from an ARCH(4) model

Steps (H)	Ger	Fra	Ita	Spa	EA	USA
0	0.22	**0.01**	0.86	0.99	0.79	0.36
1	0.68	0.15	0.85	1.00	0.98	0.73
3	0.93	0.72	0.99	1.00	1.00	0.96

Table 21: Test of symmetry of the response of diesel prices to a 2 s.d. spot gasoil price shock - shocks simulated from an ARCH(4) model

	Jan. 1999 - Apr. 2004						Apr. 2004 - Sep. 2009					
Steps (H)	Ger	Fra	Ita	Spa	EA	USA	Ger	Fra	Ita	Spa	EA	USA
0	0.40	0.85	**0.04**	0.63	0.19	0.68	0.58	0.74	0.73	0.96	0.98	0.84
1	0.65	0.99	0.29	0.94	0.61	0.95	0.79	0.65	0.85	0.48	1.00	0.99
3	0.94	1.00	0.68	1.00	0.92	1.00	0.99	0.97	0.96	0.88	1.00	0.98

Table 22: Test of symmetry of the response of petrol prices to a 1 s.d. oil price shock - sample split analysis

	Jan. 1999 - Apr. 2004						Apr. 2004 - Sep. 2009					
Steps (H)	Ger	Fra	Ita	Spa	EA	USA	Ger	Fra	Ita	Spa	EA	USA
0	0.06	**0.03**	0.27	0.40	0.27	0.10	**0.01**	0.28	0.85	0.99	0.85	0.56
1	0.07	0.26	0.71	0.64	0.56	0.50	0.27	0.66	0.95	0.96	0.97	0.60
3	0.57	0.73	0.99	0.96	0.85	0.93	0.84	0.98	0.94	0.99	0.95	0.97

Table 23: Test of symmetry of the response of diesel prices to a 1 s.d. oil price shock - sample split analysis

	Jan. 1999 - Apr. 2004						Apr. 2004 - Sep. 2009					
Steps (H)	Ger	Fra	Ita	Spa	EA	USA	Ger	Fra	Ita	Spa	EA	USA
0	0.21	0.27	0.88	0.47	0.75	0.72	0.68	0.25	0.37	0.65	0.88	0.19
1	0.69	0.38	0.54	0.47	0.97	0.60	0.96	0.71	0.63	0.85	0.75	0.19
3	0.98	0.88	0.94	0.94	1.00	0.93	1.00	0.96	0.95	0.99	0.99	0.72

Table 24: Test of symmetry of the response of petrol prices to a 2 s.d. oil price shock - sample split analysis

	Jan. 1999 - Apr. 2004						Apr. 2004 - Sep. 2009					
Steps (H)	Ger	Fra	Ita	Spa	EA	USA	Ger	Fra	Ita	Spa	EA	USA
0	**0.04**	0.08	0.20	0.77	0.35	**0.05**	0.40	**0.01**	0.48	0.27	**0.00**	**0.01**
1	0.33	0.45	0.66	0.47	0.80	0.38	0.73	0.21	0.52	0.54	0.07	0.08
3	0.86	0.50	0.93	0.93	0.99	0..87	0.98	0.80	0.91	0.93	0.58	0.63

Table 25: Test of symmetry of the response of diesel prices to a 2 s.d. oil price shock - sample split analysis

	Jan. 1999 - Apr. 2004						Apr. 2004 - Sep. 2009					
Steps (H)	Ger	Fra	Ita	Spa	EA	USA	Ger	Fra	Ita	Spa	EA	USA
0	0.50	0.22	0.95	0.05	0.10	0.35	0.21	0.34	0.77	0.77	0.42	0.64
1	0.87	0.52	0.85	0.40	0.49	0.81	0.65	0.73	0.61	0.81	0.59	0.93
3	1.00	0.94	0.99	0.89	0.90	0.95	0.92	0.92	0.96	0.99	0.83	1.00

Table 26: Test of symmetry of the response of petrol prices to a 1 s.d. spot gasoline price shock - sample split analysis

	Jan. 1999 - Apr. 2004						Apr. 2004 - Sep. 2009					
Steps (H)	Ger	Fra	Ita	Spa	EA	USA	Ger	Fra	Ita	Spa	EA	USA
0	0.31	0.90	0.95	0.91	**0.01**	0.06	0.10	0.98	0.98	0.52	0.60	**0.01**
1	0.76	0.97	0.87	0.65	0.09	0.17	0.33	0.81	0.69	0.9	0.77	0.20
3	0.89	0.99	0.97	0.94	0.54	0.64	0.76	0.88	0.95	0.99	0.98	0.71

Table 27: Test of symmetry of the response of diesel prices to a 1 s.d. spot gasoil price shock - sample split analysis

	Jan. 1999 - Apr. 2004						Apr. 2004 - Sep. 2009					
Steps (H)	Ger	Fra	Ita	Spa	EA	USA	Ger	Fra	Ita	Spa	EA	USA
0	0.60	**0.02**	0.92	0.43	0.17	0.21	0.85	**0.01**	0.67	0.55	**0.03**	0.25
1	0.79	0.20	0.94	0.11	0.42	0.52	0.96	0.18	0.96	0.09	0.29	0.61
3	0.99	0.70	0.99	0.68	0.89	0.85	0.93	0.64	1.00	0.59	0.54	0.93

Table 28: Test of symmetry of the response of petrol prices to a 2 s.d. spot gasoline price shock - sample split analysis

	Jan. 1999 - Apr. 2004						Apr. 2004 - Sep. 2009					
Steps (H)	Ger	Fra	Ita	Spa	EA	USA	Ger	Fra	Ita	Spa	EA	USA
0	0.79	**0.00**	0.72	0.64	**0.01**	**0.00**	0.39	0.22	0.28	0.65	**0.01**	**0.00**
1	0.96	0.10	0.76	0.67	0.14	**0.02**	0.81	0.64	0.74	0.83	0.19	**0.02**
3	0.93	0.63	0.88	0.95	0..70	0.33	0.82	0.68	0.99	0.94	0.73	0.41

Table 29: Test of symmetry of the response of diesel prices to a 2 s.d. spot gasoil price shock - sample split analysis

A Impulse response function approximation

In this Appendix I compare the true IRF $I_y^h(\beta, \delta)$ with (i) the IRF conditional on the mean history of the data $I_y^h(\beta, \bar{\omega}_{t-1}, \delta)$ (ii) the CAF. Given that the true IRF requires a large amount of simulations, as one needs to average across all possible histories and future shocks, I conduct this comparison only for one set of parameters, that is the least square estimate $\beta = \beta_{ls}$ and only for a shock of one standard deviation.[18] The results (Figures 20 and 21) show that the approximate IRF (IRF (mean) in the graphs) tracks very well the correct one.[19] The CAF, on the other hand, strongly understates the long run pass-through of oil prices to liquid fuels prices. This is because the CAF ignores the endogenous response of oil prices following an own shock, and therefore approximates the long run oil price response with the initial one (which in turn equals the size of the shock).

[18]I do not repeat this exercise for large shocks, since Kilian and Vigufsson (2009) show that the importance of conditioning on the history of the data vanishes when the shock becomes large. Results for a negative shock as well as for refined products are in line with those for a positive shock; I do not present them for brevity.

[19]All the IRF are rescaled by the long run response of the oil price obtained in the simulation that takes into account all sources of uncertainty.

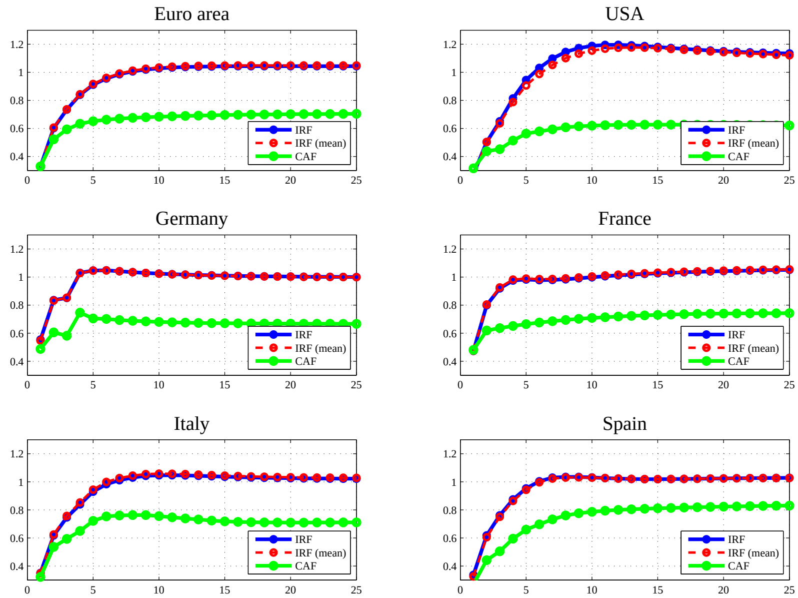

Figure 20: Correct and approximate IRF (petrol prices)

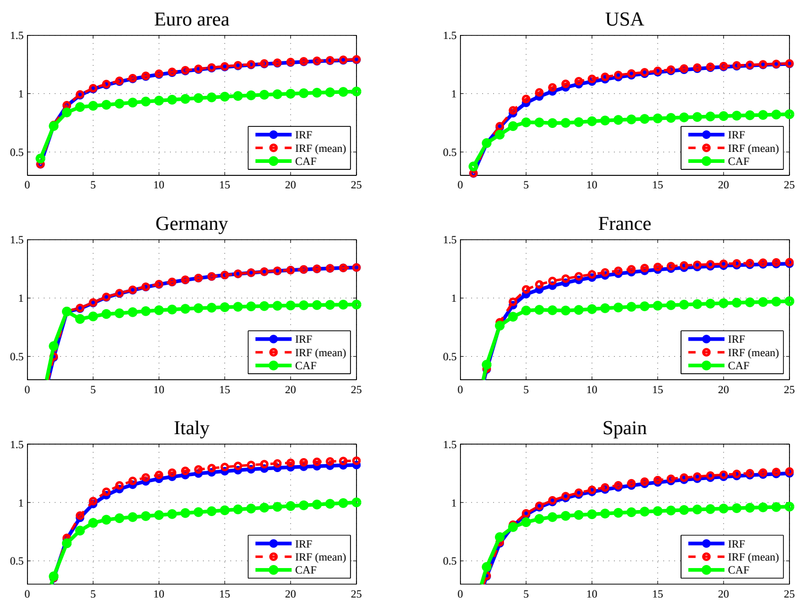

Figure 21: Correct and approximate IRF (diesel prices)

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2007

S. SIVIERO and D. TERLIZZESE, *Macroeconomic forecasting: Debunking a few old wives' tales*, Journal of Business Cycle Measurement and Analysis , v. 3, 3, pp. 287-316, **TD No. 395 (February 2001).**

S. MAGRI, *Italian households' debt: The participation to the debt market and the size of the loan*, Empirical Economics, v. 33, 3, pp. 401-426, **TD No. 454 (October 2002).**

L. CASOLARO. and G. GOBBI, *Information technology and productivity changes in the banking industry*, Economic Notes, Vol. 36, 1, pp. 43-76, **TD No. 489 (March 2004).**

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M. PAIELLA, *Does wealth affect consumption? Evidence for Italy*, Journal of Macroeconomics, Vol. 29, 1, pp. 189-205, **TD No. 510 (July 2004).**

F. LIPPI. and S. NERI, *Information variables for monetary policy in a small structural model of the euro area*, Journal of Monetary Economics, Vol. 54, 4, pp. 1256-1270, **TD No. 511 (July 2004).**

A. ANZUINI and A. LEVY, *Monetary policy shocks in the new EU members: A VAR approach*, Applied Economics, Vol. 39, 9, pp. 1147-1161, **TD No. 514 (July 2004).**

D. JR. MARCHETTI and F. Nucci, *Pricing behavior and the response of hours to productivity shocks*, Journal of Money Credit and Banking, v. 39, 7, pp. 1587-1611, **TD No. 524 (December 2004).**

R. BRONZINI, *FDI Inflows, agglomeration and host country firms' size: Evidence from Italy*, Regional Studies, Vol. 41, 7, pp. 963-978, **TD No. 526 (December 2004).**

L. MONTEFORTE, *Aggregation bias in macro models: Does it matter for the euro area?*, Economic Modelling, 24, pp. 236-261, **TD No. 534 (December 2004).**

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A. DALMAZZO and G. DE BLASIO, *Production and consumption externalities of human capital: An empirical study for Italy*, Journal of Population Economics, Vol. 20, 2, pp. 359-382, **TD No. 554 (June 2005).**

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L. GAMBACORTA and S. IANNOTTI, *Are there asymmetries in the response of bank interest rates to monetary shocks?*, Applied Economics, v. 39, 19, pp. 2503-2517, **TD No. 566 (November 2005).**

P. ANGELINI and F. LIPPI, *Did prices really soar after the euro cash changeover? Evidence from ATM withdrawals*, International Journal of Central Banking, Vol. 3, 4, pp. 1-22, **TD No. 581 (March 2006).**

A. LOCARNO, *Imperfect knowledge, adaptive learning and the bias against activist monetary policies*, International Journal of Central Banking, v. 3, 3, pp. 47-85, **TD No. 590 (May 2006).**

F. LOTTI and J. MARCUCCI, *Revisiting the empirical evidence on firms' money demand*, Journal of Economics and Business, Vol. 59, 1, pp. 51-73, **TD No. 595 (May 2006).**

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A. BRANDOLINI, *Measurement of income distribution in supranational entities: The case of the European Union*, in S. P. Jenkins e J. Micklewright (eds.), Inequality and Poverty Re-examined, Oxford, Oxford University Press, **TD No. 623 (April 2007).**

M. PAIELLA, *The foregone gains of incomplete portfolios*, Review of Financial Studies, Vol. 20, 5, pp. 1623-1646, **TD No. 625 (April 2007).**

K. BEHRENS, A. R. LAMORGESE, G.I.P. OTTAVIANO and T. TABUCHI, *Changes in transport and non transport costs: local vs. global impacts in a spatial network*, Regional Science and Urban Economics, Vol. 37, 6, pp. 625-648, **TD No. 628 (April 2007).**

M. BUGAMELLI, *Prezzi delle esportazioni, qualità dei prodotti e caratteristiche di impresa: analisi su un campione di imprese italiane,* v. 34, 3, pp. 71-103, Economia e Politica Industriale, **TD No. 634 (June 2007).**

G. ASCARI and T. ROPELE, *Optimal monetary policy under low trend inflation*, Journal of Monetary Economics, v. 54, 8, pp. 2568-2583, **TD No. 647 (November 2007).**

R. GIORDANO, S. MOMIGLIANO, S. NERI and R. PEROTTI, *The Effects of Fiscal Policy in Italy: Evidence from a VAR Model*, European Journal of Political Economy, Vol. 23, 3, pp. 707-733, **TD No. 656 (January 2008).**

B. ROFFIA and A. ZAGHINI, *Excess money growth and inflation dynamics*, International Finance, v. 10, 3, pp. 241-280, **TD No. 657 (January 2008).**

G. BARBIERI, P. CIPOLLONE and P. SESTITO, *Labour market for teachers: demographic characteristics and allocative mechanisms,* Giornale degli economisti e annali di economia, v. 66, 3, pp. 335-373, **TD No. 672 (June 2008).**

E. BREDA, R. CAPPARIELLO and R. ZIZZA, *Vertical specialisation in Europe: evidence from the import content of exports*, Rivista di politica economica, v. 97, 3, pp. 189, **TD No. 682 (August 2008).**

2008

P. ANGELINI, *Liquidity and announcement effects in the euro area*, Giornale degli Economisti e Annali di Economia, v. 67, 1, pp. 1-20, **TD No. 451 (October 2002).**

P. ANGELINI, P. DEL GIOVANE, S. SIVIERO and D. TERLIZZESE, *Monetary policy in a monetary union: What role for regional information?*, International Journal of Central Banking, v. 4, 3, pp. 1-28, **TD No. 457 (December 2002).**

F. SCHIVARDI and R. TORRINI, *Identifying the effects of firing restrictions through size-contingent Differences in regulation*, Labour Economics, v. 15, 3, pp. 482-511, **TD No. 504 (June 2004).**

L. GUISO and M. PAIELLA,, *Risk aversion, wealth and background risk*, Journal of the European Economic Association, v. 6, 6, pp. 1109-1150, **TD No. 483 (September 2003).**

C. BIANCOTTI, G. D'ALESSIO and A. NERI, *Measurement errors in the Bank of Italy's survey of household income and wealth*, Review of Income and Wealth, v. 54, 3, pp. 466-493, **TD No. 520 (October 2004).**

S. MOMIGLIANO, J. HENRY and P. HERNÁNDEZ DE COS, *The impact of government budget on prices: Evidence from macroeconometric models*, Journal of Policy Modelling, v. 30, 1, pp. 123-143 **TD No. 523 (October 2004).**

L. GAMBACORTA, *How do banks set interest rates?*, European Economic Review, v. 52, 5, pp. 792-819, **TD No. 542 (February 2005).**

P. ANGELINI and A. GENERALE, *On the evolution of firm size distributions*, American Economic Review, v. 98, 1, pp. 426-438, **TD No. 549 (June 2005).**

R. FELICI and M. PAGNINI, *Distance, bank heterogeneity and entry in local banking markets*, The Journal of Industrial Economics, v. 56, 3, pp. 500-534, **No. 557 (June 2005).**

S. DI ADDARIO and E. PATACCHINI, *Wages and the city. Evidence from Italy*, Labour Economics, v.15, 5, pp. 1040-1061, **TD No. 570 (January 2006).**

M. PERICOLI and M. TABOGA, *Canonical term-structure models with observable factors and the dynamics of bond risk premia*, Journal of Money, Credit and Banking, v. 40, 7, pp. 1471-88, **TD No. 580 (February 2006).**

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S. FEDERICO and G. A. MINERVA, *Outward FDI and local employment growth in Italy,* Review of World Economics, Journal of Money, Credit and Banking, v. 144, 2, pp. 295-324, **TD No. 613 (February 2007).**

F. BUSETTI and A. HARVEY, *Testing for trend*, Econometric Theory, v. 24, 1, pp. 72-87, **TD No. 614 (February 2007).**

V. CESTARI, P. DEL GIOVANE and C. ROSSI-ARNAUD, *Memory for prices and the Euro cash changeover: an analysis for cinema prices in Italy,* In P. Del Giovane e R. Sabbatini (eds.), The Euro Inflation and Consumers' Perceptions. Lessons from Italy, Berlin-Heidelberg, Springer, **TD No. 619 (February 2007).**

B. H. HALL, F. LOTTI and J. MAIRESSE, *Employment, innovation and productivity: evidence from Italian manufacturing microdata*, Industrial and Corporate Change, v. 17, 4, pp. 813-839, **TD No. 622 (April 2007).**

J. SOUSA and A. ZAGHINI, *Monetary policy shocks in the Euro Area and global liquidity spillovers,* International Journal of Finance and Economics, v.13, 3, pp. 205-218, **TD No. 629 (June 2007).**

M. DEL GATTO, GIANMARCO I. P. OTTAVIANO and M. PAGNINI, *Openness to trade and industry cost dispersion: Evidence from a panel of Italian firms,* Journal of Regional Science, v. 48, 1, pp. 97-129, **TD No. 635 (June 2007).**

P. DEL GIOVANE, S. FABIANI and R. SABBATINI, *What's behind "inflation perceptions"? A survey-based analysis of Italian consumers,* in P. Del Giovane e R. Sabbatini (eds.), *The Euro Inflation and Consumers' Perceptions. Lessons from Italy,* Berlin-Heidelberg, Springer, **TD No. 655 (January 2008).**

R. BRONZINI, G. DE BLASIO, G. PELLEGRINI and A. SCOGNAMIGLIO, *La valutazione del credito d'imposta per gli investimenti,* Rivista di politica economica, v. 98, 4, pp. 79-112, **TD No. 661 (April 2008).**

B. BORTOLOTTI, and P. PINOTTI, *Delayed privatization,* Public Choice, v. 136, 3-4, pp. 331-351, **TD No. 663 (April 2008).**

R. BONCI and F. COLUMBA, *Monetary policy effects: New evidence from the Italian flow of funds,* Applied Economics , v. 40, 21, pp. 2803-2818, **TD No. 678 (June 2008).**

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A. SILVESTRINI and D. VEREDAS, *Temporal aggregation of univariate and multivariate time series models: a survey,* Journal of Economic Surveys, v. 22, 3, pp. 458-497, **TD No. 685 (August 2008).**

2009

F. PANETTA, F. SCHIVARDI and M. SHUM, *Do mergers improve information? Evidence from the loan market*, Journal of Money, Credit, and Banking, v. 41, 4, pp. 673-709, **TD No. 521 (October 2004).**

M. BUGAMELLI and F. PATERNÒ, *Do workers' remittances reduce the probability of current account reversals?*, World Development, v. 37, 12, pp. 1821-1838, **TD No. 573 (January 2006).**

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M. PERICOLI and M. SBRACIA, *The CAPM and the risk appetite index: theoretical differences, empirical similarities, and implementation problems*, International Finance, v. 12, 2, pp. 123-150, **TD No. 586 (March 2006).**

U. ALBERTAZZI and L. GAMBACORTA, *Bank profitability and the business cycle,* Journal of Financial Stability, v. 5, 4, pp. 393-409, **TD No. 601 (September 2006).**

S. MAGRI, *The financing of small innovative firms: the Italian case*, Economics of Innovation and New Technology, v. 18, 2, pp. 181-204, **TD No. 640 (September 2007).**

S. MAGRI, *The financing of small entrepreneurs in Italy*, Annals of Finance, v. 5, 3-4, pp. 397-419, **TD No. 640 (September 2007).**

V. DI GIACINTO and G. MICUCCI, *The producer service sector in Italy: long-term growth and its local determinants*, Spatial Economic Analysis, Vol. 4, No. 4, pp. 391-425, **TD No. 643 (September 2007).**

F. LORENZO, L. MONTEFORTE and L. SESSA, *The general equilibrium effects of fiscal policy: estimates for the euro area*, Journal of Public Economics, v. 93, 3-4, pp. 559-585, **TD No. 652 (November 2007).**

R. GOLINELLI and S. MOMIGLIANO, *The Cyclical Reaction of Fiscal Policies in the Euro Area. A Critical Survey of Empirical Research*, Fiscal Studies, v. 30, 1, pp. 39-72, **TD No. 654 (January 2008).**

P. DEL GIOVANE, S. FABIANI and R. SABBATINI, *What's behind "Inflation Perceptions"? A survey-based analysis of Italian consumers,* Giornale degli Economisti e Annali di Economia, v. 68, 1, pp. 25-52, **TD No. 655 (January 2008).**

F. MACCHERONI, M. MARINACCI, A. RUSTICHINI and M. TABOGA, *Portfolio selection with monotone mean-variance preferences,* Mathematical Finance, v. 19, 3, pp. 487-521, **TD No. 664 (April 2008).**

M. AFFINITO and M. PIAZZA, *What are borders made of? An analysis of barriers to European banking integration*, in P. Alessandrini, M. Fratianni and A. Zazzaro (eds.): The Changing Geography of Banking and Finance, Dordrecht Heidelberg London New York, Springer, **TD No. 666 (April 2008).**

G. FERRERO and A. NOBILI, *Futures contract rates as monetary policy forecasts,* International Journal of Central Banking, v. 5, 2, pp. 109-145, **TD No. 681 (June 2008).**

L. ARCIERO, C. BIANCOTTI, L. D'AURIZIO and C. IMPENNA, *Exploring agent-based methods for the analysis of payment systems: A crisis model for StarLogo TNG,* Journal of Artificial Societies and Social Simulation, v. 12, 1, **TD No. 686 (August 2008).**

A. CALZA and A. ZAGHINI, *Nonlinearities in the dynamics of the euro area demand for M1*, Macroeconomic Dynamics, v. 13, 1, pp. 1-19, **TD No. 690 (September 2008).**

L. FRANCESCO and A. SECCHI, *Technological change and the households' demand for currency,* Journal of Monetary Economics, v. 56, 2, pp. 222-230, **TD No. 697 (December 2008).**

G. ASCARI and T. ROPELE, *Trend inflation, taylor principle, and indeterminacy*, Journal of Money, Credit and Banking, v. 41, 8, pp. 1557-1584, **TD No. 708 (May 2007).**

S. COLAROSSI and A. ZAGHINI, *Gradualism, transparency and the improved operational framework: a look at overnight volatility transmission*, International Finance, v. 12, 2, pp. 151-170, **TD No. 710 (May 2009).**

M. BUGAMELLI, F. SCHIVARDI and R. ZIZZA, *The euro and firm restructuring*, in A. Alesina e F. Giavazzi (eds): Europe and the Euro, Chicago, University of Chicago Press, **TD No. 716 (June 2009).**

B. HALL, F. LOTTI and J. MAIRESSE, *Innovation and productivity in SMEs: empirical evidence for Italy,* Small Business Economics, v. 33, 1, pp. 13-33, **TD No. 718 (June 2009).**

2010

S. MAGRI, *Debt maturity choice of nonpublic Italian firms* , Journal of Money, Credit, and Banking, v.42, 2-3, pp. 443-463, **TD No. 574 (January 2006).**

FORTHCOMING

L. MONTEFORTE and S. SIVIERO, *The Economic Consequences of Euro Area Modelling Shortcuts*, Applied Economics, **TD No. 458 (December 2002).**

M. BUGAMELLI and A. ROSOLIA, *Produttività e concorrenza estera*, Rivista di politica economica, **TD No. 578 (February 2006).**

G. DE BLASIO and G. NUZZO, *Historical traditions of civicness and local economic development,* Journal of Regional Science, **TD No. 591 (May 2006).**

R. BRONZINI and P. PISELLI, *Determinants of long-run regional productivity with geographical spillovers: the role of R&D, human capital and public infrastructure,* Regional Science and Urban Economics, **TD No. 597 (September 2006).**

E. IOSSA and G. PALUMBO, *Over-optimism and lender liability in the consumer credit market,* Oxford Economic Papers, **TD No. 598 (September 2006).**

S. NERI and A. NOBILI, *The transmission of US monetary policy to the euro area,* International Finance, **TD No. 606 (December 2006).**

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A. CIARLONE, P. PISELLI and G. TREBESCHI, *Emerging Markets' Spreads and Global Financial Conditions*, Journal of International Financial Markets, Institutions & Money, **TD No. 637 (June 2007).**

Y. ALTUNBAS, L. GAMBACORTA and D. MARQUÉS, *Securitisation and the bank lending channel,* European Economic Review, **TD No. 653 (November 2007).**

M. IACOVIELLO and S. NERI, *Housing market spillovers: evidence from an estimated DSGE model,* American Economic Journal: Macroeconomics, **TD No. 659 (January 2008).**

F. BALASSONE, F. MAURA and S. ZOTTERI, *Cyclical asymmetry in fiscal variables in the EU*, Empirica, **TD No. 671 (June 2008).**

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F. D'AMURI, O. GIANMARCO I.P. and P. GIOVANNI, *The labor market impact of immigration on the western german labor market in the 1990s,* European Economic Review, **TD No. 687 (August 2008).**

A. ACCETTURO, *Agglomeration and growth: the effects of commuting costs,* Papers in Regional Science, **TD No. 688 (September 2008).**

L. FORNI, A. GERALI and M. PISANI, *Macroeconomic effects of greater competition in the service sector: the case of Italy,* Macroeconomic Dynamics, **TD No. 706 (March 2009).**

Y. ALTUNBAS, L. GAMBACORTA, and D. MARQUÉS-IBÁÑEZ, *Bank risk and monetary policy,* Journal of Financial Stability, **TD No. 712 (May 2009).**

V. DI GIACINTO, G. MICUCCI and P. MONTANARO, *Dynamic macroeconomic effects of public capital: evidence from regional Italian data,* Giornale degli economisti e annali di economia, **TD No. 733 (November 2009).**

F. COLUMBA, L. GAMBACORTA and P. E. MISTRULLI, *Mutual Guarantee institutions and small business finance,* Journal of Financial Stability, **TD No. 735 (November 2009).**

L. FORNI, A. GERALI and M. PISANI, *The macroeconomics of fiscal consolidations in euro area countries,* Journal of Economic Dynamics and Control, **TD No. 747 (March 2010).**