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DO FOOD COMMODITY PRICES HAVE ASYMMETRIC EFFECTS ON EURO-AREA INFLATION?

by Mario Porqueddu* and Fabrizio Venditti*

Abstract

This paper analyzes the relationship between commodity prices and consumer food prices in the euro area and in its largest economies (Germany, France and Italy) and tests whether the latter respond asymmetrically to shocks to the former. The issue is of particular interest for those monetary authorities that target headline consumer price inflation, which has been heavily influenced by pronounced swings in international commodity prices in the past decade. The empirical analysis is based on two distinct but complementary approaches. First, we employ a structural model to identify a shock to commodity prices and verify using formal econometric tests whether the Impulse Response Functions of food consumer prices is invariant to the sign of the commodity price shock. Next, we employ predictive regressions and examine the relative forecasting ability of linear models compared with that of models that allow for sign-dependent nonlinearities. Overall, the empirical analysis uncovers very little evidence of asymmetries.

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

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1 Introduction

The relationship between commodity price fluctuations and macroeconomic variables has attracted considerable interest in the past decade, fostered by the sharp rise in commodity prices observed between 2003 and 2008 and then their pronounced volatility which followed the financial crisis. Besides oil and metals, food prices recorded an unprecedented surge, after a prolonged period of relative stability (Figure 1).

At the root of the food commodity boom lay a number of factors from both the demand and the supply side (IMF, 2008). On the demand side, a significant contribution came from the rapid increase in per capita income in the emerging economies. Considering that the income elasticity of food demand is much larger in developing than in developed countries this channel boosted the demand for food more than proportionally. A composition effect was also at play, since higher living standards imply a shift in the nutrition pattern towards a higher protein diet with a multiplier effect on the demand for grains that are used for livestock feeding. On the supply side, the key factors were the protracted adverse weather conditions in some important exporters of wheat like Australia, and a decrease in productivity, mainly reflecting the secular fall in the real price of agricultural output, which, in turn, had lowered incentives for investing in this sector. According to some commentators loose monetary policy conditions and the increased complexity of financial markets, which allowed traders to exchange commodities also in future markets, played a decisive role, although the empirical evidence supporting this argument is rather weak (Anzuini et al., 2012).

The persistent inflationary pressure exerted by food commodity prices raised concerns for those monetary authorities whose objective is stated in terms of headline inflation. The relevance of these shocks was instead downplayed by those who claimed that monetary policy should look at a "core" measure of inflation, typically defined as the headline index net of food and energy, items whose price fluctuations are mostly seen as volatile and short-lived. This view found both advocates (Evans, 2011, Rosengren, 2011) as well as fierce opponents (Bullard, 2011) within the U.S. Federal Reserve.[1] Core inflation measures obtained by focusing on a subset of the consumption basket have, however,

[1] President Trichet explicitly made this distinction between the ECB and the Fed attitude towards core inflation in an interview with the Wall Street Journal in January 2011: "In the U.S. the Fed considers that core inflation is a good predictor for future headline inflation. In our case we consider that core inflation is not necessarily a good predictor for future headline inflation." See `http://www.ecb.int/press/key/date/2011/html/sp110123.en.html`, accessed on the 16*th* of April 2012.

some major shortcomings. First, they do not appear to be good predictors of headline inflation, and therefore cannot serve as a useful intermediate target to control inflation in the medium term (Crone et al., 2011). Second, insofar as the recent trend in commodity prices reflects growing demand from emerging countries, it could be a secular issue rather than a short-term phenomenon, playing an important role for global inflation for many years ahead. Third, theoretical macro models in which food plays a distinctive role in the utility function of the agents and in which the variance of food commodity prices is realistically large, suggest that monetary policies that target headline inflation are welfare-superior (Catao and Chang, 2010). In a historical shift, the Fed has recently taken a clear stand in this debate, getting closer to the position held by the ECB, and quantified a long run inflation target of 2% defined in terms of the annual change in the price index for personal consumption expenditures.[2]

Contrary to the effect of oil prices, which has attracted a large number of studies since the first oil shock of the Seventies, the impact of food prices on the macroeconomy of developed countries has been much less investigated, partly owing to the fact that until the last decade the real price of agricultural output had constantly fallen, driven by persistent productivity gains. Moreover, existing empirical studies that aimed at quantifying the pass-through from upstream to downstream prices reached mixed conclusions. In particular, while it was possible to establish a robust statistical link between commodity and retail food prices in most countries outside the euro area, results for the euro area pointed to a rather weak correlation between upstream and downstream food prices (IMF, 2008). However, a study by the National Bank of Belgium (2008, henceforth NBB), pointed out that in Europe the effect of international agricultural price fluctuations might have been shadowed by the subsidies granted to domestic producers by the Common Agricultural Policy (CAP). These subsidies, acting as a wedge between international and domestic prices, partly mitigate the transmission of international raw material shocks to consumer prices. In turn, once internal EU prices are used, instead of the index of food commodity prices on the world market calculated by the Hamburg Institute of International Economics (HWWI) customarily employed in previous studies, a robust relationship between

[2]"The Committee judges that inflation at the rate of 2 percent, as measured by the annual change in the price index for personal consumption expenditures, is most consistent over the longer run with the Federal Reserve's statutory mandate. Communicating this inflation goal clearly to the public helps keep longer-term inflation expectations firmly anchored.", from minutes of the meeting of the FOMC meeting held on the 25*th* of January 2012, available at `http://www.federalreserve.gov/monetarypolicy/fomcminutes20120125.htm`, accessed on the 12*th* of April 2012.

commodity and euro area food consumer prices emerges. Using the same dataset as NBB, Ferrucci, Jimeneze-Rodriguez and Onorante (2012, henceforth FJRO) further analyze the speed and magnitude of the pass-through from raw materials to retail prices in the euro area. In particular they investigate the hypothesis that the transmission to food consumer prices is asymmetric, i.e. that positive upstream shocks generate a faster and stronger downstream response than negative ones. They conclude that "...asymmetries and non-linearities are statistically and economically significant, and hence have to be accounted for in order to precisely measure the impact of a commodity price shock on consumer prices". This conclusion is supported solely by an Information Criterion argument, i.e. models that include some form of asymmetry in the relationship between consumer and commodity prices are found to score better than linear models in terms of penalized measures of fit.

In this paper we revisit the empirical link between commodity and consumer food prices in the euro area with a focus on the issue of possible asymmetries in the transmission from upstream shocks to downstream prices. Our paper presents three novel contributions to the literature. First, extending the analysis in FJRO, we look not only at the euro area as a whole but also at the three largest countries (France, Germany and Italy), in order to uncover any cross-country heterogeneity in the transmission of food shocks. This might be an important input for national regulatory authorities since differences across countries might indicate the presence of some form of collusive behavior in a given market. Second, rather than relying on Information Criteria, we base our inference on recently developed econometric tools that are explicitly designed to test for this form of nonlinearities. Our third contribution is methodological, as we propose a linearized version of the nonlinear Impulse Response Function (IRF) used by Kilian and Vigfusson (2011a, 2011b) to test for nonlinearities in structural models with asymmetric terms. In a small Monte Carlo study we show that our linearized IRF matches very closely the nonlinear one while reducing computational time by an order of magnitude. Overall, our results indicate that, when the appropriate econometric methods are used, it is very hard to argue for any form of asymmetry in the response of retail to commodity prices in the food sector. The paper is structured as follows. Section 2 reviews more in detail the literature on asymmetric price adjustments. Section 3 describes the data. Section 4 reports some preliminary evidence on the differences across countries of the speed of adjustment of consumer prices to shocks to raw material prices, based on a linear model. Section 5 discusses the nonlinear IRF based tests, section 6 presents the forecasting exercise and section 7 concludes.

2 Literature review

The wealth of empirical studies on asymmetries price adjustment contrasts starkly with the lack of a clear theoretical justification for why prices should respond more strongly to cost increases than to cost decreases. Often quoted reasons, like adjustment costs (menu costs, search costs, implicit contracts and so forth) rationalize in fact the existence of a lag in the transmission from marginal costs to prices, but do not imply any asymmetry, a point stressed by Peltzman (2000). Other formal motives for the existence of asymmetries include collusion among market players (Balke et al., 1998), imperfect competition (Borenstein et al., 1997) and inventory management (Reagan and Weitzman, 1982). The weakness of these justifications is discussed at length in Peltzman (2000), Manera and Frey (2007) and Meyer and von Cramon-Taubadel (2004). A stronger theoretical underpinning for asymmetric price adjustment has been recently provided by L'Huillier (2012), who builds a model in which firms are better informed than consumers about the nature of the aggregate shocks and an exchange of goods occurs only after consumers have engaged in a costly search. In this setup an asymmetric response of prices to marginal cost shocks emerges as an optimal pricing behavior.

In contrast with the scarcity of theoretical papers, the empirical evidence on the existence of asymmetries in the food pricing chain is quite rich. Aguiar and Santana (2002) use monthly data to study the pass-through from farm to consumer prices in Brazil. They find significant evidence of asymmetries in most markets. In a report by London Economics (2003), focussing on the relationship between producer and consumer of prices of a number of unprocessed and processed food products in Austria, Denmark, France, Germany, Ireland, Netherlands, Spain and UK, evidence of asymmetries is found mainly in the dairy sector, where asymmetric transmission from producer to consumer prices occurs in the UK, France and Denmark. Similar results are reported by Zachariasse and Bunte (2003) for meat products in the Netherlands. In a widely quoted study, Peltzman (2000) analyzes the producer and retail prices of more than a hundred agricultural and food items, finding that asymmetric transmission is a pervasive feature of prices setting, both at the producer and consumer level.

From a methodological point of view empirical models of asymmetric transmission have recently come under scrutiny in the literature. Two important contributions have shaped this debate.

The first is a study by Kilian and Vigfusson (2011a, henceforth KV), which revisits the issue of the effect of energy shocks on the U.S. macroeconomy, questioning the methodology employed in all previous empirical studies. To understand their argument consider the case in which a researcher wants to test the hypothesis that a variable y_t responds asymmetrically to the variable x_t, so that a positive change in x_t, defined as $x_t^+ = max\{0, x_t\}$ generates a stronger response than a negative one $x_t^- = min\{0, x_t\}$. In most studies this question is tackled by estimating a dynamic model in which y_t is regressed on its own lags, lags of x_t and lags of x_t^+.[3] Evidence in favor or against asymmetries is then provided by testing that the coefficients associated with the positive changes x_t^+ are not significantly different from zero. KV argue that such a test cannot provide a convincing case either in favor or against asymmetries because the presence of censored endogenous variables makes the model nonlinear.[4] In this setup there is no clear correspondence between the slopes of the estimated equations and the dynamic response of the endogenous variables to an exogenous shock, since small asymmetries in the coefficients might progressively accumulate and result in significant asymmetries in the response. Furthermore, different linear combinations of these coefficients can be subject to a statistical test. Some researchers, for example, test the null hypothesis that these coefficients are jointly zero, other that they sum to zero and so forth. In their review of the literature Manera and Frey (2007) count up to six different ways to test for asymmetries. The weakness of such an empirical strategy is quite clear as one can reach different conclusions based on different tests. Not surprisingly, in their review of the literature of asymmetric price transmission in agricultural markets Meyer and von Cramon-Taubadel (2004) lament the lack of a unified framework for testing the hypothesis of asymmetric price adjustment. KV provide such a unified framework by proposing to investigate the issue of asymmetric responses in a *structural* setup. In practice their proposal consists of writing down a general model of the price transmission in which (i) all the variables of interest are modeled jointly and (ii) some restrictions are imposed on the dynamics of the model so that it is possible to identify a structural commodity price shock and to derive the Impulse Response Functions (IRFs) of the endogenous variables to such a shock. A formal test on the shape of these responses can then be performed. An IRF based asymmetry test, in particular, consists of verifying the hypothesis that the response to a positive structural shock is significantly different from the response to a negative one. Such testing procedure is used by Kilian and Vigufsson (2011b), who cannot reject the hypothesis linearity of the relationship be-

[3]Including x_t and x_t^+ in the regression is equivalent to including x_t^+ and x_t^-

[4]Such a model can be seen as a threshold model with the threshold set at 0.

tween oil and GDP in the U.S., by Herrera et al. (2011) in the context of the relationship between oil prices and sectoral industrial production in the U.S., and in Venditti (2010) for studying the pass-through from oil to liquid fuels prices in the U.S. and in the euro area. It presents a clear advantage over tests based on the regression coefficients, since it provides a univocal answer to the question of whether the pass-through from upstream to downstream prices is asymmetric. It has, however, one major shortcoming since it is quite computationally intensive, see Section 5 for further details.

A second important contribution is the reply to KV by Hamilton (2011). He argues that the need to identify a structural shock in the testing procedure advocated by KV leads to unnecessary restrictions on the contemporaneous relationship between y_t and x_t. If asymmetries are really present then reduced form models should be able to detect them. In particular, if the censored terms x_t^+ are informative for the dynamics of y_t a forecasting model that includes these terms should forecast future values of y_t better than a model that excludes them. Asymmetries can then be tested for by comparing the forecast accuracy of models that include nonlinear terms relative to that of a standard linear model. Using this strategy, for example, Hamilton (2011) finds that augmenting a linear VAR with oil price increases leads to more accurate predictions of the U.S. GDP. The appealing aspect of the forecast based asymmetry test proposed by Hamilton (2011) is its computational simplicity. On the other hand, in small samples the results of forecast based tests might be affected by parsimony issues. Models with asymmetric terms, in fact, nest models that exclude them and are therefore more heavily parameterized. Greater parsimony might therefore give linear models an edge in a forecast competition, regardless of whether *in population* asymmetric terms are part of the data generating process.

Since both KV and Hamilton (2011) approaches have their own merits, in our empirical analysis we take an eclectic view and use both testing procedures to examine whether food commodity prices affect asymmetrically retail prices. First we estimate a bivariate model of commodity and consumer food prices, identify a structural shock to commodity prices and use Impulse Response Function (IRF) based asymmetry tests to investigate the hypothesis that consumer prices respond more strongly to positive than to negative shocks. Rather than using the computationally intensive procedure advocated by KV, we propose a linearized version of the IRF and show in a simple simulation study that our simplified algorithm produces IRF that match closely the original ones, reducing computational time substantially. We then turn to forecast based tests and design a pseudo out of sample forecast exercise in which we use multistep predictive regressions to analyze the relative predictive accuracy of models with or without asymmetric terms at different forecast horizons.

3 The Data

Our dataset is composed of monthly data data from January 1997 to September 2011. Starting from consumer prices, we use series of the Harmonised Index of Consumer Prices (HICP) provided by Eurostat for four categories. We look at aggregate *processed food prices net of alcohol and tobacco*[5] and at a subset of its underlying sub-components, namely *bread and cereals*, *milk cheese and eggs*, *oils and fats* and *coffee tea and cocoa*. We exclude from the analysis some items, namely *sugar, jam, honey, syrups, etc.*, *food products n.e.c.* and *mineral water and soft drinks*. The *sugar, jam, honey, syrups, etc.* item could, in principle, be relevant for our analysis since it could be affected by international sugar prices. In practice the results presented by FJRO indicate that it is not possible to find a significant relationship between this food category and sugar commodity prices.[6] A plausible explanation for this result is that this specific category collects products whose inputs include many other raw materials whose weight in the structure of the production costs is likely to be larger than that of sugar.

Table 1 shows the weights of these items in the HICP basket for the euro area, France, Germany and Italy. Overall, processed food account for around 8-9% of the consumption basket in these countries. Over half of this is represented by *bread and cereals* and *milk cheese and eggs*. *Oils and fats* and *coffee tea and cocoa* account for a much smaller fraction of the consumption basket. There are no stark differences across countries, although Italian households allocate a relatively slightly higher share of their expenditure to the consumption of *bread and cereals* and *oils and fats*.

In the remainder of the paper in order to simplify the taxonomy of the HICP categories we will refer to *processed food prices*, *bread and cereals*, *milk cheese and eggs*, *oils and fats* and *coffee tea and cocoa* simply as food, cereals, dairy, oils and coffee.

Turning to food commodity prices, for cereals, milk and oil we rely on the price statistics collected by the European Commission (the same data sources used by FJRO and NBB), which explicitly account for the role of the CAP. For coffee, whose prices are not subject to the CAP, we use instead internationally quoted prices. We also construct

[5]In what follows we will refer to this category simply as processed food prices.

[6]We replicated FJRO analysis and reached the same conclusions.

an aggregate commodity food series as a weighted average of the relevant commodity prices using the corresponding HICP weights displayed in Table 1. All commodity prices are expressed in euros.

4 The baseline model

Before moving to the main empirical analysis where we test for asymmetric effects we run a preliminary exercise to check whether there are significant differences across euro area countries in the pass-through of commodity price shocks to consumer prices. We model the dynamic relationship between food commodity and consumer prices[7] with a linear structural VAR and derive the IRFs to a 10% commodity price shock for all the countries and the products considered.

Commodity price shocks are identified by assuming that an unexpected variation in the price of raw materials is exogenous relative to the contemporaneous values of the consumer prices included in the VAR. We are therefore imposing the restriction that consumer prices do not impact *contemporaneously* on commodity prices but that the latter can respond to the former only with a lag. This assumption is quite standard in studies of the pass-through of upstream price shocks along the pricing chain. It has been employed, for example, by Hahn (2003) to examine the transmission of external shocks to euro area inflation and also by NBB and FJRO in their studies of the pass-through from commodity to consumer food prices. In the context of the transmission of oil price shocks to the macroeconomy recent studies that postulate this identifying restriction include, among many others, Blanchard and Galí (2009), Kilian and Park (2008) and Blanchard and Riggi (2012).

One difference worth remarking between our approach and the one followed by NBB and FJRO is that, whereas we rely on a simpler bivariate model, they work with a richer structure, which includes not only commodity and consumer prices but also producer prices. Since, however, we (and also NBB and FJRO) assume that food commodity prices are predetermined (i.e. contemporaneously exogenous) with respect to all included and excluded macroeconomic variables, a bivariate model is not sensitive to the omitted variable bias and delivers fully consistent IRFs. The intuition behind this result, discussed at length in Kilian (2011, section 2.3.4), is that since commodity prices are ordered first in the recursive VAR, the identification of a commodity price shock does not depend

[7] Both food and commodity prices are log-differenced

on the contemporaneous correlation of any of the reduced form shocks in the system, regardless of the number of variables included in the VAR. Things are different, instead, in recursive VAR models used to identify a monetary policy shock where the interest rate reacts *contemporaneously* to all the structural shocks in the system (i.e the interest rate is ordered as the last variable of the system). In these models omitting a relevant variable hinders the identification of the structural shock and produces incorrect IRFs.

The estimated IRFs are shown in figure 2.[8] The main findings on the effect of an unexpected 10% shock to raw material prices can be summarized as follows.

- Food consumer prices in the euro area are strongly affected by commodity prices. Following the shock to raw material prices, they gradually increase, reaching a cumulated long-run response of around 1.5%. This result masks a significant degree of heterogeneity across countries: the long-run reaction of downstream prices in Germany (2%) is, respectively, twice and four times as strong as that observed in Italy (1%) and in France (0.5%). The response of consumer prices is very protracted, as it takes around a year for the transmission of the shock to be completed.

- The response of retail prices to wheat and coffee commodity prices is rather homogenous across countries. In the case of wheat prices, two years after a shock consumer prices rise by around above 1% in the euro area and in Italy, slightly less than 1% in Germany and around 4 decimal points in France. The pass-through of a coffee price shocks is slightly more muted as the cumulated impact after two years lies between 0.5 and 1% in Italy, Germany and the euro area. No significant reaction is instead found in France.

- The responses of dairy and oils are markedly different across countries. Dairy consumer prices increase by 4% in the euro area, with responses at the country level ranging from zero (in France and in Italy) to 10% (in Germany). In the case of oils the effect of a raw material price shock on retail prices amounts to 3% in the euro area. The reaction is again stronger in Germany (5%), milder in France (around 1%), and not significantly different from zero in Italy.

Summing up, preliminary evidence obtained in a linear setup confirms that euro area food prices are indeed significantly affected by commodity prices but also highlight considerable heterogeneity at the country/product level. In particular, food prices appear to

[8] Confidence bands are obtained via Monte Carlo simulation.

be much more sensitive to upstream shocks in Germany, especially in the case of oils and dairy products, than in Italy and France. While an investigation of the reasons behind these diverging patterns is beyond the scope of this paper, a recent cross-country analysis (ECB, 2011) suggests that different features of the distribution sector might play a role. In particular, the size of the pass-through in Germany is plausibly related to the pronounced presence of discount shops, which account for around 40% of the retail grocery market. In turn, in France, where hypermarkets dominate and vertical integration is pervasive, cost shocks appear to get buffered by varying profit margins along the whole pricing chain.

5 A structural model of asymmetric price adjustment

Having collected some stylized facts on the pass-through of food commodity prices we augment the baseline model to allow for an asymmetric reaction of downstream prices to upstream shocks.

In line with Herrera et al. (2011) we look at two different measures of asymmetries. The former is the positive change over the previous period, defined as:

$$x_t^{\#} = x_t^1 = max\{0, x_t - x_{t-1}\} \tag{1}$$

where x_t is the log of prices. This transformation, equivalent to right censoring price changes, was first introduced by Mork (1989) and is customarily used in the empirical literature on the asymmetric transmission of oil price shocks. The latter is the net price increase over the previous 12 periods maximum as in Hamilton (1996):

$$x_t^{\#} = x_t^{12} = max\{0, x_t - max(x_{t-1}, x_{t-2}, \ldots, x_{t-12})\} \tag{2}$$

This transformation tries to capture the notion that asymmetries are not systematically embedded in the price setting behavior of retailers, but rather occur only when marginal costs deviate significantly from their recent behavior. By construction the net increase x_t^{12} will therefore take value zero much more often than the simple increase x_t^1, consistently with the idea that nonlinear responses should be observed only in response to shocks drawn from the right tail of the distribution. To clarify this point we apply the above transformations to food commodity prices and show the results in Figure 3. The net increase takes a positive value in about a dozen instances over the sample (around 10%

of the available observations), while simple price increases constitute around half of the available data points. These differences reflect the different notion of asymmetry that these two data transformations aim at capturing: while the net increase separates only persistent and exceptional shocks from the other observations, Mork's measure attributes to every price increase the possibility to trigger a stronger than the average reaction of the variables in the system.

We investigate asymmetries in the context of the following structural model:

$$\Delta x_t = a_x + \sum_{i=1}^{p} b_{11,i}\Delta x_{t-i} + \sum_{i=1}^{p} b_{12,i}\Delta y_{t-i} + \epsilon_{xt} \quad (3)$$

$$\Delta y_t = a_y + \sum_{i=0}^{p} b_{21,i}\Delta x_{t-i} + \sum_{i=1}^{p} b_{22,i}\Delta y_{t-i} + \sum_{i=0}^{p} b^{\#}_{21,i} x^{\#}_{t-i} + \epsilon_{yt} \quad (4)$$

where $\epsilon_{xt} \sim N(0,\sigma_x^2)$, $\epsilon_{yt} \sim N(0,\sigma_y^2)$, $Cov(\epsilon_{xt},\epsilon_{yt}) = 0$. In the empirical analysis x_t is the log of food commodity prices and y_t is the log index of the corresponding HICP item. The key identifying assumption in model (3)-(4) is that upstream prices are predetermined with respect to fuels prices: a shock to the upstream price (ϵ_{xt}) affects within the same month consumer food prices (its impact is $b_{12,0}$ times ϵ_{xt}, $b_{12,0} + b^{+}_{12,0}$ times ϵ_{xt} if $\epsilon_{xt} > 0$), while a shock to consumer food prices (ϵ_{yt}) feeds back to commodity prices only with a lag. This recursive identifying assumption, akin to the Choleski identification used in section (3) in the linear SVAR, also implies that the regression parameters can be consistently estimated equation by equation via OLS since $Cov(x_t,\epsilon_{yt}) = 0$.

In most of the empirical papers on asymmetric price adjustment a test of asymmetric behaviour is cast in terms of an exclusion test on the coefficients associated with the nonlinear terms in equation (4), that is $b^{\#}_{21,0} = b^{\#}_{21,1} = \cdots = b^{\#}_{21,p} = 0$, which can be investigated via a standard likelihood ratio test. Under this restriction the system of equations (3)-(4) reduces to a linear SVAR. The problem with this empirical strategy, as noticed by KV and discussed in section 2, is that although under the null the model is linear, under the alternative the model is nonlinear, so that the hypothesis that the dynamic response to a structural shock depends on the sign of the underlying shock can only be investigated via a full blown simulation of the model. How to perform such a simulation is the topic of the next subsection.

5.1 Impulse Response Function: computation and approximation

In a nonlinear setup like the one described by equations (3)-(4) the IRF at horizon h will depend not only on the parameters of the model (like in a linear SVAR) but also on the size of the initial shock to x_t, on the history of the data and on future shocks to both x_t and y_t. To understand why all these factors matter consider feeding equation (3) with a very large positive shock. For a given x_{t-1} it is very likely that Δx_t will be positive and that $x_t^{\#}$ will be different from zero, affecting downstream prices through the coefficient $b_{21,0}$. Conversely, the smaller the size of the shock the higher the probability that the term $x_t^{\#}$ will be zero, resulting in a more muted response of y_t. This example also shows that the starting point x_{t-1} matters, as the response of y_t depends on how distant from zero is x_{t-1}. If x_{t-1} is negative and large in absolute value, only positive shocks of a very large magnitude will make $x_t^{\#}$ positive while for shocks of a standard size Δx_t will be negative and $x_t^{\#}$ will be zero, with obvious implications for the dynamic response of y_t to upstream shocks. Finally, the size and the sign of future shocks will also affect the likelihood that the censored terms $x_t^{\#}$ will hit the zero bound, therefore switching on and off their effect on the future values of y_t. In this setup the IRF needs to be estimated via simulation methods. KV propose an algorithm to perform such a simulation. Their algorithm requires picking a starting point ($\omega_{t-1} = x_{t-1}, \ldots, x_{t-p}, y_{t-1}, \ldots, y_{t-p}$) and using model (3)-(4) to simulate two paths of the endogenous variables. The first is obtained by hitting both equations with a sequence of random shocks drawn from the empirical distributions of ϵ_{xt} and ϵ_{yt}, keeping the first shock to x_t at a fixed value, say δ. This generates a response of y_t, say y_t^s. The second path is obtained by feeding both equations a different sequence of random shocks, drawn again from the empirical distributions of ϵ_{xt} and ϵ_{yt}. Call the response of y_t in this second case y_t^b. The difference between the two different trajectories $y_t^s - y_t^b$ is the response of y_t to a shock to x_t of size δ and conditional on a given path of future shocks and on a given data history ω_{t-1}. Repeating the exercise for (1) a large number of simulated shocks (2) all the possible starting points ω_{t-1} and averaging across all the resulting responses $y_t^b - y_t^s$, one obtains an IRF that is only conditional on the size of the initial shock: $I_y^h(\delta)$ (to simplify the notation we have omitted the dependence on the parameters of the model).

Once IRF to positive and negative shocks have been computed, a formal IRF based asymmetry test can be implemented as a Wald test on the cumulated responses up to a specific horizon H:

$$I_y^h(\delta) - I_y^h(-\delta) = 0 \qquad h = 0, 1, 2, ..., H \tag{5}$$

To carry out this test an estimate of the variance of $I_y^h(\delta) - I_y^h(-\delta)$ is needed. This is obtained via bootstrap simulation, that is, a number N of artificial samples are generated using model (3)-(4) and for each artificial sample the IRF to a positive and negative shock is estimated as described above. The variance of $I_y^h(\delta) - I_y^h(-\delta)$ is then computed as the sample analogue from the N IRFs obtained on the artificial samples.

The algorithm proposed by KV is computationally intensive. Imagine that you have a time series model of T observations and that you want to carry out the IRF based test proposed by KV. This requires performing T (the number of histories) times 2S (the number of paths of future shocks) times N (the number of artificial samples used to estimate the covariance matrix of the IRF) simulations. If T=120 (say 30 years of quarterly data, a relatively short sample), S=250, and N=1000 (a relatively small number of artificial samples for bootstrap simulations) the number of simulations required to perform a test on a single model is around sixty million. If one wants to consider, as we do, 16 models the figure rises to seven hundred 20 million simulations, quite a high number even for modern computers.

To make the problem manageable while staying as close as possible to the true IRF we propose a slight modification of the KV algorithm. The idea is very simple: instead of performing a simulation for all the possible histories ω_{t-1}, we select 7 percentiles of the empirical distribution of y_t and x_t (the $12^{th}, 25^{th}, 37^{th}, 50^{th}, 63^{th}, 75^{th}, 88^{th}$) and evaluate the IRF at each of these data points. Our approximate IRF is then a weighted average of the seven IRFs obtained conditionally on each of these percentiles, where we weigh the different percentiles with a triangular window.[9] The computational benefits of such an approximation are sizable. In our previous example the number of simulations to compute a single IRF drops from sixty to three and a half millions.

To gauge the precision loss of such an approximation we conduct a small simulation exercise and compare the correct IRF, with the approximate one.[10] The data generating

[9]The weights are [0.04,0.07,0.11,0.14,0.11,0.07,0.04]. They therefore give maximum weight to the median, linearly decaying weights to both sides of the distribution and they sum to 1.

[10]It would be desirable to compare the size and power of the IRF asymmetry tests based on the correct and on the approximate IRFs. Such an exercise is, however, unfeasible for computational reasons.

process is a two equations system similar to the one used in the empirical application:

$$x_t = \alpha_1 y_{t-1} + \alpha_2 x_{t-1} + \epsilon_t \quad (6)$$

$$y_t = \gamma y_{t-1} \beta_1 x_t + \beta_2 x_{t-1} + \beta_1^+ x_t^+ + \beta_2^+ x_{t-1}^+ + u_t \quad (7)$$

The structural shocks ϵ_t and u_t are zero mean unit variance. We consider two different parameterizations (see Table 2) characterized by, respectively, a mild and a strong degree of asymmetries. The model is simulated 200 times for a sample size of $T = 100$ and the average response of the variable y_t to positive and negative shocks ϵ_t of different sizes (1 and 2 times the standard deviation) is computed using the true IRF and the approximate one. The number of future simulated shocks S is set to 250. The results are shown in Figure 4. The top two graphs refer to the model with mild asymmetries, the bottom two to the model with strong asymmetries. The cumulated responses of y_t to positive and negative shocks are reported for 16 steps ahead. It can be seen that for both configurations of parameters and across all horizons the approximate IRF matches very closely the true one.

5.2 Impulse Response Function based test results

We start with the results obtained using Mork's measure of asymmetry $x_t^{\#} = x_t^1$. We test for asymmetries in two different scenarios, in the former the system is hit by a commodity price shock of standard size (Table 3), in the latter we look at the response to a shock of larger magnitude, i.e. twice the historical standard deviation (Table 4). The tables are organized in five different panels corresponding to the different products, from aggregate food prices (top panel) to coffee prices (bottom panel) and show the p-values of the Wald asymmetry test discussed above. Values below 0.05 indicate that the null hypothesis of symmetry can be rejected at the 10% confidence level and are highlighted in bold. We trace the IRF for up to one year after the shock since, as expected, only a few months after the shock the p-values of the test converge to 1, indicating that at longer horizons, as the effect of the initial impulse vanishes, the responses to shocks of different signs converge to each other. The results of the IRF based tests are quite sharp. Considering a shock of standard size (Table 3) in only two cases the tests reject the null hypothesis of symmetry. The former is the response on impact and at one month lag case of processed food prices in Germany. The latter is the contemporaneous response of cereal prices in

Italy. In all other cases the p-values are above the confidence level, although in some cases only marginally (as in the case of the contemporaneous response of the prices of cereal products in France and of dairy products in Germany). In the case of a shock of larger size (Table 4) the results are even more clearly in favor of the null hypothesis of symmetry since the tests do not detect any significant difference in the IRF to shocks of different signs.

The results obtained with the models in which asymmetries are measured on the basis of the net increase ($x_t^{\#} = x_t^{12}$), are presented, respectively, in Tables 5 and 6. In the case of shocks of standard size (Table 5) the tests do not indicate any rejections of the null hypothesis, regardless of the product, country and horizon considered. Shocks of large magnitude (Tables 6) generate some asymmetry on impact only in the case of processed food prices in Italy.

Summarizing, tests based on a structural model of the food commodity/consumer price pass-through do not indicate the existence of systematic nonlinearities in the response of downstream prices to upstream shocks. Some rejections of the null hypothesis of linearity do indeed crop up sporadically in the analysis. The relevance of these rejections, in our view, should not however be overstated since (i) they are very short lived, manifesting themselves mostly on impact and (ii) when they appear in aggregate processed food prices they are hard to interpret as they are not backed up by any rejection in the underlying food categories.

6 Forecasting models

As discussed in the introduction, an alternative way to test for asymmetries without imposing identifying restriction on the structure of the model is to compare the relative forecast accuracy of linear models with respect that of models that include asymmetric terms. The rationale behind forecast based tests is then quite straightforward: if models that allow for asymmetries describe the data better than linear models do, the former can be expected to provide more accurate forecasts than the latter. In this section we therefore explore this route and conduct an analysis of the predictive performance for HICP food prices of models with and without asymmetries. Since our interest is in testing a dynamic relationship between future consumer prices and current commodity prices we derive forecasts directly from a predictive regression, rather than positing a multivariate model and iterating forecasts of both consumer and commodity prices. The approach, commonly used in empirical finance, consists of running a regression in which the dependent variable is the multistep ahead value being forecast. In particular, for each forecast horizon h we produce forecasts of consumer food prices on the basis of an h steps ahead regression of the following form:

$$\Delta y_t = a_y + \sum_{i=1}^{p} b_{21,i}\Delta x_{t-i-h+1} + \sum_{i=1}^{q} b_{22,i}\Delta y_{t-i-h+1} + \sum_{i=1}^{p} b^{\#}_{21,i} x^{\#}_{t-i-h+1} + \epsilon_{yt} \quad (8)$$

The competing linear model is nested in the one described by equation (8) and is obtained by setting to zero the coefficients associated with the nonlinear terms, that is $b^{\#}_{21,1} = \cdots = b^{\#}_{21,p} = 0$. In this respect the forecast exercise can be seen as an out-of-sample Granger causality test, see Busetti and Marcucci (2012).

A problem posed by the use of the direct, rather than the iterated, forecasting method is the presence of compounded prediction errors, which generate serial correlation of order $h-1$ in the regression residuals. As the forecast horizon rises, sampling error variance increases and the costs paid for using the direct method can more than offset the benefits (see Marcellino et al., 2006). The performance of the direct method, however, is substantially improved by the use of a lag length selection criterion that includes a correction for serial correlation (see Pesaran et al., 2011). In Appendix A we discuss how we implement such correction.

Two popular testing frameworks for evaluating the relative forecasting performance of two competing models are (i) equal forecast accuracy tests, like the one developed by Diebold and Mariano (1995, DM-test) and (ii) forecast encompassing tests like the one proposed by Harvey, Leybourne and Newbold (1998, HLN-test). Formally, equal forecast accuracy tests scrutinize the null hypothesis that two competing models provide the same mean squared prediction errors. The null hypothesis of forecast encompassing tests, instead, is that, when two competing forecasts for the same target variable are available, there is no gain in combining the prediction coming from model 1 with those obtained from model 2. Formally, given a target variable y_t, two forecasts provided by two competing models $\hat{y}_{t,1}$ and $\hat{y}_{t,2}$ and a combined forecast $\hat{y}_{t,c}$, where $\hat{y}_{t,c} = (1-\lambda)\hat{y}_{t,1} + \lambda\hat{y}_{t,2}$, for some nonnegative weight λ, model 1 encompasses model 2 if $\lambda = 0$. In standard settings both the DM and the HLM tests have Gaussian critical values. In our case, where the competing models are nested, some technical issues on the distribution of both tests arise. We refer the interested reader to Appendix B for further details on how to solve such problems.

6.1 Forecast based asymmetry tests results

The out-of-sample forecasting exercise is designed as follows. The first estimation period goes from the January 1997 (1999 in the case of the euro area) to April 2005. Each equation is estimated using information up to T-h, then the h steps ahead forecast is computed, where T runs from May 2005 to April 2011 and h= 1,2,. . . ,12. We employ rolling regressions, i.e. the estimation window is kept fixed at 100 observations (76 in the case of the euro area) as the forecast exercise proceeds.[11] Root Mean Squared Forecast Errors (RMSFE) are therefore computed on 72 data points spanning the May 2005-April 2011 period. At each step of the forecast exercise the lag length of the models is optimized on the basis of the Modified Akaike Criterion of Pesaran et al. (2011).

The results of the forecast competition obtained using Mork's measure of asymmetry are shown in Tables 7 and 8, those obtained using the net increase are displayed in Tables 9 and 10. Each cell shows the RMSFE obtained with the linear model relative to the one obtained on the basis of the correspondent model that includes asymmetric terms. Values higher than 1 therefore indicate that the predictive accuracy of the linear model is relatively higher than that of the nonlinear one. The cells highlighted in bold are those for which the equal forecast accuracy DM test or the encompassing HLM test reject the null hypothesis at the 10% confidence level.[12] The following results are worth remarking.

- When Mork's measure is used, linear models generally prevail in the forecast competition: in around 70% of the cases, the RMSFE ratios shown in Tables 7 and 8 are in fact above 1. There is, however, some heterogeneity across product categories. For example, while in the case of coffee, linear models improve upon nonlinear ones 98% of the times, the percentage falls to 44 in the case of dairy. Point estimates are surrounded by considerable uncertainty. In the case of equal predictive accuracy tests, the DM test rejects the null hypothesis only in 14 cases out of 240, and all the rejections are in favor of linear models (Table 7). According to forecast encompassing tests, nonlinear models should receive a positive weight in an optimal forecast

[11] An expanding estimation window gives broadly similar results. Results are available upon request from the authors.

[12] In the case of the encompassing test the null hypothesis is that the linear model encompasses the nonlinear one, i.e. that in an optimal forecast combination of the the models the latter receives zero weight.

combination only in 10 of the 240 cases considered: the null hypothesis that the linear models encompass the nonlinear ones is rejected at some horizons in the case of cereals for the euro area and for Italy, in the case of oils in the euro area and Germany, and in the case of coffee for the euro area at long horizons.

- When asymmetries are measured on the basis of the net increase (Tables 9 and 10), the prevalence of linear models is less sharp (54% of the cases). Also heterogeneity across product categories is even more pronounced as the percentage of times linear models outperform nonlinear ones goes from 79% (cereals) to 6% (dairy). Uncertainty is again very high. Equal predictive accuracy tests reject only 7% of the times (16 cases out of 240), and only three times in favor of nonlinear models. In the case of dairy products, where nonlinear models seem to provide on average more accurate forecasts, the DM test delivers only one rejection. The percentage of times the HLN tests reject the null also remains quite low (6%). One specific case worth highlighting is that of oils in Germany where the improvements yielded by the inclusion of asymmetric terms in the forecasting models are both sizeable and statistically significant according to both equal predictive accuracy and encompassing tests, at least at very short horizons.

All in all, the picture emerging from the forecasting exercise is in line with that provided by structural analysis. With the exception of dairy and oils in Germany, introducing asymmetric terms does not provide a *generalized* improvement in the forecasting performance of predictive regressions.

7 Conclusions

This paper analyzes the relationship between food commodity prices and processed food consumer prices in the euro area and in its largest countries (Germany, France and Italy), and investigates whether the latter respond asymmetrically to shocks to the former. Results from a baseline linear structural models indicate that there exists a significant degree of heterogeneity across countries and products in the response to commodity price shocks. In particular, retail prices of food products are generally more reactive in Germany than in Italy and in France. Also oils and dairy consumer prices respond more strongly to upstream shocks than cereals and coffee. Regarding the question whether positive shocks generate stronger responses than negative ones, based on two distinct but complementary approaches our analysis does not point to any compelling evidence of asymmetries. First, tests based on a structural model of the food commodity/consumer price pass-through do not indicate the existence of systematic differences in the the response of downstream prices to positive or negative upstream shocks. Second, the relative forecasting ability of linear models is generally superior to that of models that allow for sign-dependent non-linearities, although in the forecast competition results are less sharp, with some evidence of asymmetries showing up in the case of dairy and oils in Germany.

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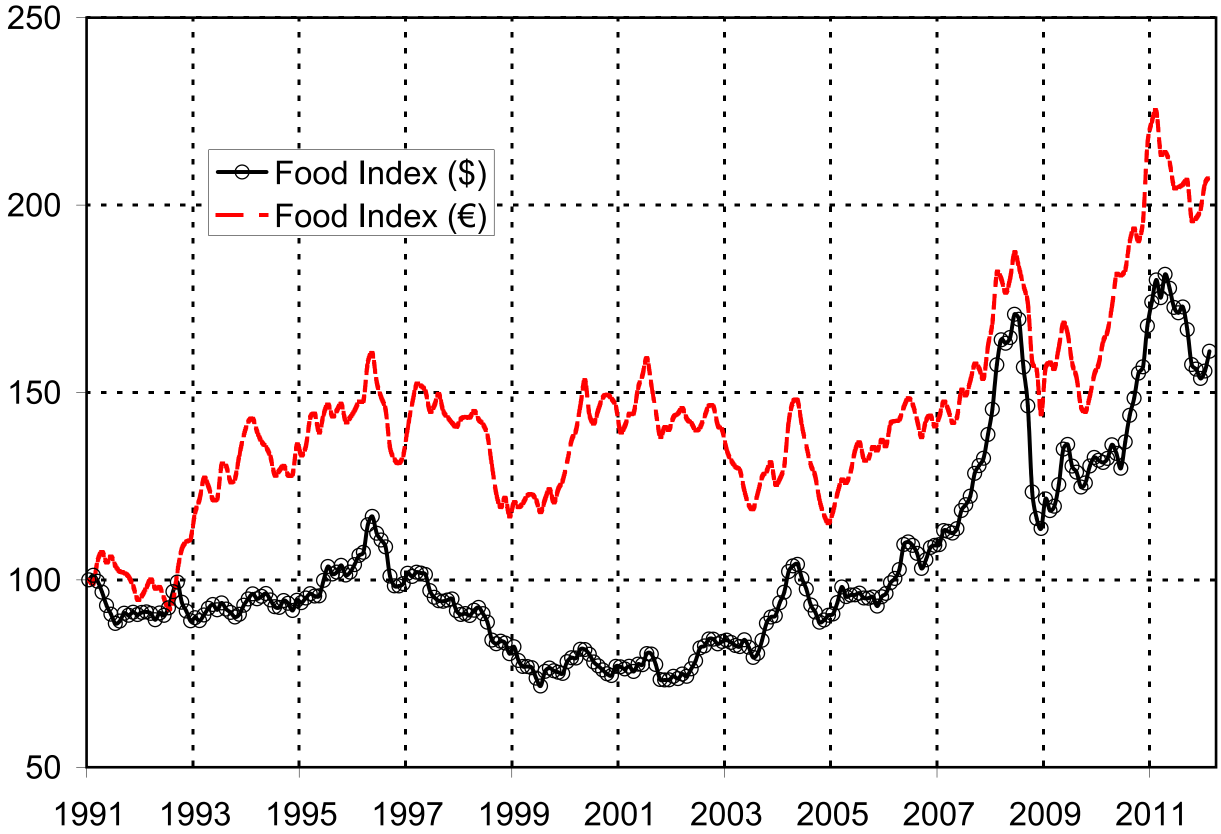

Figure 1: Food Commodity Index (Jan 1991=100), in US $ and in €
Source: IMF.

	Euro area	France	Germany	Italy
Processed food excluding alcohol and tobacco	8.0	8.2	6.8	8.9
Sub-components included in the analysis				
Bread and cereals	*2.6*	*2.3*	*1.9*	*3.4*
Milk, cheese and eggs	*2.2*	*2.4*	*1.7*	*2.4*
Oils and fats	*0.5*	*0.4*	*0.3*	*0.7*
Coffee, tea and cocoa	*0.4*	*0.5*	*0.4*	*0.3*
Total sub-components analyzed	5.6	5.5	4.4	6.7
Sub-components excluded from the analysis				
Sugar, jam, honey, syrups, etc.	*1.0*	*1.1*	*0.9*	*1.2*
Food products n.e.c.	*0.5*	*0.6*	*0.5*	*0.1*
Mineral waters, soft drinks, etc	*0.9*	*1.0*	*1.1*	*0.9*

Table 1: HICP weights - processed food (tot=100)

Note to Table 1: weights refer to 2011.

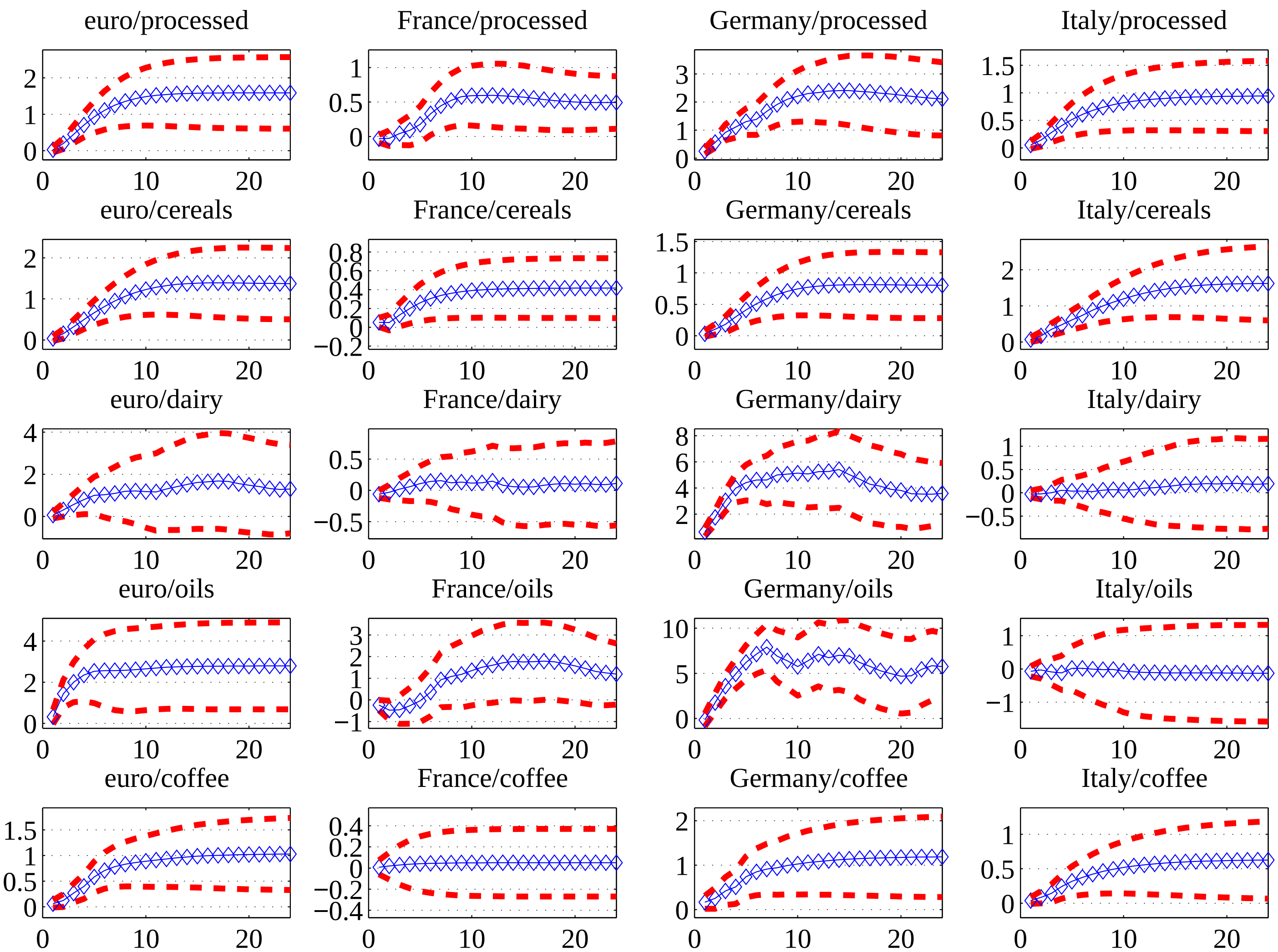

Figure 2: IRF to a 10 % commodity price shock in a linear structural VAR

Note to Figure 2: the Figure shows the cumulated impulse responses of HICP prices to a 10% shock in commodity prices. The impulse responses are calculated from a VAR model that includes food commodity prices HICP food prices. The estimation period is January 1999-April 2011. The lag length is selected with the Akaike Information Criterion. Confidence bands are obtained via Monte Carlo simulation.

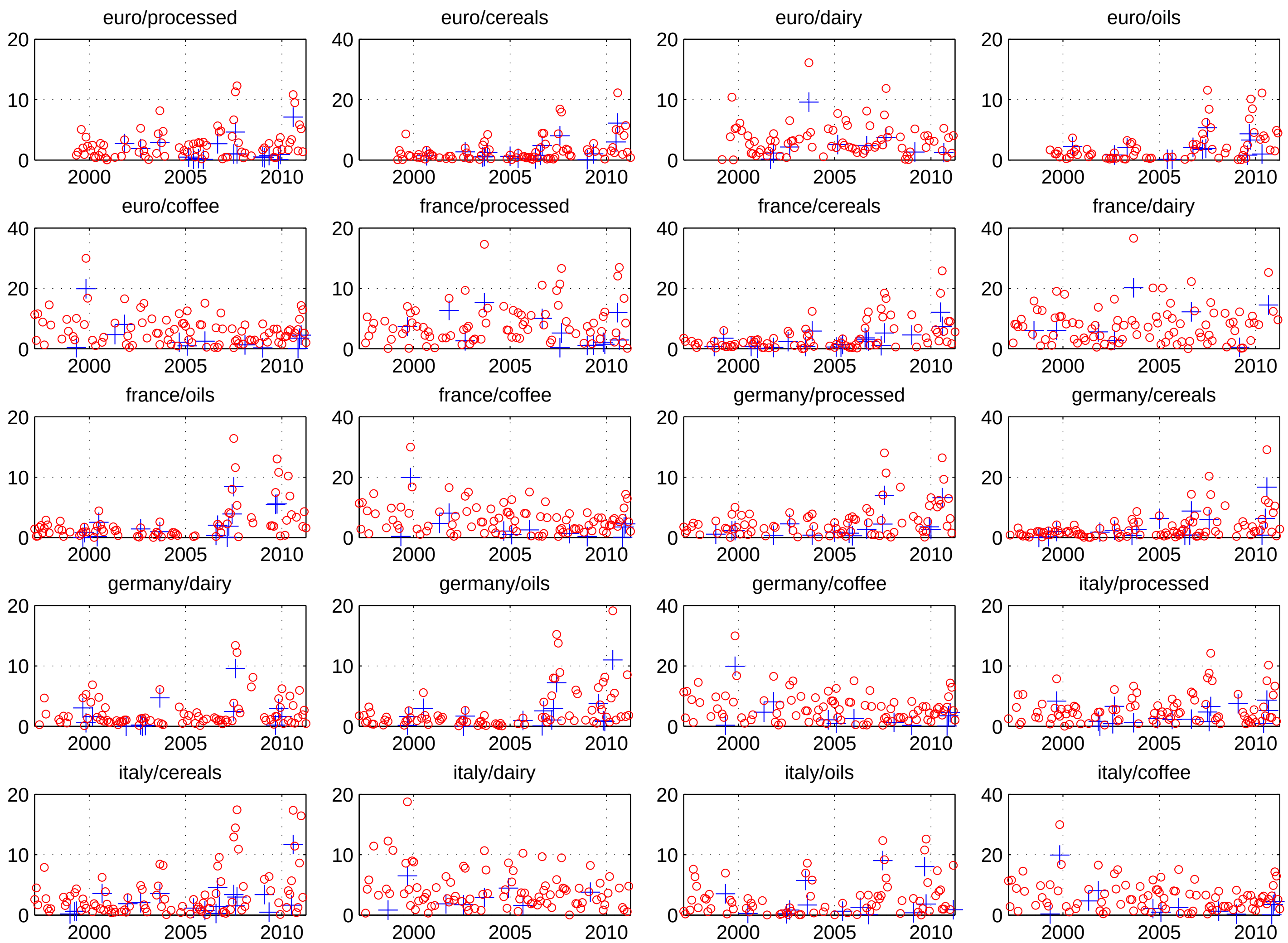

Figure 3: Food commodity prices: net (blue crosses) and simple increase (red circles)

Note to Figure 3: The red circles represent the data point in which Mork's measure of asymmetry $x_t^{\#} = x_t^1$ is different from zero, the blue crosses those for which the net increase $x_t^{\#} = x_t^{12}$ is positive.

	Mild asymmetry	Strong asymmetry
α_1	0.2	0.2
α_2	0.4	0.4
β_1	0.4	0.4
β_2	0.4	0.4
β_1^+	0.2	0.5
β_2^+	0.2	0.5
γ	0.2	0.2

Table 2: Different parameterizations for the simulation of the model described by equations (6)-(7).

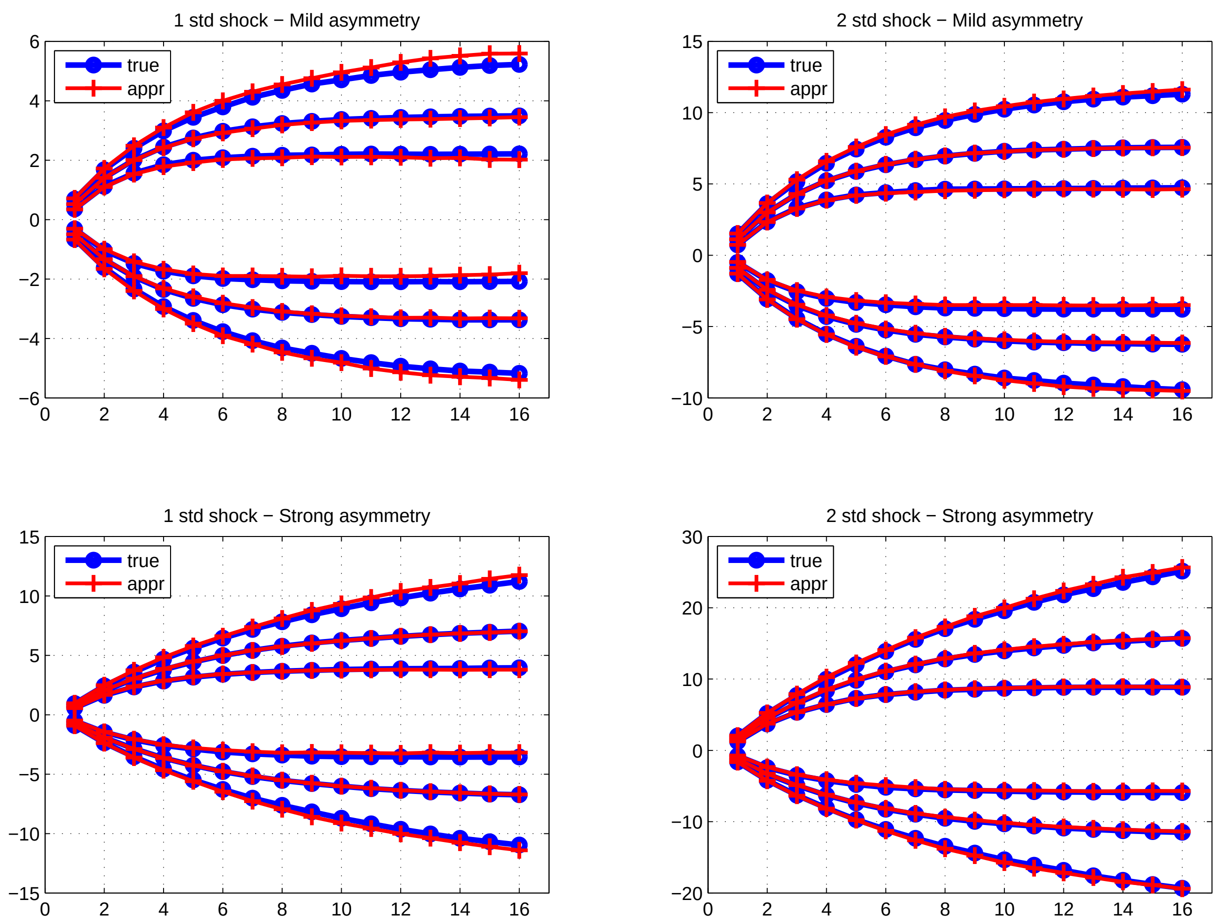

Figure 4: True and approximated impulse response function

	Horizon												
	0	1	2	3	4	5	6	7	8	9	10	11	12
							Processed food						
Euro area	0.27	0.57	0.47	0.72	0.88	0.95	0.98	0.99	1.00	1.00	1.00	1.00	1.00
Germany	**0.00**	**0.07**	0.18	0.37	0.60	0.76	0.89	0.95	0.98	0.99	1.00	1.00	1.00
France	0.49	0.78	0.95	0.96	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.54	0.75	0.87	0.96	0.93	0.94	0.98	0.99	1.00	1.00	1.00	1.00	1.00
							Cereals						
Euro area	0.64	0.63	0.21	0.45	0.71	0.87	0.95	0.98	0.99	1.00	1.00	1.00	1.00
Germany	0.25	0.72	0.75	0.92	0.95	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00
France	0.11	0.29	0.60	0.84	0.92	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00
Italy	**0.07**	0.42	0.76	0.93	0.93	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00
							Dairy						
Euro area	0.80	0.86	0.87	0.96	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.11	0.23	0.20	0.48	0.73	0.86	0.93	0.97	0.99	1.00	1.00	1.00	1.00
France	0.44	0.36	0.71	0.89	0.92	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.27	0.60	0.74	0.88	0.96	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00
							Oils						
Euro area	0.39	0.70	0.83	0.95	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.79	0.66	0.90	0.98	0.98	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00
France	0.18	0.37	0.32	0.61	0.80	0.92	0.97	0.99	1.00	1.00	1.00	1.00	1.00
Italy	0.79	0.78	0.95	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
							Coffee						
Euro area	0.56	0.84	0.94	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.46	0.83	0.96	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
France	0.96	0.86	0.91	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.26	0.13	0.24	0.51	0.64	0.83	0.92	0.97	0.99	0.99	1.00	1.00	1.00

Table 3: IRF based test of symmetry for 1 s.d. shock, $x_t^{\#} = x_t^1$,

Note to Table 3. Each cell shows the p-value of the χ^2_{H+1} test of equal IRF to a positive and a negative shock.

	Horizon												
	0	1	2	3	4	5	6	7	8	9	10	11	12
	Processed food												
Euro area	0.70	0.96	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.87	0.91	0.89	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
France	0.87	0.98	1.00	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.84	0.94	0.91	0.93	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00
	Cereals												
Euro area	0.75	0.96	0.98	1.00	1.00	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.65	0.95	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
France	0.47	0.87	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.55	0.91	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
	Dairy												
Euro area	0.59	0.92	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.89	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
France	0.98	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.92	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
	Oils												
Euro area	0.98	0.97	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
France	0.65	0.91	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.57	0.92	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
	Coffee												
Euro area	0.72	0.76	0.94	0.99	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.56	0.91	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
France	0.87	0.99	1.00	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.98	0.98	0.94	0.59	0.81	0.86	0.89	0.96	0.98	0.99	1.00	1.00	1.00

Table 4: IRF based test of symmetry for 2 s.d. shock, $x_t^{\#} = x_t^1$,

Note to Table 4. Each cell shows the p-value of the χ^2_{H+1} test of equal IRF to a positive and a negative shock.

	Horizon												
	0	1	2	3	4	5	6	7	8	9	10	11	12
	Processed food												
Euro area	0.13	0.51	0.72	0.89	0.95	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.63	0.84	0.95	0.99	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00
France	0.96	0.94	0.98	0.98	0.94	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.70	0.96	0.97	0.91	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00
	Cereals												
Euro area	0.73	0.86	0.95	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.37	0.81	0.91	0.83	0.90	0.96	0.99	1.00	1.00	1.00	1.00	1.00	1.00
France	0.77	0.94	0.95	0.97	0.95	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.13	0.55	0.84	0.96	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
	Dairy												
Euro area	0.21	0.28	0.63	0.85	0.94	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.37	0.65	0.87	0.97	0.83	0.93	0.98	0.98	0.99	1.00	1.00	1.00	1.00
France	0.36	0.71	0.92	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.89	0.98	0.98	0.84	0.94	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00
	Oils												
Euro area	0.53	0.86	0.96	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.70	0.51	0.79	0.80	0.90	0.97	0.99	1.00	1.00	1.00	1.00	1.00	1.00
France	0.76	0.83	0.59	0.79	0.72	0.88	0.90	0.95	0.98	0.99	1.00	1.00	1.00
Italy	0.57	0.53	0.84	0.95	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
	Coffee												
Euro area	0.46	0.25	0.60	0.81	0.93	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.35	0.28	0.56	0.81	0.93	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00
France	0.56	0.86	0.97	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.37	0.78	0.95	0.99	1.00	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00

Table 5: IRF based test of symmetry for 1 s.d. shock, $x_t^{\#} = x_t^{12}$,

Note to Table 5. Each cell shows the p-value of the χ^2_{H+1} test of equal IRF to a positive and a negative shock.

	Horizon												
	0	1	2	3	4	5	6	7	8	9	10	11	12
							Processed food						
Euro area	0.59	0.90	0.71	0.91	0.97	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.83	0.86	0.95	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
France	0.78	0.97	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	**0.03**	0.23	0.57	0.77	0.91	0.97	0.99	1.00	1.00	1.00	1.00	1.00	1.00
							Cereals						
Euro area	0.91	0.78	0.93	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.47	0.57	0.84	0.91	0.93	0.97	0.98	1.00	1.00	1.00	1.00	1.00	1.00
France	0.94	0.96	1.00	0.88	0.95	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.42	0.78	0.95	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
							Dairy						
Euro area	0.91	0.96	0.90	0.94	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.66	0.88	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
France	0.65	0.87	0.97	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.98	0.99	0.95	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
							Oils						
Euro area	0.39	0.82	0.67	0.88	0.94	0.98	0.98	0.99	1.00	1.00	1.00	1.00	1.00
Germany	0.92	0.99	1.00	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
France	0.30	0.43	0.77	0.82	0.94	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.42	0.71	0.93	0.87	0.93	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00
							Coffee						
Euro area	0.57	0.73	0.93	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.35	0.80	0.95	0.98	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
France	0.95	0.98	0.93	0.99	1.00	0.98	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.85	0.99	1.00	0.99	1.00	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00

Table 6: IRF based test of symmetry for 2 s.d. shock, $x_t^{\#} = x_t^{12}$,

Note to Table 6. Each cell shows the p-value of the χ^2_{H+1} test of equal IRF to a positive and a negative shock.

	Horizon											
	1	2	3	4	5	6	7	8	9	10	11	12
	Processed food											
Euro area	0.97	0.96	0.99	0.99	0.99	1.01	1.01	1.01	1.01	1.01	1.01	1.02
Germany	1.04	0.98	1.00	1.01	1.02	1.03	1.03	1.02	1.02	1.01	1.01	1.01
France	1.04	1.02	1.00	1.01	1.01	1.02	1.02	1.03	1.02	1.03	1.04	1.04
Italy	0.98	1.02	1.04	1.04	1.04	1.04	1.04	1.03	1.03	1.02	1.02	1.03
	Cereals											
Euro area	0.94	0.94	0.95	0.97	0.97	0.98	1.00	1.02	1.04	1.05	1.05	1.06
Germany	0.95	0.97	0.99	0.99	0.99	1.01	1.04	1.05	1.05	1.05	1.05	1.05
France	1.06	1.01	1.02	1.05	1.07	1.08	1.10	1.10	1.10	1.12	1.13	1.15
Italy	0.98	1.00	1.01	1.02	1.02	**1.02**	**1.02**	**1.02**	1.02	1.03	1.03	**1.03**
	Dairy											
Euro area	0.98	0.91	0.88	0.85	0.91	0.97	1.00	1.03	1.03	1.05	1.04	1.04
Germany	1.04	1.08	1.03	0.97	0.94	0.92	0.92	0.90	0.88	0.86	0.85	0.86
France	1.05	1.02	0.99	0.99	0.98	0.98	0.96	0.96	0.97	0.98	0.98	0.97
Italy	0.99	1.01	1.00	1.00	1.01	1.02	1.02	1.02	1.03	**1.03**	**1.03**	**1.04**
	Oils											
Euro area	0.91	1.03	1.12	1.13	1.12	1.14	1.19	1.20	1.22	1.22	1.17	1.11
Germany	0.99	0.98	1.15	1.26	1.27	1.25	1.27	1.26	1.28	1.25	1.19	1.12
France	0.86	0.86	0.86	0.89	0.91	0.95	0.98	1.01	1.03	1.05	1.05	1.05
Italy	0.97	0.98	0.99	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	0.99
	Coffee											
Euro area	1.03	1.03	1.03	1.03	1.04	1.04	1.06	1.07	1.08	**1.09**	**1.10**	**1.11**
Germany	**1.07**	1.06	1.07	1.04	1.02	1.01	**1.01**	1.00	1.01	1.01	1.02	**1.03**
France	1.01	1.02	1.04	1.04	1.06	1.05	1.05	1.05	1.07	1.08	1.09	**1.11**
Italy	0.98	1.02	1.02	1.01	1.01	1.00	1.01	1.02	1.01	1.01	1.01	1.01

Table 7: Equal predictive accuracy tests, $x_t^{\#} = x_t^1$ - rolling regressions

Note to Table 7. Each cell shows the ratio of the Root Mean Squared Forecast error (RMSFE) obtained with the model with nonlinearities to that obtained with the linear model. Values higher (lower) than 1 therefore indicate that the linear model outperforms (is outperformed by) the nonlinear one. Entries in bold are those for which the hypothesis of equal forecast accuracy can be rejected at the 5% confidence level according to the Diebold-Mariano test.

	Horizon											
	1	2	3	4	5	6	7	8	9	10	11	12
	Processed food											
Euro area	0.97	0.96	0.99	0.99	0.99	1.01	1.01	1.01	1.01	1.01	1.01	1.02
Germany	1.04	0.98	1.00	1.01	1.02	1.03	1.03	1.02	1.02	1.01	1.01	1.01
France	1.04	1.02	1.00	1.01	1.01	1.02	1.02	1.03	1.02	1.03	1.04	1.04
Italy	0.98	1.02	1.04	1.04	1.04	1.04	1.04	1.03	1.03	1.02	1.02	1.03
	Cereals											
Euro area	**0.94**	0.94	0.95	0.97	0.97	0.98	1.00	1.02	1.04	1.05	1.05	1.06
Germany	0.95	0.97	0.99	0.99	0.99	1.01	1.04	1.05	1.05	1.05	1.05	1.05
France	1.06	1.01	1.02	1.05	1.07	1.08	1.10	1.10	1.10	1.12	1.13	1.15
Italy	**0.98**	1.00	1.01	1.02	1.02	1.02	1.02	1.02	1.02	1.03	**1.03**	**1.03**
	Dairy											
Euro area	0.98	0.91	0.88	0.85	0.91	0.97	1.00	1.03	1.03	1.05	1.04	1.04
Germany	1.04	1.08	1.03	0.97	0.94	0.92	0.92	0.90	0.88	0.86	0.85	0.86
France	1.05	1.02	0.99	0.99	0.98	0.98	0.96	0.96	0.97	0.98	0.98	0.97
Italy	0.99	1.01	1.00	1.00	1.01	1.02	1.02	1.02	1.03	1.03	1.03	1.04
	Oils											
Euro area	**0.91**	**1.03**	**1.12**	1.13	1.12	1.14	1.19	1.20	1.22	1.22	1.17	1.11
Germany	**0.99**	0.98	1.15	1.26	1.27	1.25	1.27	1.26	1.28	1.25	1.19	1.12
France	0.86	0.86	0.86	0.89	0.91	0.95	0.98	1.01	1.03	1.05	1.05	1.05
Italy	0.97	0.98	0.99	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	0.99
	Coffee											
Euro area	1.03	1.03	1.03	1.03	1.04	1.04	1.06	1.07	1.08	1.09	**1.10**	**1.11**
Germany	1.07	1.06	1.07	1.04	1.02	1.01	1.01	1.00	1.01	1.01	1.02	1.03
France	1.01	1.02	1.04	1.04	1.06	1.05	1.05	1.05	1.07	1.08	1.09	1.11
Italy	0.98	1.02	1.02	1.01	1.01	1.00	1.01	1.02	1.01	1.01	1.01	1.01

Table 8: Forecast encompassing tests, $x_t^{\#} = x_t^1$ - rolling regressions

Note to Table 8. Each cell shows the ratio of the Root Mean Squared Forecast error (RMSFE) obtained with the model with nonlinearities to that obtained with the linear model. Values higher (lower) than 1 therefore indicate that the linear model outperforms (is outperformed by) the nonlinear one. Entries in bold are those for which the hypothesis of forecast encompassing can be rejected at the 5% confidence level according to the Harvey-Leybourne-Newbold test.

	Horizon											
	1	2	3	4	5	6	7	8	9	10	11	12
	Processed food											
Euro area	1.02	0.97	1.01	1.01	1.01	1.01	1.01	1.01	1.01	**1.01**	1.01	**1.01**
Germany	1.07	0.96	0.95	0.97	0.97	0.97	0.98	0.98	0.99	1.00	0.99	**1.00**
France	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	1.01	1.05	1.04	1.04	1.03	1.03	1.03	1.03	1.03	1.03	1.03	1.02
	Cereals											
Euro area	1.14	1.04	1.02	1.01	1.00	1.00	1.00	1.01	**1.01**	**1.01**	**1.01**	1.01
Germany	1.06	1.02	1.01	**1.01**	1.01	1.01	1.01	1.01	1.01	1.01	1.01	1.02
France	1.01	0.99	1.00	1.00	0.99	0.99	0.99	0.98	0.99	0.98	0.99	0.99
Italy	1.00	1.00	1.00	1.01	1.00	1.00	1.00	1.00	1.00	1.00	1.00	**1.01**
	Dairy											
Euro area	0.98	0.99	0.99	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.97	0.92	0.90	0.91	0.92	0.94	0.97	0.98	0.97	0.97	0.97	0.98
France	1.00	**1.00**	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.99	1.00	0.99	1.00	1.00	0.99	0.99	0.99	0.99	0.99	0.99	0.99
	Oils											
Euro area	0.91	0.96	0.99	1.01	1.00	1.01	1.01	**1.01**	**1.01**	**1.01**	0.99	0.99
Germany	**0.75**	**0.71**	0.68	0.76	0.78	0.80	0.82	0.82	0.76	0.76	0.74	0.73
France	1.09	1.02	1.01	1.01	1.02	1.06	1.09	1.07	1.09	1.10	1.12	1.16
Italy	1.00	1.02	1.02	1.02	1.01	1.01	1.01	1.01	1.01	1.01	1.01	**1.01**
	Coffee											
Euro area	1.00	1.01	1.01	1.01	1.01	1.01	1.01	1.01	1.01	1.00	1.00	1.00
Germany	0.98	0.99	1.00	1.00	0.99	1.00	0.99	0.99	1.00	**0.99**	0.99	0.99
France	1.00	1.00	1.00	1.01	1.01	1.01	1.01	1.02	1.02	1.01	1.01	1.01
Italy	0.99	0.99	0.99	0.99	0.99	1.00	1.00	1.00	1.00	0.99	1.00	1.00

Table 9: Equal predictive accuracy tests, $x_t^{\#} = x_t^{12}$ - rolling regressions

Note to Table 9. Each cell shows the ratio of the Root Mean Squared Forecast error (RMSFE) obtained with the model with nonlinearities to that obtained with the linear model. Values higher (lower) than 1 therefore indicate that the linear model outperforms (is outperformed by) the nonlinear one. Entries in bold are those for which the hypothesis of equal forecast accuracy can be rejected at the 5% confidence level according to the Diebold-Mariano test.

	Horizon											
	1	2	3	4	5	6	7	8	9	10	11	12
	Processed food											
Euro area	1.02	0.97	1.01	1.01	1.01	1.01	1.01	1.01	1.01	1.01	1.01	1.01
Germany	1.07	0.96	0.95	0.97	0.97	0.97	0.98	0.98	0.99	1.00	0.99	1.00
France	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	1.01	1.05	1.04	1.04	1.03	1.03	1.03	1.03	1.03	1.03	1.03	1.02
	Cereals											
Euro area	1.14	1.04	1.02	1.01	1.00	1.00	1.00	1.01	1.01	1.01	1.01	1.01
Germany	1.06	1.02	1.01	**1.01**	1.01	1.01	1.01	1.01	1.01	1.01	1.01	1.02
France	1.01	0.99	1.00	1.00	0.99	0.99	0.99	0.98	0.99	0.98	0.99	0.99
Italy	1.00	1.00	1.00	1.01	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.01
	Dairy											
Euro area	0.98	0.99	0.99	0.99	0.99	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Germany	0.97	**0.92**	**0.90**	0.91	0.92	0.94	0.97	0.98	0.97	0.97	0.97	0.98
France	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Italy	0.99	1.00	0.99	1.00	1.00	0.99	0.99	0.99	0.99	0.99	**0.99**	**0.99**
	Oils											
Euro area	**0.91**	0.96	0.99	1.01	1.00	1.01	1.01	1.01	1.01	1.01	0.99	0.99
Germany	**0.75**	**0.71**	0.68	0.76	0.78	0.80	0.82	**0.82**	0.76	0.76	0.74	0.73
France	1.09	1.02	1.01	1.01	1.02	1.06	1.09	1.07	1.09	1.10	1.12	1.16
Italy	1.00	1.02	1.02	1.02	1.01	1.01	1.01	1.01	1.01	1.01	1.01	1.01
	Coffee											
Euro area	1.00	1.01	1.01	1.01	1.01	1.01	1.01	1.01	1.01	1.00	1.00	1.00
Germany	**0.98**	0.99	**1.00**	1.00	0.99	1.00	**0.99**	**0.99**	1.00	**0.99**	0.99	0.99
France	1.00	1.00	1.00	1.01	1.01	1.01	1.01	1.02	1.02	1.01	1.01	1.01
Italy	**0.99**	0.99	0.99	0.99	0.99	1.00	1.00	1.00	1.00	0.99	1.00	1.00

Table 10: Forecast encompassing tests, $x_t^{\#} = x_t^{12}$ - rolling regressions

Note to Table 10. Each cell shows the ratio of the Root Mean Squared Forecast error (RMSFE) obtained with the model with nonlinearities to that obtained with the linear model. Values higher (lower) than 1 therefore indicate that the linear model outperforms (is outperformed by) the nonlinear one. Entries in bold are those for which the hypothesis of forecast encompassing can be rejected at the 5% confidence level according to the Harvey-Leybourne-Newbold test.

A Model Selection in multistep predictive regressions

At each step in the forecast exercise we choose the number of lags p and q in the predictive regression (8) as the ones that minimize the following modified Akaike Criterion (AIC), see Pesaran et al. (2011):

$$AIC(h) = ln[\hat{u}_t'\hat{u}_t/(w-h)] + \frac{2tr(\Pi)}{w-h} \tag{9}$$

where $\hat{u}_t$ are the OLS residuals obtained from model 8, w is the number of observations used to estimate the model and h is the forecast horizon. The matrix Π is defined as follows:

$$\Pi = \widehat{\Sigma_{zz}}^{-1}\Omega\widehat{\Sigma_{zz}} \tag{10}$$

where $\widehat{\Sigma_{zz}}$ is the variance of the coefficients of the predictive regression and Ω is the Newey-West long-run covariance matrix of the residuals $\hat{u}_t$ with bandwith parameter equal to $min(h, w^{1/3})$.

B Forecast accuracy tests

In the case of nested models Clark and McCraken (2001, 2005) argue that both the DM and the HLN statistics have non standard distributions that depend on several nuisance parameters. In this case the critical values should be obtained through bootstrap simulations. The power and size of these tests has been more recently analyzed by Busetti and Marcucci (2012). On the basis of Monte Carlo simulations they confirm that for nested models and multistep forecasts the original DM test is undersized, but they also find that the original HLN test displays good size and power properties, especially as the prediction sample increases. In this case the Guassian distributed HLN statistics becomes relative more attractive than the computationally intensive procedure advocated by Clark and McCraken (2005). Also, Clark and McCraken (2012) find that the DM test has reasonable size provided that the long-run variance of the test statistics is estimated nonparametrically with a *rectangular window* rather than with a triangular one.

Considering the recommendations emerging from these literature we correct the DM test as in Clark and McCraken (2012) and use the HLN test in its standard formulation. We consider only one-sided encompassing tests, in the sense that the null hypothesis is that the nested (linear) model encompasses the nonlinear one, and the alternative is that the nesting (nonlinear) model yields better forecasts. If the test cannot reject this hypothesis the linear model is said to encompass the one with asymmetric terms.

The test statistics are defined as follows. Let T be the total number of observations, R be the number of in-sample observations used to estimate the model, P the number of out-of-sample predictions, h the forecast horizon and $e_{t,m}$ the forecast errors of model m, where $m = 1, 2$. The DM test statistics is the following:

$$DM = P^{\frac{1}{2}} \bar{f} / \hat{\sigma}_{DM} \tag{11}$$

where $\bar{f} = \frac{1}{P} \sum_{t=R+h}^{T} f_t$ and $f_t = e_{t,1}^2 - e_{t,2}^2$. The denominator, $\hat{\sigma}_{DM}$, is a nonparametric estimator of the long-run variance of f_t:

$$\hat{\sigma}_{DM} = P^{-1} \sum_{t=R+h}^{T} (f_t - \bar{f}) + 2P^{-1} \sum_{j=1}^{m} w(j,m) \sum_{t=j+R+h}^{T} (f_t - \bar{f})(f_{t-j} - \bar{f}) \tag{12}$$

Taking into account the results of Clark and McCraken (2012) we set $m = h - 1$ and $w(j,m) = 1$.

Turning to forecast encompassing, the forecast errors of the competing models and those of the combined forecast satisfy the relationship $e_{t,1} = \lambda(e_{t,1} - e_{t,2}) + e_{t,c}$. In population, under the null hypothesis that $e_{t,1}$ encompasses $e_{t,2}$ $\lambda = 0$, or equivalently, $g_t = e_{t,1}(e_{t,1} - e_{t,2}) = 0$. The HLN test statistics is:

$$HLN = P^{\frac{1}{2}} \frac{\bar{g}}{\hat{\sigma}_{HLN}} \tag{13}$$

where $\bar{g} = \frac{1}{P} \sum_{t=R+h}^{T} g_t$. The variance $\hat{\sigma}_{HLN}$ is a non-parametric estimator of the long-run variance of g_t, analogue to 12. Like in the model selection step we use a triangular window with $w(j,m) = 1 - j/(m+1)$ and set $m = 1.5h$, in line with Busetti and Marcucci (2012).

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