World Commodity Prices and Domestic Retail Food Inflation: Some Insights from the UK

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Abstract

We address the links between world commodity prices and retail food price inflation, focussing on two aspects. First, since world commodity prices represent a relatively small share of costs of retail food products, retail price behaviour may differ from world commodity prices and other factors (exchange rates and other input costs) will also matter in determining retail food inflation. Second, noting that the world price spike of 2007-2008 was different in the level and duration from the price spike experienced in 2011, we also emphasise an obvious but neglected fact that the effect on retail food price inflation depends on the duration of the shocks on world commodity markets, not just the magnitude of price spikes (the latter often commanding most attention). Being an open economy reliant on world commodity trade, the UK offers a natural and hitherto unexplored setting for the analysis. Applying time series methods to a sample of 259 monthly observations over the 1990(9)-2012(3) period we find substantial and significant long term partial elasticities for domestic food price inflation with respect to world food commodity prices, the exchange rate and oil prices (the latter indirectly via a relationship with world food commodity prices). Domestic demand pressures and food chain costs are found to be less substantial and significant over our data period. Interactions between the main driving variables in the system tend to moderate rather than exacerbate these partial effects. Furthermore, the persistence of shocks to these variables markedly affects their effects on domestic food prices.

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1. Introduction

The dynamics of retail food prices differ from the behaviour of retail prices in non-food sectors but also from the price changes observed on world commodity markets. There are two dimensions to this. First, since the early 2000s, food inflation in most OECD countries has on average been higher than non-food inflation: in the UK, for example, annual retail food inflation has averaged 3.2% compared to average non-food inflation of 1.6%. Second, retail food inflation has tended to be more volatile than non-food inflation: this is most obviously associated with the commodity price spikes on world markets in 2007-2008 and 2011 and the subsequent price collapse, giving rise to rates of food inflation in the UK between 13% and -3% compared to 3.7% and -0.1% for non-food inflation over the 2007-2014 period. The obvious point here is that retail food prices are tied to but distinct from the behaviour of prices on world agricultural markets. With some limited exceptions - and even in cases where there are no explicit actions to limit the impact of world price changes through trade and market support policies - these observations regarding the relative levels and volatility of retail food prices hold over a large number of countries.

Against this background, we provide new insights into the assessment of the links between events on world commodity markets and domestic retail food inflation that have gone largely ignored in the now-voluminous research on the effects of recent events. Specifically, the behaviour of domestic retail food prices is quite different from world agricultural prices, but insofar as they are linked, it is the *cumulative* effects of events on world markets that matters for domestic food inflation rather than a direct pass-through of shocks. These findings have potentially important implications for empirical research on food markets and how the interpretation of commodity market events is translated into policy advice. In particular, while recent developments on world markets have generated a capacious literature on the causes and consequences of world price spikes, it is important to note that in the developed world at least, the often-referred to 'food' products traded on world markets are not the same as 'food' products purchased by domestic consumers[2]. Indeed, though the cost shares vary across food products, unprocessed agricultural commodities typically account for a relatively small share of the final processed products bought by consumers (25-30% for the US, Hobijn (2008); 15-30% in the EU, Bukeviciute *et al.* (2009)). As a consequence, factors such as labour costs other inputs into processed foods are likely to be important in determining retail

[2] Notable papers covering the determinants of world commodity prices include Ferrucci *et al.* (2012), Wright (2011) and Gilbert (2010) among others.

prices. Moreover, given that most commodities traded on world markets are priced in US dollars, the effect on domestic retail food prices will also depend on exchange rates, which may offset or exacerbate the dollar-denominated world commodity price change. Of course, there has been much valuable research addressing the transmission of world prices through to domestic food prices in the wake of the commodity price crisis (*inter alia* Gilbert 2010; IMF 2011) but in large part empirical studies have focussed on bi-variate price relations in the US (e.g. Baumeister and Kilian, 2014) and the Eurozone (e.g. Porqueddu and Venditti, 2012) in contrast to the multivariate approach that is here applied to the UK; an economy inside the EU, reliant on trade, yet operating its own exchange rate[3].

The second issue we address in linking world price developments to retail food prices is to investigate the effects of both the magnitude *and* duration of price shocks in determining the inflationary effect. While it is well-known that commodity markets are characterised by long-periods of price stability interrupted by short-lived 'spikes' (most notably from Deaton and Laroque (1992) and Williams and Wright (1991)) the impact of the longevity of the spike on retail food inflation is a dimension of the shock that has gone largely unnoticed in empirical work. However, we might expect the cumulative effect of the commodity price shock to drive domestic food inflation rather than any transitory volatility. While the one-period shocks (that are the typical currency of simulation exercises) delivers useful summaries of the estimated impacts, they contrast with the empirical reality in which commodity price shocks are idiosyncratic in both magnitude and duration, reflecting differences in their underlying causes, commodity composition and the macro-economic conditions prevailing at the time.

Figure 1 displays the FAO food commodity index since the mid-1990s. It is clear that the experience of the first global commodity price hike of 2007-2008 differed markedly from the second in 2011, the former being an archetypal spike in the series, the second being characterised by a rapid inflation but more gradual decline. Less obvious, but arguably more important, is the 'momentum' fuelling these episodes. In contrast to the spike of 2008 which emerged gradually and then abruptly from the relatively low levels of 2002, the 2011 hike in prices developed from prices what were already at historically high levels.

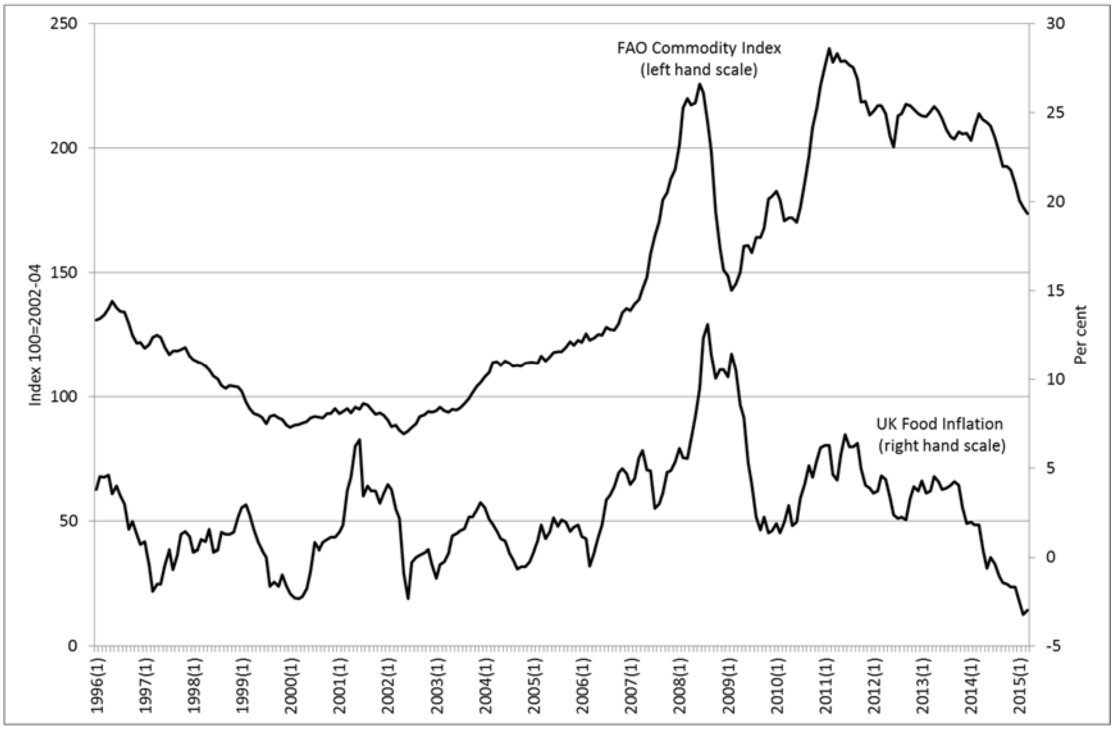
While the likely causes of these differences have been documented elsewhere (see *inter alia*, Abbott et al. (2011), Ferrucci (2012), IMF (2008) and Tadesse *et al*. (2014)) it is clear that the anatomy of these episodes was different, meaning that for any given lag structure, the

[3] The UK's reliance on food imports is evidenced by data from the UK's Department of Environment, Food and Rural Affairs (defra) which report UK self-sufficiency in 2013 for 'All Food' at around 60 per cent.

characteristics of commodity price spikes will have a different effect on domestic food price inflation. This is important in the context of the experience of UK, whose food inflation is also displayed on Figure 1. Note here that although UK food inflation reproduced the spike on commodity markets in 2008, it rose only modestly during 2011 despite the escalation in commodity prices being little different to that experienced in 2008.

While other factors (e.g. including the underlying causes, how quickly supply responds, the adjustments made by food firms and retailers as well as the macroeconomic environment) may clearly have played a role in differentiating the two responses, the observation that the 'build up' to the two world price spikes differed suggests that the *cumulative* impact of world price developments may also have contributed to the different experience of UK food inflation in these two periods.

Figure 1: FAO Food Commodity Price Index and UK Food Inflation (1996-2015)



These two dimensions, namely the multiplicity of potential drivers and the magnitude and duration of commodity price shocks, are the focus of this paper. We address these issues in the context of the recent experience of food price inflation in the UK. Using monthly data covering the 1990-2012 period, we employ a co-integrated vector autoregressive framework that allows us to account for a wider range of factors that determine domestic food inflation. In addition, by using impulse response functions and a variance decomposition approach, we distinguish the relative contribution of world commodity prices and other macroeconomic factors in food inflation. We also highlight the importance of the characteristics of commodity price spikes in driving the retail price effect using an econometric approach that readily facilitates the introduction of shocks of various (size and) duration. To summarise, the main insights are that: (i) world agricultural prices, the Sterling-US Dollar exchange rate and oil prices have been the main drivers of food inflation in the UK in recent years, though their relative contribution can vary depending on the lags in the pass-through of these variables and; (ii) the duration of world commodity price spikes affects the role of commodity prices in UK food inflation. Does a short-lived commodity price spike have a different effect on food inflation from a price shock of the same magnitude but lasting for a longer period? Our hypothesis is that persistent commodity price shocks increase retail prices to a greater extent than a one-period shock of the same magnitude. Estimates from the model suggest the effect of a permanent shock is more than seven times that of a temporary shock.

The paper is organised as follows. In Section 2, we briefly summarise empirical research on pass-through between world commodity prices and domestic inflationary impacts. The econometric framework is outlined in Section 3 together with a summary of the main estimation results. In Section 4, we explore the nature of retail food price dynamics in an attempt to separate the impact of duration and magnitude in commodity price shocks in the short term. We also report the results from variance decomposition which highlights the role of factors determining UK food inflation over different time horizons. In Section 5, we summarise and conclude.

2. Commodity Pass-Through and Food Inflation

Standard approaches to evaluating the impact of world commodity events on domestic markets typically focus on the pass-through between world and domestic prices. This approach has also been widely employed in estimating the effect of world oil prices on inflation (for example, Blanchard and Gali, 2007). Hamilton (2008) provides an extensive review of the oil-inflation pass-through literature. Estimating pass-through has also featured in the analysis of agricultural-food markets with Vavra and Goodwin (2005) providing a summary of these issues. Focussing more directly on recent world commodity market effects on food inflation, IMF (2008) investigates the pass-through effects on domestic general inflation for a wide range of countries[4]. Ferruci *et al.* (2012) address food price pass-through

[4] There is, of course, a more extensive literature on pass-through from upstream markets through to retail but where the upstream market is domestic. While these issues overlap to some extent with the concerns addressed here, our focus is more on linkages between domestic prices and commodity price spikes, these having been at the forefront of public policy concerns raised in recent years.

in the Euro area. Their analysis shows that world prices are a poor approximation for cost pressures that determine domestic retail food prices because government intervention in the form of the Common Agricultural Policy breaks the direct link between world prices and domestic agricultural prices, the latter being the main driver of food prices in Euro area countries. Other papers which have addressed pass-through of prices in upstream markets through to domestic food prices in the context of the EU include Bukeviciute *et al.* (2009), Porqueddu and Venditti (2012) and Bakucs and Fertö (2013). [5]

One potential downside of the approach typically applied is that it estimates bi-variate time series models and hence does not account for other factors that may determine domestic food inflation or influence the pass-through effect. The theoretical literature on pass-through in agricultural markets indicates that other factors are important. For example, Gardner (1975) and McCorriston *et al.* (1999) note that raw commodity inputs are only one source of costs determining retail food prices which, by extension, implies that other factors that could drive retail prices should also be accounted for[6]. Recognising this disconnect between time series work on pass-through in commodity/food markets, the econometric framework we employ accommodates a range of other factors which may also have an impact on the pass-through between world commodity and domestic (UK) retail food prices. These include macroeconomic factors such as the exchange rate, which determines the own-currency price effect of US dollar-denominated price changes, domestic agricultural prices and the price of oil as well as other demand and supply shifters that underpin the theory of price transmission in the food sector (such as the costs of labour). Thus, while the econometric approach -like many other studies in this area- is atheoretical, we relate the specification to an underlying structural story in that we recognise that (a) a range of factors can determine changes in retail prices not just raw agricultural commodity prices determined on world markets and (b) as a consequence, that raw commodity prices on world markets will behave differently from

[5] Given the focus of this paper, we acknowledge but do not address the issue of asymmetric price transmission. There is a considerable literature on this issue more generally (see Meyer and von Cramon-Taubadel (2004), Vavra and Goodwin (2005), Frey and Manera (2007) and Bakucs, Falkowski and Fertö (2014) for comprehensive surveys). Empirical evidence on the topic is mixed and to some extent model-dependent but asymmetric price transmission appears to be most evident in markets for specific products; aggregation over products tending to mask the asymmetric responses detected at more specific levels (e.g. only dairy products in the UK, London Economics 2003 and AHDB 2011). Recent empirical analyses of Porqueddu and Venditti (2012) and Hassouneh *et al.* (2015) find little evidence of asymmetry at either aggregate or commodity-specific levels using two recently-proposed test methods. Given the numerous approaches to testing (Frey and Manera (2007) catalogue six further approaches) and our interest in the macroeconomic factors affecting food price transmission, the model adopted here assumes symmetric adjustment. It should be noted however, that the tractability of asymmetric price transmission testing is a key attraction in the bi-variate setting.

[6] The difference between these two approaches is that Gardner (1975) assumes the food industry to be competitive while McCorriston *et al.* (1998) allow for market power in the food sector in determining pass-through.

domestic retail food prices. Related to the latter, we also draw on the observation that agricultural prices are characterised by long periods of relatively low and stable prices punctuated with short-lived spikes (see Deaton and Laroque (1992) and Wright and Williams (1991)). As such, the behaviour of food inflation might well depend on the characteristics of spike episodes including the build-up of the spike, where it starts from and how long it lasts. Since these factors might also feed through to retail food prices in an accumulated manner, it is not just the level raw commodity prices reach but also the duration of the spike episode that could determine the inflationary impact of commodity price movements.

3. Empirical Model

(i) Econometric approach

In an open economy such as the UK, we presume that aggregate food prices reflect a basic relationship posited by the theoretical literature on pass-through augmented by macroeconomic factors that are small in number, non-stationary and dynamically complex. In such circumstances, the co-integrated vector autogressive (C-VAR) rather than a bi-variate model offers a tractable framework. Applying this approach to the experience of UK retail food inflation, the specification of the C-VAR is given by:

$$\mathbf{x}_t = \mathbf{\Phi}_1 \mathbf{x}_{t-1} + \mathbf{\Phi}_2 \mathbf{x}_{t-2} + \ldots + \mathbf{\Phi}_p \mathbf{x}_{t-p} + \mathbf{\Psi D}_t + \boldsymbol{\varepsilon}_t \qquad (1)$$

whereby $\mathbf{x}_t$ is a vector of I(1) variables containing the UK retail food price index (r_t) and a set of dollar-denominated factors that are likely to play a role in the price transmission process, namely world agricultural prices (w_t), the world price of oil (o_t) and the Sterling-Dollar exchange rate (e_t). In addition, as an agricultural producer, domestic farm-gate prices (d_t), which may differ from world prices in both composition and timing, could also be expected to play a role. To capture the non-agricultural costs of food processing and retailing, $\mathbf{x}_t$ is augmented by UK labour costs (c_t) and the level of UK unemployment (u_t) the latter being used to proxy domestic demand for food.[7,8] Deterministic terms (constants, trends, seasonals and dummies) populate $\mathbf{D}_t$ and $\boldsymbol{\varepsilon}_t$ is a vector of disturbances, each element of which is assumed to be serially independent with zero mean and finite covariance matrix, $\mathbf{\Sigma}$. The maximum lag length (p) is determined empirically using conventional model selection criteria.

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[7] In principle it would have been preferable to use manufacturing input costs as a measure of other costs. Since this measure turned out to be statistically insignificant (possibly reflecting presence of oil explicitly in the model)) we opted for an index of labour costs to capture non-commodity costs in the food manufacturing and retailing sectors.

[8] Variables are expressed in natural logarithms. Data definitions and sources are provided in the appendix.

While (1) captures the dynamic correlations between the variables succinctly, the VAR is difficult to interpret economically. Where the variables form co-integrated relationships, then (1) is more conveniently expressed in its vector error correction (VEC) form,

$$\Delta \mathbf{x}_t = \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{x}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \mathbf{x}_{t-i} + \boldsymbol{\Psi}\mathbf{D}_t + \boldsymbol{\varepsilon}_t \qquad (2)$$

in which the co-integrated relationships are explicitly parameterised by the matrix $\boldsymbol{\beta}$, coefficients of which provide estimates of the usual long-run response elasticities. In the empirical analysis, trace and maximal eigenvalue statistics are used to assess the number of co-integrating relationships among the data. Equation (2) also defines a matrix of error correction coefficients $\boldsymbol{\alpha}$, elements of which load deviations from equilibrium (*i.e.* $\boldsymbol{\beta}'\mathbf{x}_{t-1}$) into $\Delta \mathbf{x}_t$ for correction, thereby quantifying the speed at which each variable adjusts to maintain equilibrium. The matrices of coefficients $\boldsymbol{\Gamma}_i$ for $i = 1, \ldots, p-1$ capture the short-run effect of shocks to the variables on $\Delta \mathbf{x}_t$ and thereby allow the short and long-run responses to differ.

(ii) Cointegration analysis

The empirical model to be estimated is a seven-equation vector error correction (VEC) model, consisting of UK food prices and the six potential drivers as set out above. The system is estimated using the least generalised variance estimator available in *Time Series Modelling 4.31* using *Ox* version 7.00 (Davidson 2014, Doornik 2012)[9] over a sample consisting of 259 monthly observations in logarithmic form, spanning the period September 1990 to March 2012. All these series are non-stationary, exhibiting stochastic trends, and our first step is to test for the existence of cointegrating long-run relationships. Results for Johansen cointegration tests in a C-VAR with seven lags, as chosen by the Schwartz Bayesian selection criterion, are shown in Table 1.

[9] The least generalized variance estimator is equivalent to Gaussian maximum likelihood. All data and summary computer output is available upon request.

Table 1: Co-integration Test Statistics [*p* values]

	Johansen tests of H_0: rank = r		Trend test given rank = r.
r	Maximal Eigenvalue	Trace	$\chi^2(n\text{-}r)$
0	46.9 [<0.05]	149.6 [<0.01]	57.3 [0.000]
1	41.4 [<0.05]	102.7 [<0.025]	20.1 [0.003]
2	29.3 [<0.2]	61.3 [<0.2]	16.9 [0.005]
3	20.0 [<0.5]	32.0 [<0.2]	14.0 [0.007]
4	7.0 [<1]	12.1 [<1]	14.0 [0.003]
5	4.6 [<1]	5.1 [<1]	5.7 [0.059]
6	0.4 [<1]	0.4 [<1]	3.3 [0.069]

For the maximum eigenvalues and trace tests, the square brackets contain upper bounds on the p-values according to the tabulated critical values for these tests. The final column shows tests for the existence of a deterministic trend, given each value of the cointegrating rank. These are χ^2 with *n*–*r* degrees of freedom on the hypothesis of cointegrating rank *r* and no drift, with asymptotic p-values shown.)

These tests point to the presence of two cointegrating relationships at conventional levels of significance. Examination of the unrestricted estimates suggests the first is a (vertical) price transmission relationship (denoted $\hat{\boldsymbol{\beta}}_1'\mathbf{x}_{t-1}$) between raw commodity and retail food (w_t and r_t) augmented by the exchange rate (e_t) and supply and demand shifters (c_t and u_t respectively) and the second (denoted $\hat{\boldsymbol{\beta}}_2'\mathbf{x}_{t-1}$) a (horizontal) relationship between the dollar price of oil (o_t) and the food commodity index (w_t). Normalising the coefficients of the first cointegration relation on retail food prices and the second on world food commodity prices and excluding statistically insignificant estimates yields results reported in Table 2.

Table 2: Long Run Elasticities [p values]

Elasticity of UK retail food prices (r_t) with respect to:

World food commodity prices (w_t)	0.57
	[0.00]
Exchange rate (e_t)	–0.45
	[0.00]
Labour cost shifter (c_t)	0.25
	[0.03]

Unemployment rate (u_t)	–0.15
	[0.15]
Elasticity of world food commodity prices (w_t) with respect to:	
Oil prices (o_t) pre-1999(3)	0.49
	[0.00]
Oil Prices (o_t) post-1999(3)	0.58
	[0.00]

As can be seen from the table, significant long-run influences on domestic food prices are found, notably: world food commodity prices, the exchange rate, labour costs and (barely significantly) with the demand proxy, unemployment. As far as the long run is concerned, domestic agricultural prices contribute nothing to retail price formation over and above that exerted by international commodity prices. While indicative of the UK's position as an open economy reliant on commodity trade, it is worth noting that domestic producer prices remain in the model owing to their highly significant role in the short run determination of retail food prices; the implication here being that the effect of trade on price adjustment is not instantaneous but takes time so that in the interim, domestic agricultural prices convey information that is pertinent and distinct from the signals emanating from world prices.[10] The elasticity of retail food prices with respect to world food commodity prices – the long run price transmission elasticity – suggests that, other drivers held fixed, a 10% increase in agricultural prices on the world market is associated with a 5.7% increase in retail food prices in the long run. By this measure, price transmission between commodity and retail markets is thus not one-for-one, reflecting the stabilising influence of non-agricultural components in retail food prices. The results also point to the important role played by exchange rates in domestic food prices: a 10% appreciation (depreciation) in the value of Sterling against the dollar being associated with a long-run 4.5% fall (rise) in retail food prices, *ceteris paribus*. The similarity of these two elasticities suggests that the domestic effect of changing world prices is broadly similar irrespective of the source of the dollar price change, as might be expected. The supply and demand shifters that augment the price transmission relationship have somewhat smaller effects on food prices, and suggest that a 10% increase in demand (as measured here by the rate of unemployment) and supply (labour cost) shifters lead to *ceteris paribus* long-run effects on food prices of -1.5% and 2.5% respectively.

[10] The oil price, while not significant in the first cointegrating relation, is important indirectly, through its effect on world prices, and forms the second cointegrating relation, as discussed in the main text.

The second cointegrating relation is a simple bivariate linkage between the dollar-denominated world prices of food commodities and oil. This long-run relation brings the price of oil (that was insignificant in the price transmission relationship) explicitly into the model via its co-movement with international food commodity prices. Estimates in Table 2 suggest that in the post-2000 period, a 10% increase in oil prices has been associated with a 5.8% *ceteris paribus* increase in world agricultural commodity prices. The somewhat smaller increase (4.9%) prior to that date suggests that commodity prices have indeed become more sensitive to energy prices in recent years.[11] While relationships between international commodity prices are not our principal focus, incorporating them in a second co-integrating relation does mean we are able to quantify the impact of oil prices on domestic food inflation, an aspect of policy relevance during the oil price boom and bust the sample spans. We address this issue formally using impulse response analysis in the following section and merely note here that chain-linking the long-run elasticities implies that, in the post-2000 period, a 10% change in oil prices leads to a (5.8×0.57=) 3.3% increase in food prices *ceteris paribus*. As with the other drivers, the response of food prices to oil price shocks is inelastic and, according to the estimates in Table 2, not dissimilar in magnitude to the impact of changes in the exchange rate.

In the final version of the VEC, the coefficient matrices in equation (2) are subject to a large number of restrictions, based both on the outcome of significance tests, and what is suggested by the data characteristics and the underlying economic relationships.[12] Specifically, the monthly seasonal dummies are included in the short-run equations only for UK retail food prices and UK agricultural prices, the other series being either seasonally adjusted or non-seasonal. Small and insignificant estimates in the matrices of short run effects $\boldsymbol{\Gamma}_i$ are also set to zero. Of particular interest is the matrix of error correction coefficients $\boldsymbol{\alpha}$. While corroborating the existence of any cointegrating relationships selected by the trace and maximal eigenvalue statistics they are also of interest in their own right, owing to the fact that they offer a useful summary of the speed at which the system adjusts when out of equilibrium. Results suggest that the first cointegration relation ($\hat{\boldsymbol{\beta}}_1'\mathbf{x}_{t-1}$) enters the food

[11] The switch in the elasticity of world commodity prices with respect to the price of oil is captured by including a dummy variable in the cointegrating relation defined as zero up to the break date and unity thereafter, although this refinement is not shown explicitly in equation (1). Experimentation suggests that the precise date of the structural change around this time has little effect on the estimates and their statistical significance. The 1999-3 breakpoint has been chosen on the basis of model selection criteria (SBC) rather than formal testing, so *p*-values reported in the table should be treated with caution.

[12] Diagnostic checks indicate model adequacy at conventional levels of significance. Conditional moment tests, with asymptotic p-values in brackets, are as follows: autocorrelation: $\chi^2(49) = 54.28$ [0.28]; functional form: χ^2 (49) = 53.38 [0.31]. (Note, these are whole-system tests. The test degrees of freedom are accounted for by seven test regressors, lagged residuals or squared fitted values, in the seven equations.)

price with an estimate of −0.05 that is correctly signed for stability of the system and significant with a p-value of 0.005. The second cointegration relation ($\hat{\boldsymbol{\beta}}'_2\mathbf{x}_{t-1}$) enters the world food price equation, and is estimated at −0.04, again correctly signed and with a p-value of 0.013. Together these error correction coefficients allow long run linkages in international commodity markets to permeate in to UK food prices. At a rate of 5% per month they imply a seemingly sluggish adjustment, a feature that reflects that they are (a) averages and (b) predicated on the *ceteris paribus* clause, limitations that we now relax.

4. Retail Food Price Dynamics

4.1 Methodology

Since the variables included in the model can be reasonably treated as predetermined (i.e. contemporaneously exogenous) to UK food prices, we undertake an impulse response (IR) analysis to trace the dynamic effect of shocks using a standard (rather than orthogonalised) impulse response function. In doing so, we adopt methods of stochastic simulation given the non-linear structure of the model that is induced by the structural break in the second cointegrating relationship (see, for example, Koop *et al.*, 1996). While more computationally complex than other approaches, it does rather easily facilitate the simulation of a broad spectrum of shocks, since being non-linear it does allow us to investigate the effect of commodity shock duration on UK food prices. Specifically, the standard impulse response functions are computed by Monte Carlo simulation using the following procedure:

Step 1. Model residuals are randomly resampled to provide the shocks. Here we use 1,000 simulations, the sample period being used to provide initial conditions.

Step 2. Multi-step forecasts are computed by the Monte Carlo method, the median of these 1,000 runs providing the point forecasts that represent the baseline profile. If desired, the confidence bands can also be calculated, these being given by the relevant quantiles of the Monte Carlo distribution.

Step 3. A unit shock is introduced into the equation to be perturbed (such as the world food commodity price equation) by adding a zero-one dummy, either as an impulse (one-period) shock that takes the value 1 in the first forecast period and zero otherwise, or as a step-change that takes the value 1 up to the required forecast horizon and zero otherwise. In this approach, the dummies introduce shifts in the equation intercepts, and are supplied with coefficients to fix the desired magnitude of the shock. Since the equations are in logarithmic form, perturbing the intercept can be viewed as shifting the model solution by a factor of proportionality. Thus, adding log(1.1) = 0.095 to the intercept has the effect of shifting the mean path of the process by 10%. Other magnitudes can be incorporated in a similar way.

Step 4. Similarly to Step 2, the Monte Carlo exercise is re-run using the equation in Step 3 to compute point forecasts (and confidence bands if required) for the perturbed values.

Step 5. An impulse response curve is then estimated by computing the difference of the two median paths, subtracting the perturbed values obtained in Step 4 from the baseline ones obtained in Step 2.

It is important to emphasize that this method of introducing a shift ignores the observed correlation between the model disturbances measured by $\mathbf{\Sigma}$ and thus the construction of the impulse response function differs in principle from alternative procedures (described, for example, by Lütkepohl (2004) Section 2.3.2), in which the model equations are rotated so that the shocks are orthogonal; in effect, this shocks all the equations so as to reproduce the effect of a disturbance in the target equation uncorrelated with the others. One of the difficulties with this approach is that it requires the assignment of a contemporaneous causal ordering of shocks, so that the upper triangular factorization of $\mathbf{\Sigma}$ is computed appropriately. This choice is necessarily somewhat arbitrary. By contrast, the method we adopt avoids the need to impose a causal ordering and can be thought of as measuring the effect of a shift in one driver in isolation, rather than of an orthogonal disturbance. As a consequence, the interpretation of the results is arguably also more straightforward than the alternative; in our approach shocks emanate from the empirical distribution of model residuals rather than a hypothetically derived and artificially generated disturbance, as is the case with orthogonalisation.[13]

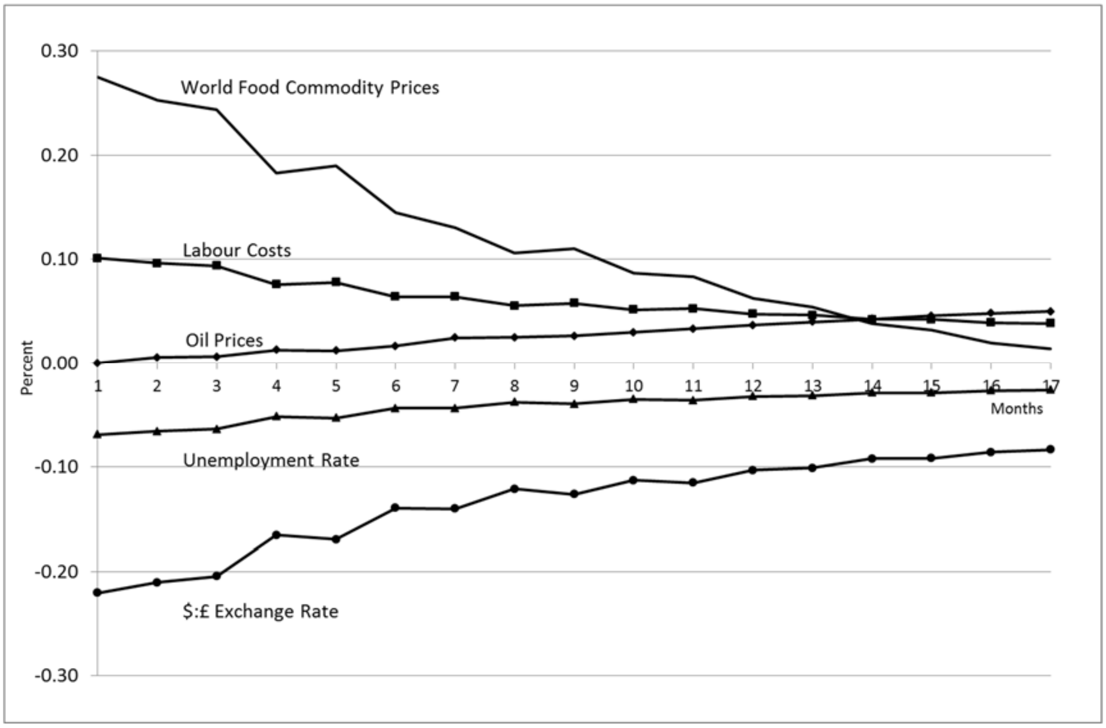
4.2 Food Price Drivers and Response Dynamics

Using the procedure outlined above, Figure 2 illustrates the dynamic effect of a 10% one-period shock in each driver on the food price index in the 18 months after the shock, the effects being expressed as a percentage of the predicted food price level in the absence of the shock. Each impulse response function measures a separate experiment (i.e. a 10% shock to each driver) and are plotted together merely for convenience[14].

Figure 2: The Percentage Change in Food Prices Following One Period 10% Shocks

[13] See Koop *et al.* (1996) on this point but notice that since the shocks are not orthogonalised they cannot be given a structural (*i.e.* causal) interpretation and hence the response functions merely represent the model's best estimate of what would happen following a shock to an individual equation.

[14] To facilitate comparison of shocks from different drivers of food inflation, we do not report the estimated confidence intervals for each impulse response function.

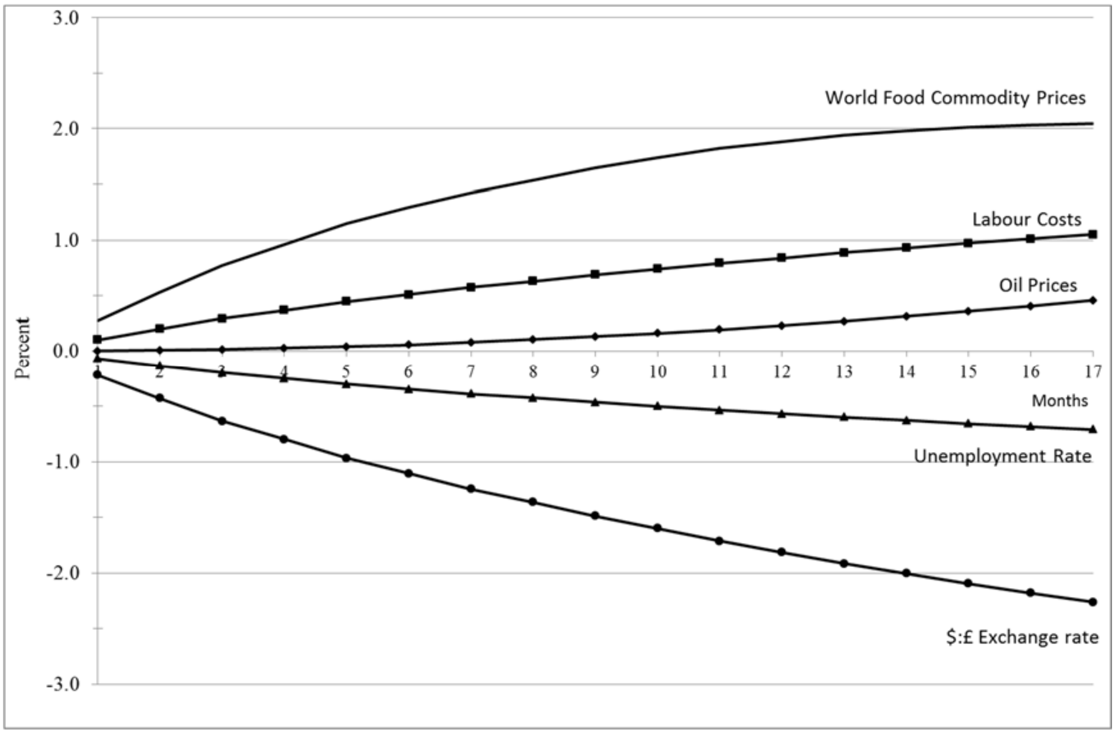


As can be seen, shocks to world food commodity prices and exchange rates have the largest quantitative impact on food prices with the maximum impact occurring in the month following the shock. Specifically, a one-period 10% increase in world food commodity prices is estimated to increase food prices by 0.3% in the month immediately following the shock, an impact that diminishes to 0.06% (effectively zero) a year after the shock. As with the long run elasticities discussed previously, the effect of exchange rate shocks is quantitatively similar albeit opposite in sign; a one-month 10% appreciation in Sterling depresses food prices by an estimated 0.2% in the month following the shock and by 0.1% one year later. One-period shocks to labour costs and unemployment produce similar patterns but with quantitatively smaller impacts. In contrast, oil price shocks appear to have negligible effects on food prices in the short run, building momentum only slowly over time, a feature we explore in more detail below.

Subject to the usual caveats regarding non-marginal changes, the effect of larger shocks, such as those generated by typical commodity price spikes, can be inferred from the graph simply by multiplying by the appropriate scalar. For example, a 50% increase in world food commodities prices for one month shifts the impulse response function by a factor of five, increasing food prices by nearly (5×0.28=) 1.5% in the month following the shock and by 0.3% a year later. While not inconsequential, the effect of such a large shock seems small. Furthermore, similarly modest effects result following shocks to the exchange rate, and even more so for the other drivers. However, the puzzle is more apparent than real since these seemingly modest effects merely reflect the short-lived (one-period) duration of the shocks. To illustrate, we repeat the experiments, this time simulating the effect of a permanent 10% increase in each of the drivers on the food price level (Figure 3). Given our modelling framework, modifying the shock duration is straightforward since it merely involves creating dummy variables that switch on for various lengths of time.

In each simulation, the shock shifts the driver to a new level that is permanently 10% higher, with the result that effects last for longer and are considerably larger in magnitude (note the different vertical scales to Figures 2 and 3). For example, a permanent 10% shock in world commodity prices leads to an initial 0.28% increase (replicating the result of a one-period shock) which then continues to grow, peaking at 2.0% some 18 months later. Hence, the effect of a 10% commodity price shock differs by a factor of (2.0/0.28 =) 7.1 depending on its duration. Similarly amplified effects are predicted when permanent shocks to the other drivers are simulated. While the simulations presented in Figure 3 assume permanent shocks (and so are more akin to shifts rather than shocks) of 10%, it is easy to see how persistent shocks of the magnitude experienced in recent commodity price crises might induce double-digit food inflation, something that we now explore in greater detail below.

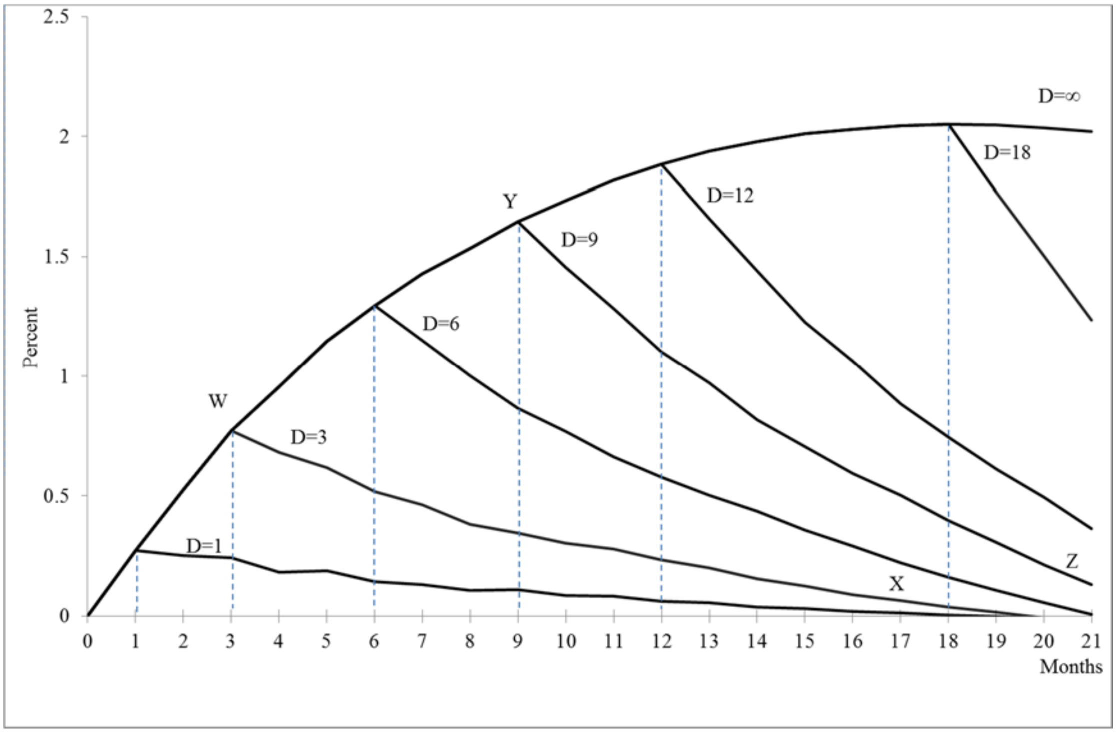
Figure 3: The Percentage Change in Food Prices Following Permanent 10% Shocks



4.3 Commodity Price Spike Duration and Food Inflation

The one-period and permanent commodity price shocks portrayed in Figures 2 and 3 underline that it is not just the size of the shock that matters for domestic food price inflation but its duration too. Since the figures represent polar illustrations (i.e. one-period and permanent shocks) it is useful to gauge the impact of some more empirically-relevant intermediate cases. Figure 4 presents the percentage response of domestic food prices to a 10% shock in world food commodity prices of various durations, with D=1 and D=∞ representing the impulse response functions of one-period and permanent shocks to world food commodity prices displayed in Figures 2 and 3. Referring to Figure 4, the effect of a 10% shock to commodity prices that persists for three months is estimated to gradually increase UK food prices peaking at point 'W' and subsiding thereafter along the line 'WX'. In a similar fashion, a 10% commodity price shocks that persists for nine months is estimated to peak at point 'Y' and decline along the line 'YZ'.

Figure 4: Percentage Effect on UK Food CPI of a 10% Shock to World Food Commodity Prices by Duration of the Shock



Despite the rather synthetic appearance of the response to the shocks, what is apparent is that the size and persistence of the effect on food prices increases with duration of the shock; the impact developing during the period in which the shock persists and declines thereafter as the shock becomes more distant. This dictates that the maximum effect on food prices does not occur at some fixed lag length, say five months after the shock, but varies with the duration of the shock; short-lived shocks creating peaks in food prices that are more immediate than with long-lasting shocks. Note also that persistent shocks, such as those lasting more than nine months or so are more akin to a permanent change than one-off shocks, suggesting that a typical commodity price spike (the 2008 spike lasted around 15 months) is likely to induce a response towards the top end of the magnitudes estimated.

The model may also be used to assess the size of commodity price spike that would be required to induce a particular level of food inflation in the UK. For example, it may be of interest to estimate the size of commodity shock required in future to reproduce the domestic food price inflation observed in 2008 or indeed 2011. Given the foregoing analysis, the answer necessarily depends on the duration of the commodity shock. Dealing first with the 2008 episode, which saw retail food inflation rise to 13% from a base of around 3%, estimates from the model suggest the world food commodity price index would need to rise (10/0.028=) 357% if it were a one-month shock, (10/0.075=) 133% if the shock were to last for three months or (10/0.20=) 50% if the shock persisted for 15 months – in fact, not dissimilar to the 58% rise in the commodity food price index reported in the Introduction. Using similar reasoning, reproducing the inflationary episode in 2011, in which retail food inflation rose by around 5 percentage points (from 2 to 7%) would have required a 33% increase in food commodity prices (assuming a commodity price shock lasting 8 months) a little under the 41% increase actually observed. While the precise duration of the shocks is debateable and the simplistic nature of the model undeniable, it does go some way to account for the marked differences in the inflationary consequences of the two commodity price spikes observed in the sample period. Less contentiously perhaps, the model does underline the message that it is both the magnitude and duration of commodity price shocks that matter for retail food inflation.

One other point is noteworthy. Sizeable though the responses to persistent commodity shocks are, they are considerably lower than those implied by the long-run elasticity (reported in Table 2), a commonly used metric of pass-through. The extent to which estimates of the permanent shock differ from the corresponding long-run elasticity depends on the importance and nature of the interactions among the variables in the system, which are incorporated in the impulse response analysis and ignored by the elasticities. Results suggest that these interactions tend to dampen the effect of world commodity price shocks; the predicted maximum impact on food prices (2.0%) is less than one-third that implied by the long run elasticity (5.7%). As a result, a (say) permanent 50% increase in world food commodity prices is predicted to raise food prices to a peak that is at most (5×2.0=) 10% above their preshocked level, considerably less than the (5×5.7=) 28.5% implied by the long-run elasticities. The key point here is that not only are the interactions between variables in the system quantitatively important they also tend to moderate, rather than exacerbate, commodity price shocks.

4.4 Assessing the Drivers of UK Food Inflation

In time series applications involving VARs, decomposing the forecast error variance provides a convenient summary of the relative importance each variable to the evolution of all other variables in the system (see for example, Lütkepohl (2006), section 2.3.3). Table 3 reports the contribution of each variable to domestic food prices at various points in the forecast horizon based on estimates from the model. Entries report the relative importance of shocks from each source so that each row sums to one. Moving down each column therefore traces the relative importance of each variable in food price variation over time.

Treating the figures at 1, 12 and 36 months as short, medium and long run respectively, estimates suggest that in the short term, idiosyncratic shocks to food prices tend dominate those from the drivers, reflecting that shocks take time to permeate in to retail food prices. Domestic agricultural price shocks are the first to register in food prices but never account for much more than 13% of their variation and are overtaken by the influence of world commodity prices, which account for around one-third of food price variation in the longer term.

Table 3: Forecast Error Variance Decomposition of UK Retail Food Prices

Months	UK Food Prices	UK agricultural prices	World commodity prices	Exchange rate	Labour costs	Unemployment	Oil prices
1	0.99	0.00	0.01	0.00	0.00	0.00	0.00
6	0.72	0.12	0.06	0.04	0.02	0.02	0.02
12	0.50	0.13	0.19	0.12	0.01	0.02	0.02
24	0.25	0.08	0.36	0.19	0.01	0.01	0.11
36	0.14	0.05	0.36	0.19	0.00	0.01	0.26

Like world commodity price shocks, exchange rates manifest in the medium term, eventually accounting for around one-fifth of food prices, whereas the proxies for manufacturing (labour) costs and consumer demand (unemployment) contribute little at any forecast horizon.[15] As noted earlier, the effect of oil prices is distinct from the other drivers, in that despite having virtually no effect in the short term (echoing the findings of Baumeister and Kilian, 2014) it is found to play an influential role in the variation of food prices at more distant horizons, something that is captured by the cointegrating (long-run) specification of the estimated model. What is also apparent from these results is the important impact of international factors on UK food prices, domestic factors only accounting for one-fifth of food price inflation over the sample.

5. Conclusions

We have highlighted two issues that are pertinent in addressing retail food price inflation and the links between domestic retail prices and world commodity prices. First, world agricultural prices are not the sole driver of domestic retail food prices; other factors matter too. To fully assess what factors drive domestic food inflation, we employed a seven variable cointegrated vector autoregression (C-VAR). With reference to the UK food inflation experience over the 1990-2012 period, a notably turbulent period in recent history, we show that world agricultural prices are indeed an important determinant of food prices but no more so than exchange rates and oil prices, the latter emerging only in the long run and hence unlikely to be captured by models that do not accommodate equilibrium relationships. Variables included in the model to proxy for food demand and industry costs were found to have statistically significant but quantitatively small effects on food inflation during the sample frame, although this need not be the case in other periods. Despite the reduced-form nature of the estimated model, embedding the price transmission relationship in a richer, more theory-consistent, framework does allows the role of other factors to be assessed in both the short and longer term. Second, we highlight that the dynamic characteristics of commodity price spikes affect significantly the inflationary effect; it is not just the level of prices reached in a 'spike' that matters but also the duration of the spike itself. For any given lag structure that determines the pass-through effect, the duration of the spike is a key determinant of the final effect on domestic food prices. Despite the seemingly obvious nature of this conclusion, little attention has been paid to it in the past, an outcome that may reflect the impulse response methods commonly employed to evaluate commodity price transmission. While our results

[15] The statistical significance of labour costs (and to a lesser extent) unemployment in the cointegration analysis suggests that the reason for their lack of contribution to UK food inflation is due to a lack of variation over the sample period. In periods where these factors are more prominent, their contribution will be greater. Of course, the imperfect nature of the proxies used may also play a confounding role but in the absence of better data, it is not possible to be more decisive on this.

pertain solely to the UK's experience, it seems likely that the principles might apply more generally where the aim is to gauge the domestic retail price impact of events on world commodity markets.

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Appendix

Definition	Source
UK Consumer Price Index (all items).	Office for National Statistics (ONS)
UK Consumer Food Price Index.	OECD, OECD Statistics. http://stats.oecd.org/index.aspx
UK Retail Price Index (all items).	Office for National Statistics (ONS)
World Food Price Index	IMF Primary Commodity Prices: http://www.imf.org/external/np/res/commod/index.asp
$:£ Exchange rate	IMF Financial Statistics
Agricultural Producer price index (UKAPPI).	UK DEFRA
Average Earnings index for the whole economy s.a.	Office for National Statistics (ONS)
Oil Price ; UK Brent, light blend 38 API, fob U.K.	IMF Primary Commodity Prices: http://www.imf.org/external/np/res/commod/index.asp
Unemployed: UK (Thousands) s.a.	Office for National Statistics (ONS)