Preliminary

Do oil prices drive food prices?

A natural experiment

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Abstract

High commodity prices have attracted close attention recently. Aside of a variety of macroeconomic explanations, some specific microeconomic factors have been proposed as the cause of a previously non-existing connection between energy prices and food prices. Specifically, ethanol promotion policies in the United States would have created a link between oil and corn prices that would be the cause of the recent rally in the price of that crop and its substitutes (especially soybeans). Even though it is intuitively appealing, one problem with this hypothesis is that ethanol policies have been in place in the US for more than 35 years, whereas the run up in food prices dates back only to 2006. I interpret a significant change in US biofuel policy during 2006 as a natural experiment that could help in identifying changes in the stochastic properties of the corn and soybean price processes. The results are sharp but to some extent unexpected: there are substantial changes in the dynamic properties of corn and soybean prices time series, they are more closely related to oil prices, but the predictive causality seems to run in reverse, from the crops to oil prices.

JEL classification: H23, O13, Q16, Q48

Keywords: oil price; food prices; ethanol; biofuels; energy policy.

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1. Introduction

The remarkable rally in commodity prices since the start of the new century, even in the face of the most acute financial and economic crisis since the Great Depression, has raised the concern of policymakers and commentators alike. Volatility has increased significantly since 2008, but after a sharp drop in the last quarter of 2008, commodity prices have quickly returned to levels that are close to, or in some cases higher than, the already lofty peaks reached before the collapse of Lehman Brothers.

Several macroeconomic explanations for the widespread rise in the prices of all sorts of commodities have been proposed.[2] The usual suspects are the expansive monetary policy pursued by central banks since the onset of the crisis in the second half of 2007, the growing financialisation of commodity markets, and increased demand for basic materials from rapidly growing emerging markets. All of these hypotheses have merits but also theoretical drawbacks, and the empirical evidence is not conclusive. However, a clear understanding of the circumstances and causes of the current strength of commodity markets is relevant for both developed and emerging markets. For developed markets, typically net importers of commodities, such understanding would allow them to adjust policy in order to counter a substantial drag on growth, and a worrying source of inflationary pressures. For many emerging markets, especially in Latin America, high commodity prices underpin the solid fiscal and external positions that these countries have enjoyed in the last few years, and such understanding would help policy-makers assess the risk for their own macroeconomic outlook.[3]

In this paper, I step out of the typical emphasis on macroeconomic drivers and, focusing on agricultural commodity prices, I explore the impact that energy policy in advanced economies (eg the promotion of biofuels) has had on food markets. Biofuels have long been considered a potential source of disruption in the market for those crops that are basic inputs for their production (mainly corn and soybeans). However, there are few quantitative assessments of their relevance and the nature of their impact. Zhang et al. (2009), using data through December 2007, found not long-term relationship between oil and food (corn and soybeans) prices. In the short run, they found no relationship, although before 1999 causality seems to run in reverse of the expected direction (from crops to oil and ethanol). Marshad and

[2] See Frankel and Rose (2010) for a summary. Also IIF (2010).

[3] See, for instance Avalos (2011), Cecchetti and Moessner (2008).

Hameed (2009), using a longer sample, find evidence of a long term relationship between oil, corn, wheat and rice, with causality flowing from the fuel to the crops. They relate this effect to cost factors, namely, the growing reliance by modern agriculture on seed fertilizer technology that is highly dependent on chemical inputs derived from oil. They also argue that biofuel production is another dimension of the problem. But they focus on the indirect effect of acreage competition between wheat and corn. As explained below, that is probably not a major factor for wheat (as it is for soybeans) since wheat and corn have limited land overlap. Trujillo-Barrera et al. (2011) study volatility spillovers in the US from energy to agricultural markets in the period 2006-2011. They found significant spillovers from oil to corn and ethanol markets, which seem to be particulary strong in high volatility periods for oil markets. They also identified significant volatility spillovers from corn to ethanol markets.

Being the largest producer of corn, and also the place of one of the longest running programs for the promotion of corn-based ethanol, the US and its energy policy are natural focal points of concern. Here, I attempt to exploit a natural experiment arising from a significant change in 2006 on the nature of ethanol policies in the US to assess the relationship between oil, corn and soybean prices. The results are at once unsurprising and intriguing: price dynamics in those two staple crops have changed significantly since 2006, but not in a way entirely consistent with biofuel critics' ex-ante concerns.

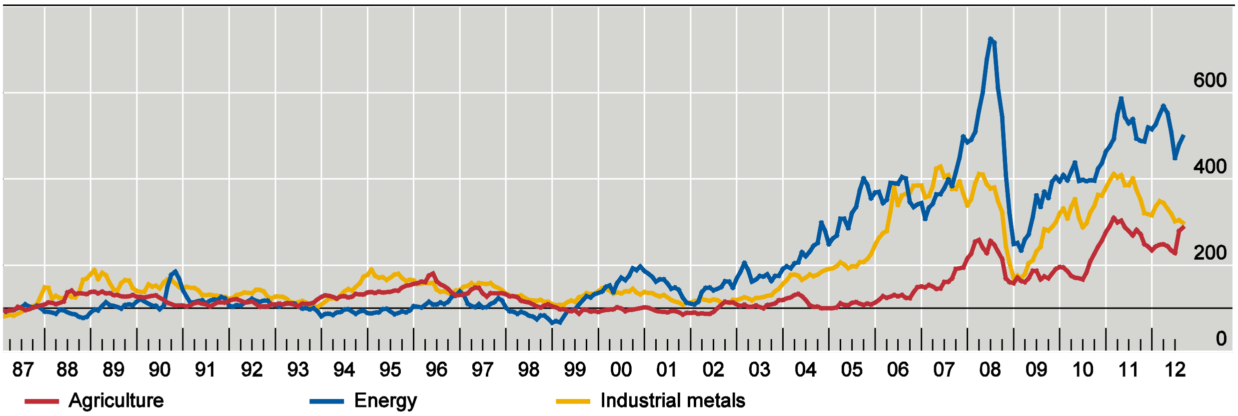
The paper is organised as follows: section 2 reviews some of the main stylised facts about commodity prices in the last 30 years; section 3 explains as clearly and briefly as possible the sequence of ethanol-promotion policies in the US, and how the aforementioned natural experiment arose. Section 4 describes the data, and analyses the time series properties of oil, corn and soybean prices before and after the breaking point mentioned. Section 5 concludes and indicates potential directions for further research.

2. Some stylised facts

For almost 20 years, starting in the early 1980s, the main classes of commodities (agriculture, industrial metals, and energy) traded in relatively narrow price bands, without any clear trend. That changed in the late 90´s, when nominal prices began rising for energy (essentially oil) and industrial metals. In roughly 6 years, those commodities increased their prices four-fold, and continued rising. Agricultural commodities took off much later, around 2006 (Graph 1). Since then, food prices have outpaced those of other commodity classes, retreating less following the financial collapse of 2008, and recovering all their losses and more afterwards.

Commodity prices[1]

1987 = 100 Graph 1



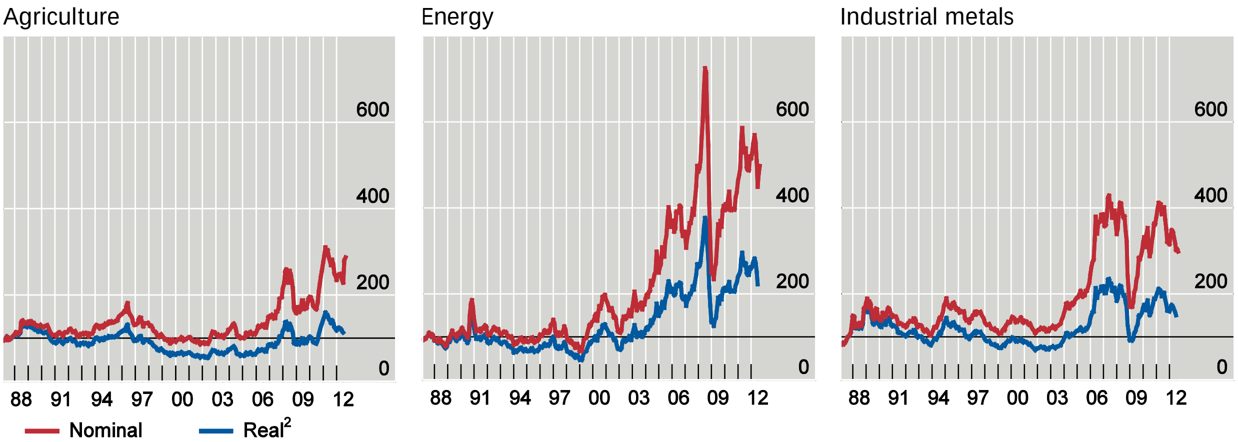
[1] S&P GSCI spot price indices, monthly average.

Sources: Standard & Poor's'; Datastream; BIS calculations.

Nominal prices can be deceptive though, especially over long horizons. Turning to real commodity prices, deflated by the US CPI, the price increases look less sharp, but they are hardly moderate (Graph 2). By early 2012, the real price of energy and industrial metals had roughly doubled since the mid-1980s, while real agricultural prices rose about 50%.

Nominal and real commodity prices[1]

1987 = 100 Graph 2



[1] S&P GSCI spot price indices, monthly average. [2] Deflated using the United States CPI.

Sources: Standard & Poor's'; Datastream; BIS calculations.

In several ways, the prices of agricultural commodities have behaved differently from those of energy and industrial metals. Graph 2 above shows that the run-ups in both nominal and real prices have been more moderate in agriculture than in the other two sectors. Moreover, in real terms, the peaks recently reached by agricultural commodities are maximums for the series, but they do not look exceptional. However, energy and industrial metal real prices reached levels vastly higher than anything observed in the last 25 years. After the initial collapse that followed the Lehman bankruptcy, agricultural prices had recovered and surpassed their 2008 levels by early 2011, whereas industrial metals and energy also recovered, but stayed significantly below their 2008 marks. This quick rebound, relatively stronger than in energy or metals is surprising, if nothing else because the supply of agricultural products is more flexible than that of mineral commodities: agricultural output can be increased (or cut) relatively quickly at relatively low cost, responding quickly to price swings, and in principle moderating price volatility. Several recent papers have remarked these unusual dynamics and attribute it to the growing integration between the oil and certain food commodity markets (more specifically corn and soybeans), fostered by the biofuel promotion policies of advanced economies.[4]

Biofuel promotion has for some time been a feature of energy policies in advanced and emerging market economies. Brazil and the United States operate two of the longest-running programmes, dating back to the 1970s. Brazil produces ethanol from sugarcane by a relatively energy-efficient process. On the other hand, the United States produces corn-based ethanol in a process that is generally regarded as less efficient, with limited net-carbon fuel savings.[5] The fiscal cost of federal ethanol subsidies in the United States was relatively small, ranging between USD 5 - 7 billion.[6] However, subsidies seemed to have a relevant impact on the industry economics, the size of the sector, and its demand of corn. On the other hand, although soy-oil can be an input in the production of biodiesel, the main connection between petrol and soybean prices would work through corn. There would exist two main transmission mechanisms from higher corn prices to higher soybean prices: first, the competition for planting acreage, since both crops share quite similar soil and climatic requirements. Moreover, corn and soybeans share several industrial uses (eg as animal feedstock) and substitution from pricier corn to soybeans could be another factor weighing on the latter´s demand and, eventually, price.

[4] For instance, see F M Arshad and A A A Hammeed (2009); Z Zhang, L Lohr, C Escalante and M Wetzstein (2009); A Trujillo-Barrera, M Mallory and P Garcia (2011).

[5] In other words, the unit cost of ethanol produced is too high (by international standards), and the amount of energy liberated by the consumption of ethanol so produced is roughly similar to the energy used to produce it. As a comparison, Brazil’s sugarcane ethanol yields roughly eight times more energy per unit of energy input. See B Yacobucci (2006).

[6] Federal subsidies and tariffs were allowed to expire as of 31 December 2011, but for reasons I will describe below, they had stopped being the main source of support for the ethanol industry a few years ago.

Disentangling the impact of ethanol-promotion policies on the price dynamics of oil, corn and soybeans would be complicated. Just tracking the marginal adjustments of subsidies at the federal and state levels over the span of almost 40 years would be a daunting task, with little potential gain. As it happens, chance has provided a natural experiment, in the form of a major policy change, which greatly simplified matters.

3. A brief history of ethanol

The ethanol industry in the United States received a major impulse with the Energy Policy Act of 1978, initially conceived as a response of the US Congress to the instability caused by the OPEC-induced oil shocks. This legislation granted the sector a tax exemption of 40 cents per gallon produced. Since then, the subsidy has ranged between 40 and 60 cents per gallon, irrespective of the price of corn or ethanol. And that is only at the federal level. There are also other state (and federal) subsidies: in 2006, the total effective subsidy was estimated at between USD 1.05–1.38 per gallon, depending on the state.[7] On top of the tax break, the federal government imposed a 54 cent per gallon tariff on imported ethanol (mainly from Brazil).

With time and lower real oil prices, concerns about pollution and global warming took precedence over energy security. In response, Congress enacted the Clean Air Act of 1990, which required vendors to ensure that their gasoline contained a minimum percentage of oxygen. This causes a more efficient combustion of fuels (ie increasing the energy produced by each gallon of gasoline burned). This was a crucial change in the nature of the energy policy, since it migrated from a pure structure of subsidies and tariffs, which predictably affected incentives, to a form of quantitative mandate whose full set of repercussions are harder to ascertain in advance. It was also the precedent for the policy change that, I argue below, caused a structural change in the corn market.

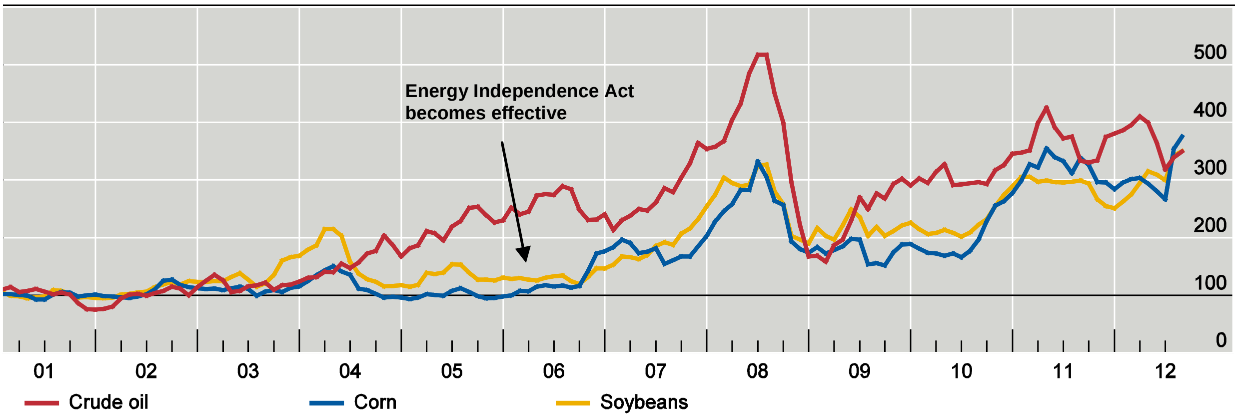
At the time, ethanol was only one of the additives that could be mixed with regular gasoline in order to increase its oxygen content. Another organic chemical compound, methyl tertiary butyl ether (or MTBE), was also widely used in the United States and for years represented the main domestic competition for corn-based ethanol. As MTBE is a petroleum or natural gas derivative, it was preferred in most non-agricultural regions because it was generally much cheaper than ethanol, as well as more widely available and easier to transport and

[7] See Koplow, D (2006).

distribute. However, the chemical properties of MTBE make it highly soluble in water, and very persistent (it can only be degraded by certain bacteria). Before long, reports started to surface of groundwater aquifers contaminated with MTBE from leaking oil tanks. The health effects for humans are not fully understood, but MTBE is a known carcinogen for animals. In short order during 2003-04, the two states with the highest MTBE consumption (California and New York) banned its use within state borders. By 2005, 19 states had partially or totally banned the use of MTBE. Mounting litigation was slowly convincing the oil industry about the hidden costs and risks of using MTBE. Once again, Congress responded by passing new legislation. The Energy Independence and Security Act of 2005 stopped short of a federal ban on the use of MTBE, but it went for a close substitute: it eliminated the oxygenate standard of the Clean Air Act as of May 2006 and replaced it with a new renewable fuels standard. This new standard required motor fuels to contain a minimum amount of fuel coming from renewable sources, such as biomass (eg ethanol), solar or wind energy. Needless to say, ethanol offered the only practical way to comply with the new standard. Therefore, as of mid-2006, MTBE was essentially finished, and ethanol became the only available gasoline additive.

Selected commodity prices[1]

2001 = 100 Graph 3



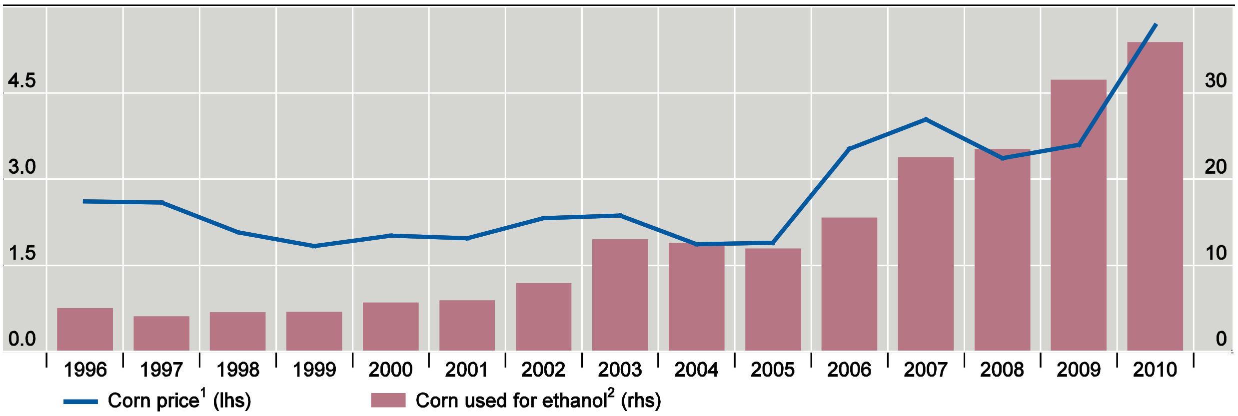
[1] S&P GSCI spot price indices, monthly average.

Sources: Standard & Poor's'; Datastream; BIS calculations.

Before the passing of the 2005 Act, critics had warned that corn prices would increase as a result. On the face of it, the price impact was swift. Since the early 80's, corn had traded in a relatively narrow band of USD 2–3 per bushel. In fact, during the 10 years before the Energy Independence Act came into force, the price had oscillated gently around USD 2 per bushel. Between August 2006 and February 2007, the price of corn almost doubled (+89%) and by June 2008 it had reached USD 6.54 per bushel, 205% more than the average price during the five years preceding the 2005 Act's implementation (Graph 3). Needless to say, the impact was not limited to corn prices, but also affected its close animal feed substitutes, mainly soybeans and, to a lesser extent, wheat. Even though their prices were typically more volatile than those of corn, soybeans traded in a range of USD 4–8 over those 20 years, with an average of USD 6. After the new renewable fuel standard become applicable, the soybeans price started a rally that by July 2008 brought it to a level almost 150% higher than the average of the previous 20 years.

Use of the United States corn for ethanol production and corn prices

1996-2010 Graph 4



[1] In US dollars per bushel. [2] As a percentage of the United States ethanol production.

Sources: United States Department of Agriculture; BIS calculations.

The apparent impact on quantities was also remarkable. In 2000, only about 5% of US corn production (the world´s largest) was used for ethanol production. This share jumped to almost 13% in 2003-2005, after California and New York banned the use of MTBE, and to almost 23% by 2007, the first full year after the Energy Independence Act came into force. By 2010, more than 35% of the US corn harvest was used to produce ethanol (Graph 4). In fact, the only use of US corn that has increased at all in the last 10 years is ethanol production, and most of it happened after 2003. All other uses (feed and residual, export, other industrial non-ethanol) have stayed roughly constant or even declined, although total US production has increased by almost 38% during this period (Graph 5). That is to say that most of the growth in US corn production since 2003 has been applied to the production of ethanol. The Food and Agricultural Policy Research Institute (FAPRI)[8] projects that the use

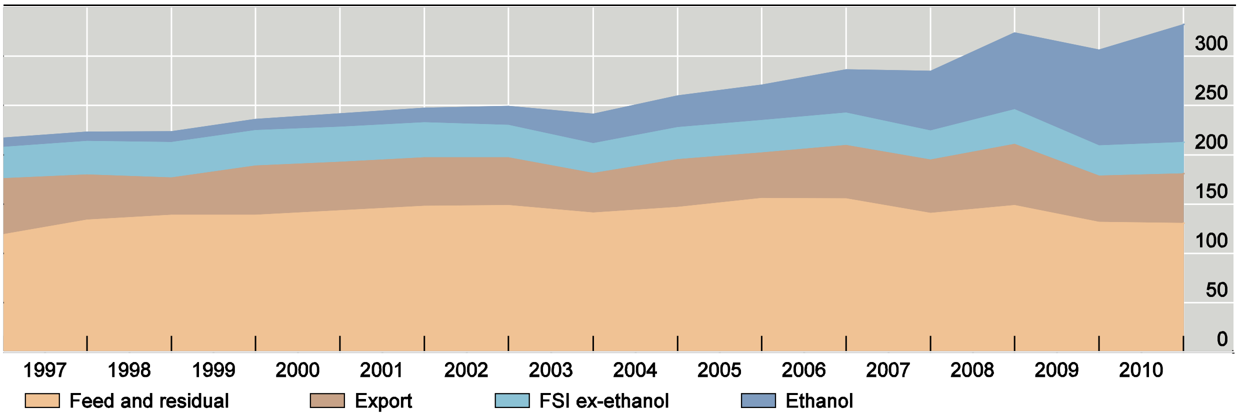
[8] This is a research programme established in 1984 by a Congressional grant to prepare baseline projections for the US agricultural sector and international commodity markets and to develop capability for policy analysis

of corn for ethanol production will stay largely constant at around 36% of the US harvest for the foreseeable future.

Use of the United States corn production

In millions of tons

Graph 5



Sources: United States Department of Agriculture; BIS calculations.

But US energy policy is hardly an exception. Worldwide government intervention has strongly favoured biofuel production, and probably will continue.[9] If the biofuel industry is having any significant impact on the market for corn and soybeans, that is unlikely to recede soon.

4. Empirical analysis

This section analyses spot price data on these three commodities to investigate whether their dynamic properties changed since the Energy Independence Act become effective. In particular, I will search for evidence that changes in corn and soybean prices have become more closely related with changes in oil prices. Once the renewable fuel standard became enforceable and MTBE stopped being a viable alternative to ethanol, the use of corn for ethanol production surged very quickly, and acquired a size large enough to have an impact on corn's global market. The intuition is straightforward: the higher the price of oil, the higher is the incentive of gasoline producers to bring to market blends with a higher content of

using comprehensive data and computer modelling systems for the world agricultural market. FAPRI is hosted by Iowa State University and the University of Missouri.

9 The EU has its own ambitious biodiesel programme, with subsidies and quantitative targets. Other countries are also developing their own, highly subsidised programmes, including Argentina, Canada, China, Colombia, India, Indonesia, Peru, the Philippines and Thailand. FAPRI estimates that these programmes will collectively almost double the output of ethanol (from several sources) over the next 15 years, while biodiesel production should increase by 45%.

ethanol. They bid up the price of ethanol, and ethanol producers in turn bid up the price of corn. As the price of corn increases, the prices of its close substitutes in other uses (industrial, animal feed, etc.) also increase. In the medium term, higher corn prices causes a larger share of arable land to be dedicated to the production of corn, restricting the supply of other crops with which it competes for acreage. Both mechanisms point to soybeans as the main receivers of these second round effects.

To test this intuition, I estimate time series models to analyse the long and short term dynamics of the prices of oil, corn and soybeans before and after May 2006, when the Energy Independence Act of 2005 became applicable.

The data

I use commodity benchmark daily price data from Datastream, starting on 1 January 1986 (the earliest available for all three commodities). I calculate monthly averages of daily data, to allow for some information aggregation and average out extreme observations. For the purpose of providing some context, I also review price data for other key commodities, in particular copper and gold. Table 1 present some basic statistics of the five commodities for the whole sample, and also for the two relevant subperiods: before implementation of the Energy Independence Act (January 1986 – April 2006) and after (May 2006 – April 2012). Mineral commodities trebled their prices from the earlier to the latter subsample, whereas the agricultural ones “only” doubled theirs. Price volatility also increased substantially, although the changes are less homogenous: the standard deviations of oil, copper and soybeans essentially doubled, corn trebled, and gold´s standard deviation multiplied by six. Normalizing the standard deviations of each sub-period by their respective means, we find that in fact only the volatilities of corn, soybeans and gold increased in the last subsample, while oil and copper almost halved theirs. Another symptom of the relatively more muted performance of agricultural commodities is that their pre-April 2006 maximum prices are very close to the averages of the next subperiod, while the average prices of the mineral commodities in May 2006 – April 2012 vastly exceed the highest prices of the previous subperiod.

The second half of Table 1 shows correlations between the prices of all five commodities in both subsamples. As expected, given the shared rally, price correlations increased in most cases. However, the correlation between the prices of oil and gold, and copper and gold decreased slightly, whereas the correlation between corn and soybean prices recorded a minimal increase. Oil and the crops were *negatively* correlated before April 2006, but in the second subsample their correlation becomes strongly positive (+0.41 for corn, and +0.44 for soybeans). Moreover, the correlation between the prices of gold and both crops also increased significantly in May 2006 – April 2012. The subsequent sections explore the connections between oil prices and crop prices systematically.

Descriptive statistics and correlation matrix

Table 1

	Oil	Corn	Soybean	Copper[2]	Gold
	Sample				
Statistic[1]	January 1986 to April 2012				
Mean	37.458	2.781	6.938	3840.587	521.127
Median	23.343	2.383	5.986	2672.138	384.913
Standard deviation	27.566	1.231	2.414	2576.155	339.416
Minimum	11.346	1.330	4.078	1377.376	256.164
Maximum	133.890	7.327	14.853	9880.938	1772.136
	January 1986 to April 2006				
Mean	24.580	2.297	5.950	2257.926	364.530
Median	20.311	2.263	5.667	1913.112	367.879
Standard deviation	11.422	0.539	1.127	839.605	60.792
Minimum	11.346	1.330	4.078	1377.376	256.164
Maximum	69.448	4.862	9.765	6389.900	609.735
	May 2006 to April 2012				
Mean	81.097	4.422	10.284	7225.724	1051.819
Median	77.289	3.695	10.158	7539.858	942.060
Standard deviation	20.585	1.482	2.607	1568.047	358.594
Minimum	39.015	2.064	5.144	3079.391	586.295
Maximum	133.890	7.327	14.853	9880.938	1772.136
Simple correaltion[3]	January 1986 to April 2006				
Oil	1.000				
Corn	–0.143	1.000			
Soybean	–0.065	0.622	1.000		
Copper	0.204	–0.012	0.173	1.000	
Gold	0.194	–0.076	0.061	0.278	1.000
	May 2006 to April 2012				
Oil	1.000				
Corn	0.314	1.000			
Soybean	0.435	0.701	1.000		
Copper	0.675	0.277	0.394	1.000	
Gold	0.165	0.294	0.291	0.208	1.000

[1] For oil, in US dollars per barrel; for corn and soybean, in US dollars per bushel; for copper, in US dollars per metric tonne; for gold, in US dollars per troy ounce. [2] Data available since June 1993. [3] Calculated over prices (in logarithms), first difference.

Sources: Datastream; BIS calculations.

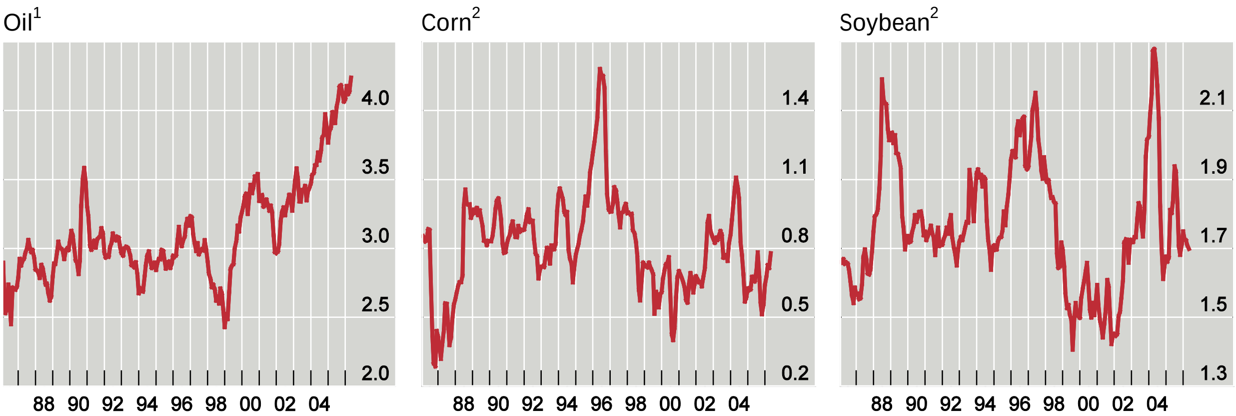
The baseline period: January 1986 - April 2006

Graph 6 shows the price paths (in logs) for the three commodities during the baseline period. On inspection, the data seem to reveal an apparent absence of trend in corn and soybean data, and also low persistence: prices seem to revert to the mean relatively quickly. That is mimicked by oil prices until early 1999, but then they start a strong rally which brings them to much higher levels than in the previous decade. Accordingly, standard unit root tests reject the non-stationary null hypothesis for corn and soy at customary 5% significance levels (Appendix Table A1). Correlograms of the first difference of both series suggest that the residuals are autocorrelated, so I also test the unit root hypothesis using the Phillip-Perron statistic, which yield similar results. On the other hand, oil prices are found to be stationary in first differences, and non-stationary in levels.[10] The coefficients of deterministic linear trends are non-significantly different from zero in most cases.

Oil, corn and soybean prices, January 1986 to April 2006

In logarithms

Graph 6



[1] In US dollars per barrel. [2] In US dollars per bushel.

Sources: Datastream; BIS calculations.

Based on these results, I estimate a vector autoregression (VAR) in first differences of all variables to investigate the short run interactions between these commodity prices during the baseline period. Differencing price data for corn and soybeans might seem unnecessary given their stationarity, but it makes for a more intuitive interpretation of the results.

[10] First differences of corn and soybean log prices strongly reject the respective unit-root hypotheses, indicating that they are stationary as well, as expected.

Moreover, the outcomes are not qualitatively different from those arising from a VAR that includes first difference of oil prices, and levels of corn and soybean prices.[11]

VAR model estimation results[1]

Sample (adjusted): March 1986 to April 2006[2] — Table 2

	Oil	Corn	Soybean
Oil (–1)	**0.201	–0.016	–0.029
	[3.307]	[–0.345]	[–0.738]
Corn (–1)	–0.152	**0.402	0.084
	[–1.547]	[5.273]	[1.333]
Soybean (–1)	0.112	0.032	**0.247
	[0.923]	[0.343]	[3.150]
Constant	0.005	0.000	0.000
	[1.040]	[0.007]	[0.090]
R-squared	0.060	0.177	0.109
Adjusted R-squared	0.048	0.167	0.098

** = significant at 5%; * = significant at 10%.

[1] Estimated over prices (in logarithms), first difference. [2] Included observations (after adjustments): 242.

Sources: Datastream; BIS calculations.

Table 2 displays the VAR estimation results. I include a single lag of the log price differences as suggested by both the Akaike and Schwartz information criteria. The model fit is poor, but the coefficients are precisely estimated. Table 3 presents block Granger-causality tests to assess whether there is a predictive causality relationship between these series. The objective is to determine whether oil prices Granger-caused corn (and soybean) prices even before the passing of the Energy Independence Act of 2005. This would be a purely statistical finding that would not answer the question of economic causality, which requires a structural analysis that I will not be performing here. Still, a positive finding would build the case for the existence of some sort of relationship unrelated to the specifics of US energy policy before 2006. The results are straightforward: none of price series seem to individually or jointly Granger cause the others, at standard significance levels. The tests strongly reject the hypothesis that corn prices in this sample were Granger caused by oil prices (or jointly by oil and soybean prices). Interestingly, the tests reject much less forcefully the opposite hypothesis: Granger-causality from corn prices to oil prices could not be rejected at a mere 12% significance level.

[11] These results are available from the author on request.

VAR model Granger causality tests[1]

Sample: January 1986 to April 2006[2]

Table 3

Excluded variable	Chi-sqared	Degrees of freedom	P-value
	Dependent variable		
	Oil		
Corn	2.392	1	0.122
Soybean	0.851	1	0.356
Both	2.395	2	0.302
	Corn		
Oil	0.119	1	0.730
Soybean	0.118	1	0.732
Both	0.228	2	0.892
	Soybean		
Oil	0.545	1	0.460
Corn	1.776	1	0.183
Both	2.646	2	0.266

[1] Tested over prices (in logarithms), first difference. [2] Included observations: 242.

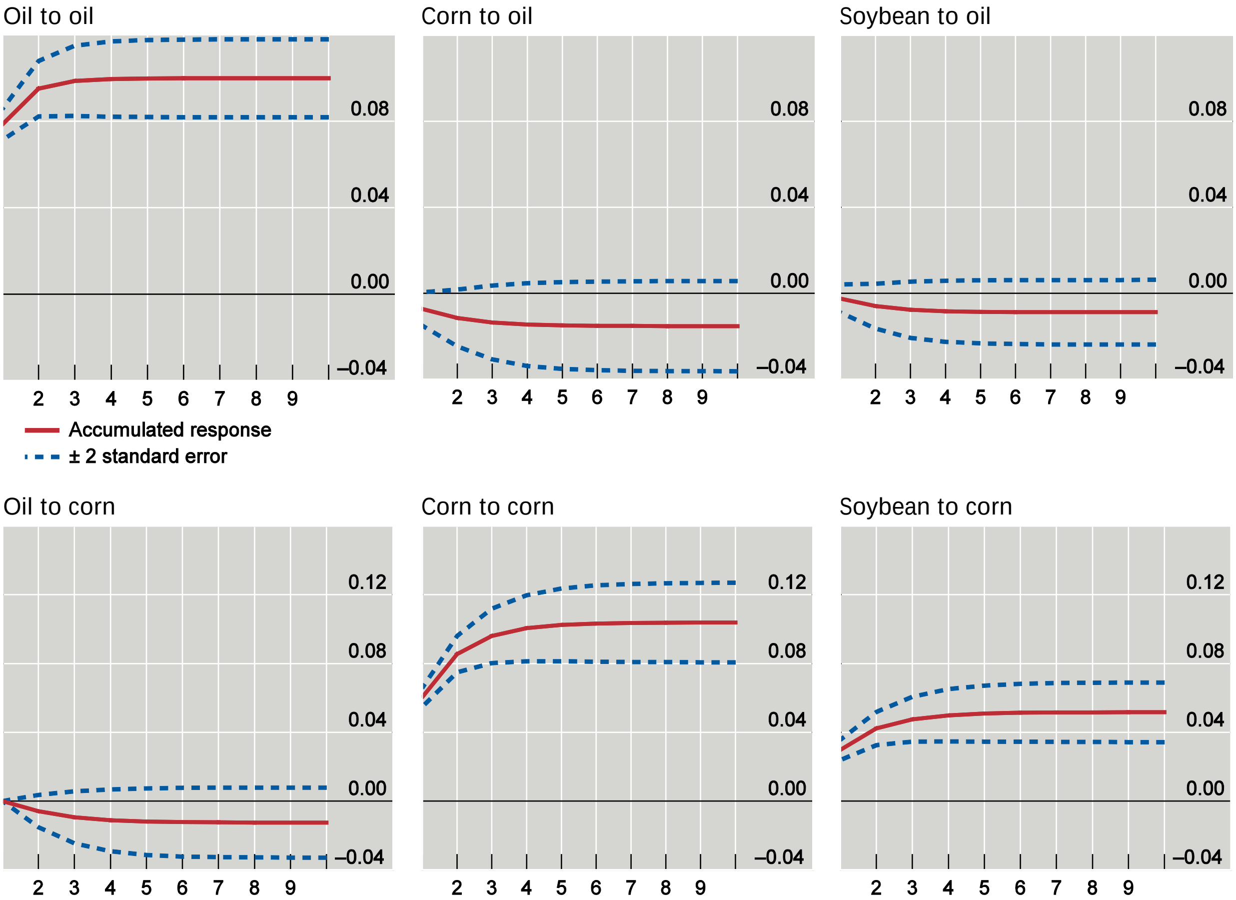
Sources: Datastream; BIS calculations.

Next I proceed to evaluate the price dynamics implied by the model, by computing impulse response functions. The elasticities to the fundamental shocks are identified through a Cholesky decomposition where the oil price is considered exogenous. The Cholesky ordering is completed by corn and soybeans. This ordering corresponds to our basic hypothesis: oil price shocks affect corn prices, which eventually impact on soybean prices. Moreover, corn price shocks affect soybean prices, but neither corn nor soybean price shocks affect oil prices. I will focus initially on the responses of prices to oil shocks. Alternatively, I will report also the responses to corn price shocks, as an alternative experiment suggested by the Granger-causality tests. Notice that this is a first step into a structural analysis, but we should be cautious in the interpretation of the results. This model is just too simple to implement a fully structural analysis, where specific shocks in global demand and monetary policy should be modelled explicitly, with actual commodity price shocks resulting as residuals. That is particularly relevant in the second part of the sample, where the magnitude of fundamental shocks to global demand and monetary policy were probably larger.

The results of a one standard deviation positive shock to oil prices are presented in Graph 7, upper row. I report cumulated impulse responses, to gauge the overall impact on price levels. The shock causes a permanent increase in oil prices, larger than the original shock, and clearly different from zero. The impact on crop prices is negative, with both the price of corn and soybeans decreasing permanently with respect to their pre-shock levels. However, the significance of this result is questionable, since zero lies inside the two standard error confidence interval around the point estimate of the impulse responses.

Accumulated impulse-responses[1] from estimated VAR model

Sample (adjusted): March 1986 to April 2006 Graph 7



[1] Prices (in logarithms), first difference. Accumulated response to Cholesky one standard deviation innovation. Cholesky ordering: oil, corn and soybean.

Sources: Datastream; BIS calculations.

Turning to the alternative experiment involving a one standard deviation positive corn price shock Graph 7 (lower row) shows that it had a significant permanent impact on soybean prices. That is consistent with our initial intuition about the effect of substitution and acreage competition effects on the soybean market. The final impact on corn prices is also positive, clearly different from zero and larger than the original shock and the response of soybean prices. The effect on oil prices is relatively small, and not very statistically significant, since the upper bound of the two standard error confidence interval overlaps with zero. Even allowing for a significant effect of corn price shocks on oil prices, this result indicates that the latter would be depressed by a positive corn price shock. In other words, before the enactment of the pre-Energy Independence Act, either corn and and oil prices had no relationship, or they behaved as complementary goods: corn price shocks somehow reduced the expenditure on oil to the extent of reducing its price, and viceversa.

In summary, before the Energy Independence Act of 2005, corn and soybean prices were stationary, whereas oil prices seemed to have a unit-root and had a relatively small and possibly negative impact in the long or short run price dynamics of the other two food commodities. In particular, oil price shocks exhibited no predictive causality over corn and soybean prices. However, as have been documented in other studies, corn prices did impact soybean prices in the short run.

Testing for a structural break: May 2006

The next step will be to test for the existence of a structural break on or around May 2006, when the new renewable fuel standard established by the Energy Independence Act of 2005 became applicable. I have not started to analyse the post-May 2006 data yet, but having characterised the stochastic processes prior to that date, I can still test whether both sub-samples can be properly described by the model that corresponds to the baseline period.

For this purpose I compute the usual Chow tests with the null hypothesis of parameter stability across sub-periods. I consider three versions of the Chow test: sample-split test (the most commonly used version), break-point test and forecast test. Candelon and Lütkepohl (2001) present a careful description of the first two tests, whereas Lütkepohl (2005) is a thorough reference for the forecast test. Sample-split and break-point tests are the most relevant for our application, since they only require that disturbances are white noise processes with an iid distribution and time invariant covariance matrix. The forecast test is reported for completeness, but it requires normality of the disturbances, which does not hold in this case. For the purpose of my test of structural break, it is important to remark that the break-point test is broader than the sample-split test, because the former allows for differences in the covariance matrix of the full sample model, whereas the latter assumes that this covariance matrix is constant across sub-samples. In other words, the sample-split test only allows for differences in the VAR equations coefficients.

Since Candelon and Lütkepohl (2001) show that these tests can be seriously distorted in small samples, I follow their suggestion of bootstrapping the residuals of the original VAR model estimation, and recomputing the tests for the bootstrapped system a large number of times (4000). I report both the asymptotic distribution and bootstrapped probability values of the tests.

The results are shown in Table 4. The break-point Chow test rejects stability of parameters, whereas the sample-split test cannot reject the null. Searching through nearby datapoints, I found that the sample-split test rejects the stability of the coefficients around the end of 2003 and beginning of 2004, ie just about the time when California and New York were banning the use of MTBE as a fuel additive. Therefore, I continue the analysis under the hypothesis that there was a structural break around this time, dating it in May 2006 for the purpose of the econometric analysis in the remainder of this paper.

Chow test for structural break[1]

Sample: March 1986 to April 2012[2] Table 4

	Test value	Bootstrapped p-value	Asymptotic p-value[3]	Degrees of freedom
Sample split Chow test	13.371	0.367	0.343	12
Break point Chow test	47.858	0.019	0.000	18
Chow forecast test	1.591	0.006	0.000	216, 708

[1] Tested break date: May 2006 (242 observations before break). [2] Included observations: 314. [3] For break point and sample split Chow tests, asymptotic Chi^2 p-value; for Chow forecast test, asymptotic F p-value.

Sources: Datastream; BIS calculations.

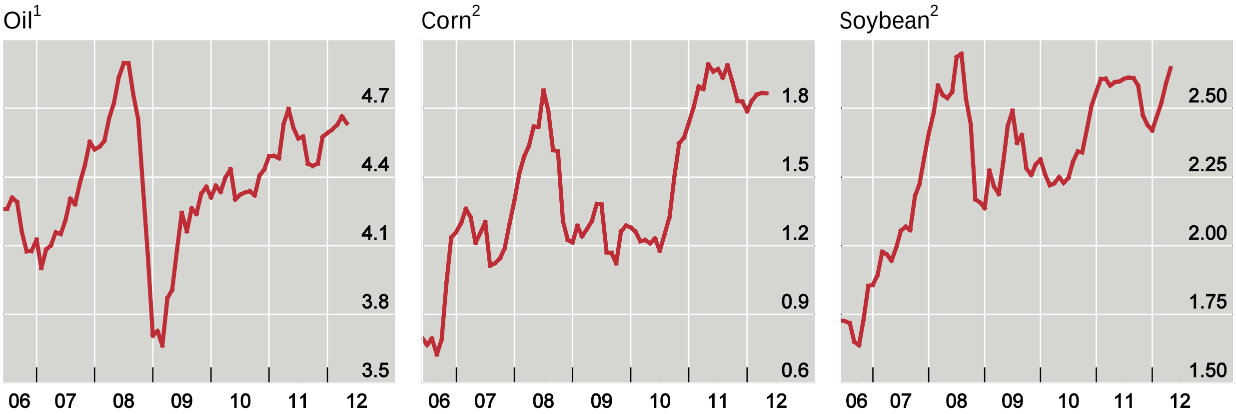
After the Energy Independence Act: May 2006 – April 2012

Graph 8 shows the path followed by the prices (in logs) of oil, corn and soybeans (respectively, from left to right). Aside from the steep plunge in all three commodity prices during 4Q08, the most remarkable feature is the positive drift in the prices of corn and soybeans, absent in the previous period. Standard unit-root tests confirm that it is no longer possible to reject the non-stationarity hypothesis for the log prices of corn and soybeans.[12] However, augmented-Dickey Fuller tests reject non-stationarity for oil prices. Presumably, the sharp drop in oil prices during the 4Q08 could be creating autocorrelation or heteroskedasticity problems that might be reducing the efficiency of the estimation. Examining the residuals of the augmented-Dickey Fuller regression, I find no evidence of autocorrelation. But White´s general test rejects the homoskedastic null hypothesis (Appendix Table A3). Therefore, I focus on the Phillips-Perron test results, which do not reject the unit root null for log oil prices (Appendix Table A2).

[12] In the rest of this section, I will be using a small sample: 72 months, or 6 full years of data. However, the small sample properties of most statistics used in the tests that follows are not fully understood. Therefore, caution is of paramount importance when interpreting the results.

Oil, corn and soybean prices, May 2006 to April 2012

In logarithms Graph 8



[1] In US dollars per barrel. [2] In US dollars per bushel.

Sources: Datastream; BIS calculations.

The next step involves testing for the existence of cointegration among the three price series. I start with bilateral Johansen tests (Appendix Table A4), which find that the log prices of oil and corn were cointegrated. Moreover, the tests also reject cointegration of the price pairs of oil-soybeans (at the 10% significance level), and corn-soybeans. The results support two important changes in the price dynamics of corn (and soybean) prices in this sub-sample: both crop prices are no longer stationary, and corn prices seem to be cointegrated with oil prices. In other words, there is now a long-run relationship between these prices that links them together in a stable fashion, which was not identified before May 2006.[13]

With cointegration established, I proceed to test whether it holds for the three price series pooled together (as opposed to conducting 3 separate tests), and based on the findings of the bilateral tests, I impose the constrain that the coefficient of soybean prices in the cointegrating vector is equal to zero. Once again, the data support the existence of a single cointegrating vector, and the restriction on the soybean price coefficient in that vector cannot be rejected (see Appendix Table A4).

Next I investigate the short run price dynamics of the three commodities by the estimation of a vector error correction model (VECM). Again it is a very simple model only involving the cointegrating vector and lags of the endogenous variables. A Wald lag exclusion test determines an optimal 2-lag structure for the model.

[13] I carried out cointegration tests for these variables during the baseline period, and it was rejected for all possible specifications. The results are available upon request.

VEC model estimation results[1]

Sample (adjusted): August 2006 to April 2012[2] Table 5

Estimated parameter [t-statistic]			
Cointegrating equation 1[3]			
Oil (–1)	Corn (–1)	Soybean (–1)	Constant
1.000	–0.551	0.000	–3.566
	[–4.209]		

Error correction model	Oil	Corn	Soybean
Cointegrating equation 1	**–0.200	0.014	–0.069
	[–3.943]	[0.233]	[–1.532]
Oil (–1)	**0.238	–0.147	–0.131
	[2.077]	[–1.082]	[–1.297]
Oil (–2)	**0.342	0.090	**0.222
	[2.873]	[0.637]	[2.104]
Corn (–1)	**0.310	*0.328	0.184
	[2.119]	[1.894]	[1.422]
Corn (–2)	–0.166	0.120	0.042
	[–1.113]	[0.681]	[0.322]
Soybean (–1)	–0.154	–0.053	0.212
	[–0.763]	[–0.221]	[1.187]
Soybean (–2)	**0.442	–0.008	–0.113
	[2.193]	[–0.035]	[–0.634]
Constant	–0.004	0.010	0.008
	[–0.420]	[0.852]	[0.962]
R-squared	0.476	0.113	0.205
Adjusted R-squared	0.416	0.011	0.114

** = significant at 5%; * = significant at 10%.

[1] Estimated over prices (in logarithms), first difference unless otherwise indicated. [2] Included observations (after adjustments): 69. [3] Estimated over prices (in logarithms).

Sources: Datastream; BIS calculations.

Table 5 presents the results of estimating the VECM. Inspection of the table reveals that oil prices have a strong autoregressive structure, and once again innovations in corn and soybean prices seem to have a significant impact on oil prices in the short run, contrary to the usual conventional wisdom. The loadings of the cointegration equation show that it only affects significantly oil prices, suggesting that when there are deviations from the long term relationship between corn and oil prices, it is oil prices which adjust to the level of corn prices to preserve the long-term relationship. The adjustment is relatively slow, as deviations are erased in about a year. Once again, this is counter-intuitive: in most policy discussion about this topic, the concern is about the swings in food prices causes by oil prices changes. This result shows that in the long-run the adjustment seem to flow from corn to oil markets.[14]

The Granger-causality tests shown in Table 6 also bring some unexpected results: first, they strongly reject that oil prices individually Granger-cause corn prices. In fact, data lends more support to the hypothesis that soybean prices are Granger-caused by oil prices, which cannot be rejected at a 10% significance level. Moreover, joint Granger-causality of oil prices from corn and soybean prices is established at the standard 5% significance level.[15] All this point to interesting connections between oil prices and these food staple prices, but quite different from those anticipated by the usual discussion about the potential impact of biofuel promotion policies.

VEC model Granger causality tests[1]

Sample: May 2006 to April 2012[2] Table 6

Excluded variable	Chi-sqared	Degrees of freedom	P-value
	Dependent variable		
	Oil		
Corn	5.116	2	0.078
Soybean	4.945	2	0.084
Both	9.718	4	0.046
	Corn		
Oil	1.276	2	0.528
Soybean	0.055	2	0.973
Both	1.644	4	0.801
	Soybean		
Oil	4.901	2	0.086
Corn	2.327	2	0.312
Both	6.994	4	0.136

[1] Tested over prices (in logarithms), first difference. [2] Included observations: 69.

Sources: Datastream; BIS calculations.

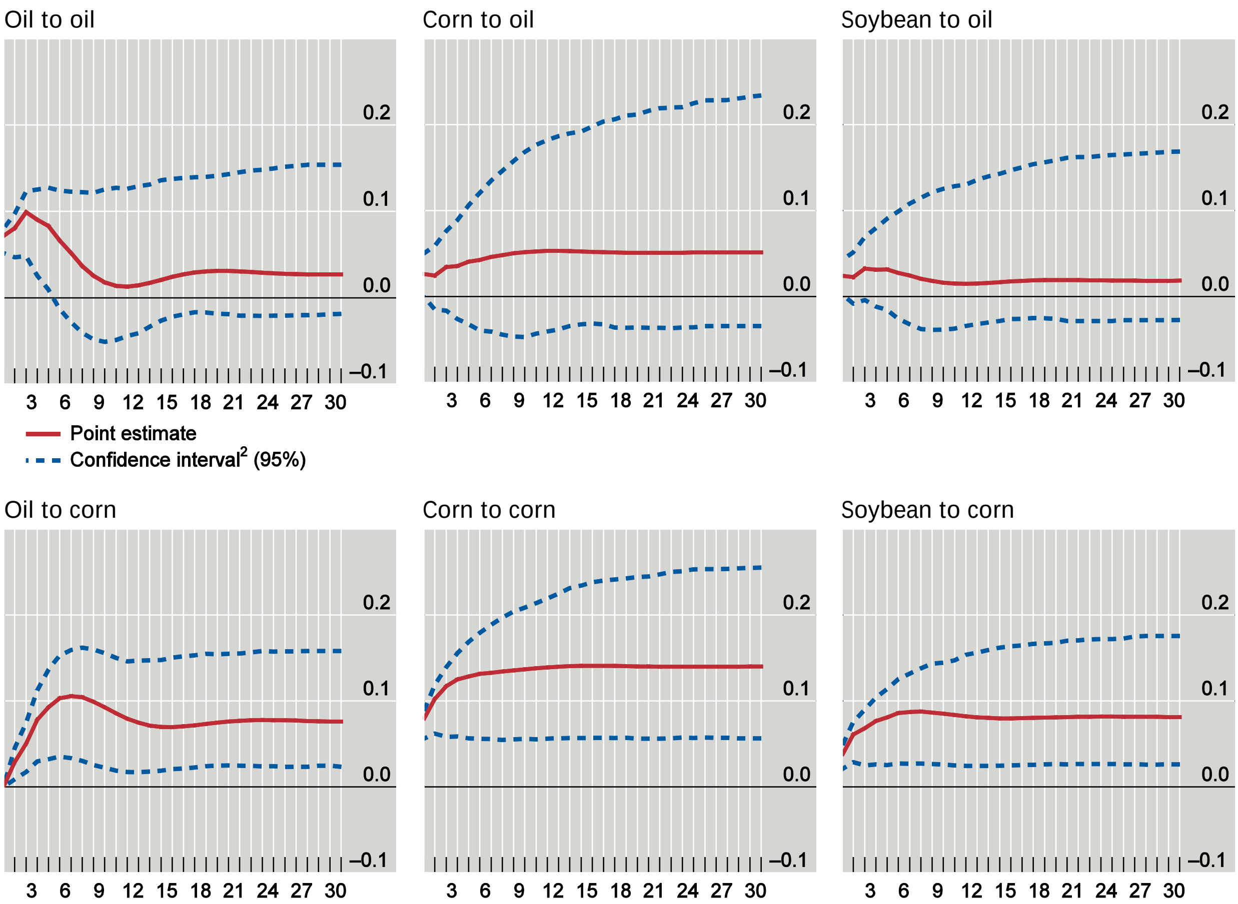
[14] In this short sample, I suspect that this can be an artifact of the sharp drop in oil prices during 4Q08 (relatively deeper than the drop in corn or soybean prices) which in practice caused a drop in the cointegration equation. Since oil prices had a steeper recovery afterwards, this period might be biasing the outcome.

[15] Individually, I reject the hypotheses that either corn or soybean prices do not Granger-cause oil prices at 10% significance.

Finally, Graph 9 displays the impulse responses implied by the estimated model to a one standard deviation shock in oil and corn prices. The upper row of the graph shows the response to an oil price shock. The response of corn prices is not very precisely estimated, but it seems significant and builds up over a period of a year. Moreover, the impact of oil price shocks on corn prices is now positive, contrary to the finding during the baseline period. The effect on soybean prices is also positive but more muted. In both cases, it is not possible to rule out the possibility that the response is trivial, since the zero response is within the 95% confidence interval. The bottom row presents impulse responses to a corn price shock. It now has a clearly significant impact in both oil and soybean prices. As in the previous case, the effect on oil prices is positive, implying that oil and corn behaved as substitutive goods in this sub-period, probably because energy can be produced now by either burning oil, or corn (as ethanol).

Impulse-responses[1] from estimated VEC model

Sample (adjusted): August 2006 to April 2012 Graph 9



[1] Prices (in logarithms). Response to Cholesky one standard deviation innovation. Cholesky ordering: oil, corn and soybean. [2] Efron's percentile method.

Sources: Datastream; BIS calculations.

In summary, after the Energy Independence Act became effective in May 2006, there have been significant changes in the properties and price dynamics of corn and soybeans. They behave as non-stationary time series, and corn prices seem to be cointegrated with oil prices. There are some unexpected results: the cointegration vector only affect oil prices (implying adjustment of oil price levels towards corn price levels), and the evidence supports Granger-causality flowing from crop prices to oil prices. As for price dynamics, oil price shocks now have a positive and significant effect on corn prices, and viceversa. As usual, soybean prices respond mutely to oil price shocks and strongly to corn price shocks.

5. Conclusions

High commodity prices are a relevant feature of the current macroeconomic landscape. They are a cause of concern for both advanced and emerging economies, for their potential impact on inflation, even in the face of the sluggish pace of global economic growth. They have arguably helped to anchor the external and fiscal positions of several Latin American countries. Even though there is certain consensus on the macroeconomic and financial factors (financialisation of commodity markets, loose monetary policy, and high global demand, especially from China and India) that should play a relevant role in explaining their behaviour, empirical tests are inconclusive. This paper documents the change observed in the price dynamics of some food staples (corn and soybeans) since 2006. Since then, both crop prices are related to oil prices in the long run, and in the short run are more affected by oil price shocks. The economic explanation proposed relates these changes to the new renewable fuel standard (RFS) imposed by the Energy Independence Act of 2005. Additional questions remain. For instance, it is surprising that both in the long and short term, impulses seem to flow from crop prices to oil prices, which is counter-intuitive. It is also suggestive that oil price shocks had a negative impact on crop prices before May 2006. One possible interpretation is that all three prices are being driven by another factor, not included in this analysis, and which had a lesser or different influence in crop prices before the Energy Independence Act of 2005. But then the next question is what that factor or factors those might be. Answering these questions will require further research involving a structural approach that, while preserving the joint analysis of oil and crop prices, would incorporate explicit measures of macroeconomic variables, including possibly global demand, monetary policy and financialisation of commodity prices.

Appendix

Unit root tests

Sample: January 1986 to April 2006 — Table A1

Commodity	Test	Variable	Specification: Model	Specification: Lag length[1]	t-statistic	P-value
Oil	ADF	log price	intercept	4	–0.523	0.883
			intercept and trend	4	–1.774	0.714
		log price, first difference	intercept	3	–8.238	0.000
			intercept and trend	3	–8.307	0.000
	PP	log price	intercept	4	–0.753	0.830
			intercept and trend	2	–2.450	0.353
		log price, first difference	intercept	11	–13.075	0.000
			intercept and trend	11	–13.085	0.000
Corn	ADF	log price	intercept	1	–3.927	0.002
			intercept and trend	1	–3.919	0.013
		log price, first difference	intercept	4	–7.438	0.000
			intercept and trend	4	–7.419	0.000
	PP	log price	intercept	1	–2.905	0.046
			intercept and trend	1	–2.901	0.164
		log price, first difference	intercept	12	–9.013	0.000
			intercept and trend	12	–8.986	0.000
Soybean	ADF	log price	intercept	1	–3.261	0.018
			intercept and trend	1	–3.283	0.072
		log price, first difference	intercept	0	–11.174	0.000
			intercept and trend	0	–11.156	0.000
	PP	log price	intercept	4	–2.891	0.048
			intercept and trend	4	–2.911	0.161
		log price, first difference	intercept	3	–11.096	0.000
			intercept and trend	3	–11.077	0.000

ADF = Augmented Dickey-Fuller; PP = Phillips-Perron.

[1] For ADF tests, lag length determined with Akaike information criterion; for PP tests, lag length determined with Newey-West bandwidth criterion, using Bartlett kernel as spectral estimation method.

Sources: Datastream; BIS calculations.

Unit root tests

Sample: May 2006 to April 2012 Table A2

Commodity	Test	Variable	Specification: Model	Specification: Lag length[1]	t-statistic	P-value
Oil	ADF	log price	intercept	2	–3.421	0.014
			intercept and trend	2	–3.801	0.022
		log price, first difference	intercept	5	–4.721	0.000
			intercept	5	–4.677	0.002
	PP	log price	intercept and trend	4	–2.235	0.196
			intercept	4	–2.466	0.344
		log price, first difference	intercept and trend	2	–5.509	0.000
			intercept	2	–5.475	0.000
Corn	ADF	log price	intercept and trend	2	–2.195	0.210
			intercept	2	–2.697	0.241
		log price, first difference	intercept and trend	1	–4.395	0.001
			intercept	1	–4.383	0.004
	PP	log price	intercept and trend	4	–1.824	0.366
			intercept	4	–2.227	0.467
		log price, first difference	intercept and trend	3	–6.290	0.000
			intercept	2	–6.219	0.000
Soybean	ADF	log price	intercept and trend	8	–1.940	0.312
			intercept	8	–2.187	0.488
		log price, first difference	intercept and trend	7	–3.771	0.005
			intercept	7	–3.761	0.025
	PP	log price	intercept and trend	1	–1.774	0.391
			intercept	2	–2.109	0.532
		log price, first difference	intercept and trend	4	–5.878	0.000
			intercept	4	–5.846	0.000

ADF = Augmented Dickey-Fuller; PP = Phillips-Perron.

[1] For ADF tests, lag length determined with Akaike information criterion; for PP tests, lag length determined with Newey-West bandwidth criterion, using Bartlett kernel as spectral estimation method.

Sources: Datastream; BIS calculations.

White heteroskedasticity test

Sample: May 2006 to April 2012 — Table A3

Statistic	Value	Degrees of freedom	Probability	
F-statistic	3.691	F(9,59)	0.001	
Observations*R-squared	24.854	Chi-Square(9)	0.003	
Scaled explained sum of squares	23.990	Chi-Square(9)	0.004	
		Test Equation[1]		
Variable[2]	Coefficient	Standard error	t-statistic	Probability
constant	0.155	0.240	0.646	0.521
oil(–1)	–0.068	0.110	–0.616	0.541
oil(–1)^2	0.008	0.013	0.609	0.545
oil(–1)*diff[oil(–1)]	–0.104	0.048	–2.164	0.035
oil(–1)*diff[oil(–2)]	–0.040	0.048	–0.838	0.406
diff[oil(–1)]	0.453	0.205	2.208	0.031
diff[oil(–1)]^2	–0.066	0.100	–0.659	0.513
diff[oil(–1)]*diff[oil(–2)]	0.562	0.157	3.590	0.001
diff[oil(–2)]	0.174	0.202	0.863	0.392
diff[oil(–2)]^2	–0.027	0.099	–0.269	0.789
R-squared	0.360			
Adjusted R-squared	0.263			

For a given variable ‘X’, diff[X] denotes first difference; X^2 denotes square.

[1] Dependent variable: residuals^2. Sample: August 2006 to April 2012; included observations: 69. [2] Prices (in logarithms).

Sources: Datastream; BIS calculations.

Cointegration rank tests

Sample: May 2006 to April 2012 Table A4

Variables[1]	Null: number of cointegrating equations	Trace statistic (probability)	Maximum Eigen-value (probability)	LR statistic (probability)
	Unrestricted tests			
Oil-Corn[2]	0	25.767	20.122	
		(0.008)	(0.010)	
	1	5.644	5.644	
		(0.220)	(0.220)	
Oil-Soybean[2]	0	22.327	14.768	
		(0.026)	(0.074)	
	1	7.559	7.559	
		(0.100)	(0.100)	
Corn-Soybean[2]	0	10.791	6.232	
		(0.562)	(0.762)	
	1	4.559	4.559	
		(0.335)	(0.335)	
Oil-Soybean-Corn[3]	0	30.791	20.005	
		(0.038)	(0.071)	
	1	10.786	6.134	
		(0.225)	(0.596)	
	Restricted test: oil coefficient = 1; soybean coefficient = 0			
Oil-Soybean-Corn[3]	1			0.067
				(0.795)

[1] Prices (in logarithms). [2] No deterministic trend assumption (restricted constant). [3] Linear deterministic trend assumption.
Sources: Datastream; BIS calculations.

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