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Joscha Beckmann and Robert Czudaj[1]

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Abstract

Although the link between oil prices and dollar exchange rates has been frequently analyzed, a clear distinction between prices and nominal exchange rate dynamics and a clarification of the issue of causality has not been provided. In addition, previous studies have mostly neglected nonlinearities which for example may stem from exogenous oil price shocks. Using monthly data for various oil-exporting and oil-importing countries, this study contributes to the clarification of those issues. We discriminate between long-run and time-varying short-run dynamics, using a Markov-switching vector error correction model. In terms of causality, the results differ between the economies under observation but suggest that the most important causality runs from exchange rates to oil prices, with a depreciation of the dollar triggering an increase in oil prices. On the other hand, changes in nominal oil prices are responsible for ambiguous real exchange rate effects mostly through the price differential and partly also through a direct influence on the nominal exchange rate. Overall, the fact that the adjustment pattern frequently differs between regimes underlines the fact that the relationships are subject to changes over time, suggesting that nonlinearities are an important issue when analyzing oil prices and exchange rates.

JEL Classification: C32, E31, F31, G15

Keywords: Bayesian econometrics; cointegration; exchange rates; Markov-switching model; oil prices; oil-importing and oil-exporting countries

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1 Both University of Duisburg-Essen. – Thanks for valuable comments on previous versions of the paper are due to the participants of the 10th INFINITI Conference on International Finance, Trinity College Dublin/Ireland, the 14th INFER Annual Conference, University of Coimbra/Portugal, the 87th WEAI Annual Conference, San Francisco/California, the 44th Annual Money, Macro and Finance Conference, Trinity College Dublin/Ireland, the 11th EEFS Annual Conference, Koc University, Istanbul/Turkey, the 14th Workshop on International Economics, University of Göttingen/Germany and the 7th International Finance Conference, ISC Paris School of Management/France. The current version of this paper will be presented at the 28th Annual Congress of the European Economic Association, University of Gothenburg/Sweden and at the Annual Congress of the German Economic Association (Verein für Socialpolitik), University of Düsseldorf/Germany. – All correspondence to Joscha Beckmann, University of Duisburg-Essen, Department of Economics, Universitätsstr. 12, 45117 Essen, Germany, E-Mail: joscha.beckmann@uni-due.de.

1 Introduction

Since oil prices and exchange rates against the dollar both experienced long swings after the breakdown of Bretton Woods, the link between both quantities has attracted considerable interest from both policymakers and researchers.[1] In a nutshell, empirical research has provided evidence for two directions of causalities: A popular finding is that real exchange rates and real oil prices are cointegrated over the recent floating period. In this vein, shocks to real oil prices have been identified as a possible explanation for the non-stationarity of real exchange rates [Chaudhuri and Daniel, 1998]. In terms of dollar exchange rates, other studies have provided evidence for a real appreciation of the dollar in the case of an increase in real oil prices and, in addition, for an increase in the nominal price of oil as a result of a depreciation of the dollar [Yousefi and Wirjanto, 2004].

However, although different relationships between oil prices and exchange rates have been identified, some important issues remain to be solved. Firstly, the question of causalities between oil prices and exchange rates has not been broadly examined in an unrestrictive fashion. This is due to the fact that many studies which provide evidence for one direction of causality have focused either on particular bilateral exchange rates or on effective exchange rate dynamics. A second caveat of previous studies stems from the possibility of nonlinearities in the relationship between exchange rates and oil prices. Since different degrees of volatility and co-movements between oil prices and exchange rates vis-à-vis the dollar can be identified, it seems important to allow for structural changes when analyzing the corresponding long-run relationships [Breitenfellner and Crespo Cuaresma, 2008; Reboredo, 2012]. Nonlinearities may also stem from major oil prices shocks triggered by exogenous factors.[2] Finally, surprisingly little attention has been paid to the dissection of nominal price and exchange rate effects when examining causalities between oil prices and exchange rates. For example, many studies, such as the recent work provided by Chen and Chen [2007] which conducts panel cointegration methods, have directly analyzed the relationship between real exchange rates and the real oil price. However, as known from the literature on purchasing power parity, such an approach applies symmetry restrictions with regard to prices and nominal exchange rates without putting the underlying dynamics under closer scrutiny [Sarno and Valente, 2006]. Moreover, due to a diverging degree of importance of a commodity such as oil for oil-exporting and oil-importing countries it makes much more sense analyzing the causality pattern for each economy separately instead of applying a panel approach which produces results based on an average of all individual economies.

Adopting a multivariate Markov-switching vector error correction model (MS-VECM) framework, this paper contributes to the literature to tackle those caveats. Firstly, our approach is unrestricted in the sense that it allows an analysis of the issue of causality between oil prices and exchange rate dynamics while distinguishing between short- and long-run dynamics. Secondly, we allow for time-

[1] For instance, Lammerding, Stephan, Trede and Wilfling [2013] recently provided evidence for the existence of speculative bubbles in oil price dynamics.

[2] The common view is that major oil price shocks are triggered by exogenous factors. However, according to Kilian [2008], this view is supported by the data for the 1980/81 and 1990/91 oil price shocks, but not necessarily for shocks occurring afterwards.

varying causality patterns in terms of an error correction mechanism regarding deviations from a long-run equilibrium. This is a huge improvement compared with studies which rely on a linear specification. Finally, our framework also enables us to dissect the contribution of prices and nominal exchange rate dynamics to real exchange rate movements while investigating the causality on oil prices. In addition, our analysis, which is based on an evaluation of twelve different exchange rates against the US dollar, also allows us to answer the question of whether a general pattern for oil-exporting or oil-importing countries can be observed or if the relative contribution of an oil price or exchange rate shock depends on the level of oil importance to the production sector of the national economy and the net position of the economy in the oil market. In a broader context, using a structural vector autoregression (SVAR) the study of Wang, Wu and Yang [2013] recently showed that the impact on stock markets to oil price shocks highly depends on whether the country is a net importer or exporter in the world oil market.

The remainder of this paper is organized as follows. The following section provides a brief description of theoretical considerations and summarizes previous empirical findings providing evidence for two ways of causlity. Section 3 describes our data and provides a description of and motivation for our empirical framework. Our results are presented and analyzed in Section 4. Section 5 concludes.

2 Theoretical considerations and review of the literature

2.1 Causalities running from oil prices to exchange rates

Various theoretical relationships between oil prices and exchange rates have been established in the literature, with causalities going in both directions. Two transmission channels of oil prices to exchange rates can be roughly distinguished: the 'terms of trade channel' and the 'wealth effect channel'. The 'terms of trade channel' focuses on real exchange rates and was originally introduced by Amano and Van Norden [1998a,b]. Their basic approach may be illustrated as a simple two-economy model consisting of two sectors which produce tradable and nontradable goods, respectively:

$$p_t = \alpha p_t^T + (1-\alpha) p_t^N, \tag{1}$$

$$p_t^* = \alpha^* p_t^{T^*} + (1-\alpha^*) p_t^{N^*}, \tag{2}$$

where p_t^T and p_t^N correspond to the logarithm of prices for tradable and nontradable goods, respectively. p_t indicates the logarithmic general price level and the foreign economy is always denoted by an asterisk. The weights α and α^* give the corresponding expenditure shares on tradable goods [Chen and Chen, 2007]. Oil enters both production functions as an input factor while the price of nontradable goods is determined solely by labor costs.

In the original model, the real exchange rate is expressed in internal terms as the ratio between prices of tradable and nontradable goods in a small open economy with a relative increase in the price of tradable goods corresponding to a real depreciation of the domestic currency. Assuming that the tradable output price is fixed internationally, one can see that the price for nontradable goods determines the reaction of the real exchange rate [Bénassy-Quéré, Mignon and Penot, 2007]. If the nontradable sector is more dependent on oil than the tradable sector, the price of nontradable goods increases to a greater extent and the domestic currency experiences a real appreciation once the oil price increases. Expressed the other way around, a rise in oil prices results in a real depreciation in economies with large oil dependence in the tradable sector [Chen and Chen, 2007].[3] This channel of transmission operates for oil-exporting economies, whereas movements in oil prices, by definition, dominate the terms of trade of industrialized economies [Backus and Crucini, 2000; Buetzer, Habib and Stracca, 2012].

Turning to the more popular definition of the real exchange rate, in external terms the logarithm of the real exchange rate q_t may be expressed as follows:

$$q_t = s_t + p_t^* - p_t, \tag{3}$$

where s_t corresponds to the logarithm of the nominal exchange rate. Substituting the price levels based on Equations (1) and (2) in Equation (3) gives

$$q_t = (s_t + p_t^{T^*} - p_t^T) + (1-\alpha)(p_t^T - p_t^N) - (1-\alpha^*)(p_t^{T^*} - p_t^{N^*}). \tag{4}$$

According to Equation (4), the dynamics stemming from an increase in oil prices become more complicated. Strictly speaking, a country with a greater increase in inflation experiences a real appreciation, a mechanism which mirrors the Balassa-Samuelson effect [Buetzer *et al.*, 2012]. Under the assumption that the price of a tradable good is internationally fixed, and that $\alpha = \alpha^*$, the reaction of the real exchange rate is then again determined by the relative oil-dependence of the tradable and nontradable sectors [Chen and Chen, 2007].

The 'terms of trade channel' therefore works mainly through relative prices and may leave the nominal exchange rate unchanged. In Equation (4), purchasing power parity (PPP) is assumed to hold only for the price of tradable goods, with a relative rise in the price differential of tradable goods being matched by a proportional depreciation of the nominal exchange rate. However, the literature on the validity of PPP is not only extensive but controversial (see Sarno and Taylor [2002] for an overview). Considering recent empirical results, which have delivered evidence in favor of a nonlinear PPP adjustment of nominal and real exchange rates based on exponential smooth transition regression (ESTR) models [Taylor, Peel and Sarno, 2001; Kilian and Taylor, 2003; Wu and Hu, 2009], a reasonable view would be that the price differential between two countries is important for the long-run path of the nominal exchange rate, although the relationship is not necessarily strictly proportional.[4] Hence, oil price shocks might also introduce changes in nominal exchange rates in the long-run, which are not proportional to the price differential. In our framework, consumer prices are used without distinguishing between tradable and nontradable goods. For reasons mentioned above, one can also expect

[3] See Schnabl and Baur [2002] for an analysis of the importance of export prices to the yen-dollar exchange rate.

[4] This view also corresponds to a weak version of PPP introduced by Cheung and Lai [1998] and requires only that a linear combination of exchange rates and prices is found to be stationary. From a theoretical point of view, nonlinear real exchange rate dynamics can be formally derived in the context of international arbitrage costs [Taylor *et al.*, 2001].

the nominal exchange rate to react to changes in the price differential without expecting the real exchange rate remaining constant. To sum up, the overall effect on the real exchange rate depends on the nominal exchange rate response relative to the first-round impact of the rise in the relative price level. In terms of exchange rates against the US dollar, one would expect currencies of countries with large oil-dependence relative to the US to depreciate in real terms.

On the other hand, transmission through the 'wealth effect channel' emphasizes effects on the nominal exchange rate and focuses on the impact of oil price changes on international portfolio decisions and trade balances. Such frameworks have been provided, for instance, by Krugman [1983] and Golub [1983], who adopt a three-country framework. According to this view, oil-exporting countries experience a wealth transfer if the oil price rises [Bénassy-Quéré *et al.*, 2007]. In general, the effects on exchange rates depend on the portfolio choices of oil-importing and oil-exporting countries. Assuming that oil-exporting countries reinvest their revenues in US dollar assets, the dollar will appreciate in the short-run. However, the long-run reaction of the US dollar against other currencies is less clear-cut and determined by the weight of oil in US total imports compared to the US weights in OPEC imports [Bénassy-Quéré *et al.*, 2007; Coudert, Mignon and Penot, 2008]. Summing up the theoretical record, positive oil price shocks lead to a real appreciation (depreciation) of the exchange rates of oil-exporting (oil-importing) economies [Buetzer *et al.*, 2012].

As a next step, we turn to empirical findings referring to a causality running from oil prices to exchange rates. The relationship between the real oil price and real exchange rates against the US dollar has been analyzed for several countries in various studies covering diverse spans of data. Applying cointegration techniques, many authors have provided evidence of a real effective appreciation of the US dollar in the case of rising oil prices in the long-run [Amano and Van Norden, 1998b; Bénassy-Quéré *et al.*, 2007; Coudert *et al.*, 2008]. Clostermann and Schnatz [2000] focus on the real exchange rate of the dollar against the euro and also find indications of a real appreciation of the dollar in the case of a rise in real oil prices. However, evidence for the effect on the real effective exchange rates of other countries is less clear-cut. Habib and Kalamova [2007] do not find a long-run relationship between real effective exchange rates and oil prices for Norway and Saudi Arabia, but report evidence for a long-run real appreciation for Russia if oil prices rise. In an earlier study, Rautava [2004] also finds that the Russian economy is affected significantly by fluctuations in oil prices and the real exchange rate through both long-run equilibrium conditions as well as short-run impacts. Narayan, Narayan and Prasad [2008] provide evidence in favor of a rise in oil prices leading to an appreciation of the Fijian dollar vis-à-vis the US dollar while adopting a GARCH framework. Using annual data from 1975 to 2008, Al-Mulali [2010] surveys this relationship for Norway and finds evidence for a real effective appreciation in the case of rising oil prices. From a broader perspective, Lizardo and Mollick [2010] embed the real oil price in a simple form of the monetary model of exchange rate determination and find that an increase in the real oil price leads to a nominal depreciation of the dollar against net oil importers, while the currencies of oil importers depreciate against the dollar. Chen and Chen [2007] examine a panel of G-7 countries and find that real oil prices have significantly contributed to real exchange rate movements using panel cointegration techniques. Based on the convenience yield Beutler [2012] shows that commodity prices have forecasting power with regard to 'commodity currencies' (i.e. currencies of commodity exporters). Finally, the most recent study was provided by Buetzer *et al.* [2012], who identify different shocks to real oil prices in a SVAR and find no evidence that the exchange rates of oil exporters systematically appreciate against those of oil importers.

2.2 Causalities running from (US dollar) exchange rates to oil prices

From a general point of view, a causality running from exchange rates to commodity prices can be derived using an asset-pricing approach of exchange rate determination which links the present exchange rate to the discounted sum of futures fundamentals. However, owing to the fact that fundamentals and exchange rates are jointly determined in equilibrium, convincing empirical support for this view has not yet been delivered [Chen, Rogoff and Rossi, 2008].

A direct transmission from US dollar exchange rates to oil prices through changes in supply and demand stems from the exceptional role of the international dollar as a settlement currency. Abstracting from transaction costs, consider the following relationship between the logarithms of the oil price denominated in US dollar (o_t) and another currency (o_t^*) based on the law of one price:

$$o_t^* = s_t - o_t, \tag{5}$$

where s_t again indicates the logarithm of the nominal exchange rate of a domestic currency against the US dollar (domestic currency per one unit of US dollar). If a commodity such as oil is denominated in US dollar, a domestic appreciation against the dollar lowers the price of oil measured in terms of the domestic currency, which increases demand and may result in a general rise in oil prices [Akram, 2009].[5] In the following, this transmission channel is labeled as the 'denomination channel'. The impact on the supply side is not clear-cut. Positive effects may stem from an exchange rate-driven rise in the price of oil on drilling activities and production capacities, although the latter causality has changed over time. On the other hand, a depreciation of the domestic currency may reduce purchasing power and shift resources away from oil production, which would result in a decrease in supply [Coudert *et al.*, 2008]. Although a direct distinction of this causality from the 'denomination channel' is not possible in our empirical framework, we will refer to this mechanism as the 'adjustment channel'.

In the following, we turn to empirical findings referring to a causality running from exchange rates to oil prices. Providing evidence for such causality, Cheng [2008] identifies an increase in the real (nominal) oil price as a response to a real (nominal) effective US dollar appreciation. Other studies also conclude that the causality runs mainly from dollar exchange rates to oil prices. Yousefi and

[5] Similar to PPP, empirical studies have delivered mixed evidence for the law of one price, also referring to nonlinear adjustment processes as a possible explanation. For the law of one price, the theoretical concepts of bands of transaction and iceberg costs are related to threshold autoregressive (TAR) models, where the transition from one regime to another is discrete once the threshold is reached. See Sarno, Taylor and Chowdhury [2004] for a survey of empirical studies related to this issue.

Wirjanto [2004] analyze five OPEC countries and provide evidence that crude oil export prices respond positively to depreciations against the dollar for the purpose of stabilizing export revenues. Focusing on nominal effective US dollar exchange rates, Krichene [2005, 2006] concludes that an appreciation of the nominal effective dollar exchange rate may lead to both an increase and a decrease in oil prices. Zhang, Fan, Tsai and Wei [2008] detect that the US dollar drives international crude oil prices by using standard cointegration, VAR, and ARCH type models for daily data running from 4 January, 2000 to 31 May, 2005. Beckmann and Czudaj [2013b] apply two different measures of effective US dollar exchange rates and find evidence for a causality pattern mainly running from nominal exchange rates to nominal oil prices by using cointegration techniques for a sample period from January 1974 to November 2011. They also incorporate the three-month treasury bill rate in their study to account for dynamics stemming from monetary policy decisions. With respect to the general link between exchange rates and commodity prices, Chen *et al.* [2008] find robust power of commodity currencies in predicting global commodity prices, while their results provide little evidence for exchange rate predictability based on commodity prices. In a recent study, Ferraro, Rogoff and Rossi [2012] investigate the predictive content of the oil price for future exchange rate developments and only find a robust explanatory power for the short-run. However, other studies frequently find that commodity prices are weakly exogenous with respect to the exchange rate, a result which may mirror the fact that commodities are priced in competitive world markets [Buetzer *et al.*, 2012].

2.3 Nonlinearities and oil-exchange rate causalities

Although empirical research has provided evidence for different kinds of nonlinearity in analyzing exchange rates, this issue has mostly been neglected when putting the relationship between exchange rates and oil prices under closer scrutiny. An exception is the study by Akram [2004], which supports the view of a nonlinear relationship between the value of the Norwegian krona and oil prices. In addition, Wang and Wu [2012] conduct linear as well as nonlinear Granger causality tests to put the causality between energy prices and US dollar exchange rates under closer scrutiny. Their results indicate that until the beginning of the financial crisis a linear causality running from petroleum prices to exchange rates and a nonlinear causality running from exchange rates to natural gas prices can be observed. For the period after the financial crisis, they find evidence for a nonlinear bidirectional causal relationship between petroleum prices and exchange rates and no causality between exchange rates and natural gas prices. They attribute nonlinear causality to a volatility spillover and a regime shift. Furthermore, Tiwari, Dar and Bhanja [2013] uncover linear and nonlinear causal relationships between the oil price and the real effective exchange rate of Indian rupee at higher time scales (lower frequency) using a wavelet based analysis, but they do not find any causality at the lower time scales.[6]

However, besides Beckmann and Czudaj [2013b], who focus on effective exchange rates rather then bilateral country pairs, most studies mentioned in the previous subsections base their analysis on a

[6]Moreover, Kisswani and Nusair [2013] indicate that the data generating process of the oil price itself is characterized by nonlinear behavior.

conventional linear vector error correction model (VECM), which provides stable long-run equilibrium relations between observed variables. In the spirit of Engle and Granger [1987], these are allowed to depart from their equilibrium path in the short-run in response to random shocks, but those deviations are corrected in the long-run. The traditional methodology assumes that all parameters of the data-generating process are fixed over the whole sample period. However, because of the above-mentioned frequent changes in the equilibrium relationships, which may for example stem from exogenous shocks, the parameter-constancy assumption seems to be too restrictive. Thus, we use an approach which allows some parameters to vary.

Turning to the choice of an adequate framework, the empirical literature roughly distinguishes between two kinds of nonlinear models. If the data is generated primarily by market forces, a threshold VECM where the regime variable is endogenous might adequately describe the dynamics [Balke and Fomby, 1997; Lo and Zivot, 2001; Hansen and Seo, 2002]. However, if exogenous factors such as policy interventions or abnormal global economic crises affect the data, an MS-VECM is more suitable, since the regime variable is treated as an exogenous stochastic process [Ihle and von Cramon-Taubadel, 2008]. The oil price series under investigation obviously include several shocks from historical events such as the Iranian revolution of 1978/79, the Gulf War between 1980 and 1988, the abandonment of the Saudi Arabian swing producer role in 1985, the Iraqi invasion of Kuwait in 1990, the Asian financial crisis in 1997/98, production target cuts by OPEC in 1999, the attack on the World Trade Center on September 11, 2001, the shortage of spare capacity in 2005, the global financial crisis that started in 2007, and further production target cuts by OPEC in 2009. These exogenous events that affect our dataset (see Figure 1) can, in our view, appropriately be accounted for by the use of an MS-VECM. Compared to threshold models such as the class of smooth transition models introduced by Teräsvirta [1994], such a framework has the advantage that it does not require the specification of a particular variable that is held responsible for the switching between different regimes. Another benefit of the MS-VECM is that it allows the identification of potentially latent regimes in the data and helps to adequately specify the nonlinear dynamics between the variables.

The MS-VECM framework has been proven to be useful in cases where the data seems to be mainly driven by exogenous events. Against this background, it is not surprising that a large amount of research is based on the application of Markov-switching models. In a broad context, Francis and Owyang [2005] have conducted an MS-VECM approach while analyzing the long-run impacts of monetary policy of the US Fed allowing for state changes which in their view can help to explain the price puzzle. In their study the timing of switches to the 'high inflation expectations state' tends to be nearly synchronous with events such as recession and oil price shocks. Beckmann and Czudaj [2013a] followed this methodology to examine the inflation hedging function of gold for the US, the UK, the Euro Area, and Japan while distinguishing between periods characterized by a slow as well as a fast adjustment of the price level to gold price changes. Moreover, Calza and Zaghini [2009] provide evidence for an error correction of the Euro Area M1 money demand with Markovian shifts. In the context of exchange rates, Sarno and Valente [2006] demonstrate that allowing for regime-switching in the underlying data-generating process is an adequate characterization that is capable of capturing the impact of an either fixed or flexible exchange rate regime of the last century in France, Germany, Japan, and the UK on the dynamics of exchange rates and relative prices. Turning to the use of Markov-switching models in oil price literature from a broader perspective, Raymond and Rich [1997] have analyzed the relationship between oil price shocks and US business cycle fluctuations by the use of a generalized Markov-switching model. Lammerding *et al.* [2013] recently provided evidence for the existence of speculative bubbles in the price of oil by the use of a present-value oil price model in state-space form with Markovian shifts, which distinguish between two phases in the bubble process, i.e. one in which the oil price bubble is a stable process and one in which the bubble explodes. Cologni and Manera [2009] show asymmetric effects of oil price shocks on output growth by applying different Markov-switching autoregressive models for the G-7 countries. Reboredo [2010] applies Markov-switching models to show that oil price shocks have nonlinear effects on stock returns. Finally, Balcilar and Ozdemir [2013] examine the causal linkage between oil futures price changes and a sub-grouping of S&P 500 stock index changes while using a Markov-switching vector autoregressive (MS-VAR) model.

3 Data and econometric methodology

3.1 Data

In our study we use a monthly dataset including the oil price (o) and the index of consumer prices (CPI) of the USA as the foreign country (p^*) as well as the CPIs of twelve states regarded as domestic countries (p). Following Lizardo and Mollick [2010], we use the series of the West Texas Intermediate (WTI) nominal oil price expressed in US dollars per barrel provided by the Federal Reserve Bank of Saint Louis. The log of the series is shown in Figure 1. As can be seen, the latter shows the same movements, such as the Europe Brent and the Dubai crude oil price series, which are frequently used as well. However, observations prior to 1986 are not available for both.

Figure 1 about here

We consider Brazil, Canada, Mexico, Norway, and Russia as the major oil-exporting countries as well as the Euro Area, India, Japan, South Africa, South Korea, Sweden, and the UK as oil-importing countries, and use their nominal exchange rates (e) against the US dollar (domestic currency per one unit of US dollar, i.e., an increase in the nominal exchange rate implies a depreciation of the domestic currency). Besides data availability, the country choice was also motivated by the relative importance of the exchange rate against the dollar.[7] Despite the UK being categorized as a crude oil

[7] We have selected countries that satisfy the following criteria: (1) our set of countries should include both net oil importing as well as exporting countries; (2) the countries should be important trading partners of the US, since we analyze dollar exchange rates; (3) their currencies should be actively traded; and (4) there should be available data for the price level as well as the exchange rate against the US dollar for a reasonable time span. This is important, since our regime switching approach should only be applied if the series under investigation are long enough to display a switching pattern. To compare our findings with previous studies we have also paid attention to the set of countries included by other studies. We have selected this set of countries according to Chen and Chen [2007], Habib and Kalamova [2007],

importer, we bear in mind that its trade dependence on crude oil could be significantly different from other oil-importers due to its own oil-producing capacity. Generally speaking, the simple classification of oil exporters and oil importers may not be adequate with regard to the country-specific import-export structure of the economies under observation. However, the chosen distinction facilitates the interpretation of our findings and is sufficient for a comparison of previous empirical results. In addition, our approach treats all variables as endogenous, implying that each economy is examined in a similar fashion regardless of its classification. Table 1 summarizes our dataset and gives the sources.

Table 1 about here

A crucial question is the choice of an adequate starting point for the different economies. Since we are interested in disentangling nominal and real dynamics, including periods of fixed nominal exchange rate regimes is not sensible. To tackle this issue, a de facto classification of the exchange rate regime has been adopted for each economy. Hence, our study starts in January 1981, in January 1993, in January 1994, in January 1995, in October 1998, and in February 1999 for South Korea, India, South Africa, Mexico, Russia, and Brazil, respectively.[8] The analysis of the Euro Area starts in February 1980, for which the exchange rate prior to 1999 is represented by the European Currency Unit (ECU), a former basket of the currencies of the European Community. For the remaining economies, we apply the beginning of 1974 as the starting date. Hence, our sample period starts right after the occurrence of a major oil price shock due to the Arab oil embargo in 1973/74 and the breakdown of Bretton Woods.[9]

As is common practice, we take each series as natural logarithm in the following and display each series for the particular time horizon in Figures 2 and 3, respectively.

Figures 2 and 3 about here

In order to analyze the underlying long-run dynamics, it is important to assure the integration order of the time series under observation. In the present context, an important question is whether some variables, in particular prices, are integrated of order two, i.e. $I(2)$.[10] By the application of the augmented Dickey-Fuller (ADF) test and the more powerful Ng-Perron MZ_α test, we find that each series may be approximated as integrated of order one, i.e. $I(1)$ [Dickey and Fuller, 1979; Ng and Perron, 2001].[11] However, because the evidence is mixed in some cases, we have treated some prices in alternative specifications as $I(2)$. The results did not change qualitatively in those cases.[12] Applying

Lizardo and Mollick [2010], and Wang *et al.* [2013], who use a similar selection of countries. Due to data availability we have included more oil importers than exporters. We do not include two of the largest oil-exporting countries, Saudi Arabia and Iran, towing to the lack of reliable data.

[8]For South Korea, we have considered two models, one starting in 1981 and one after the large depreciation in 1997:07. Since the findings are similar in terms of causality, these are only reported for the former case, where a shift dummy ($D97$) has been introduced in the long-run relationship to account for the depreciation in 1997:07.

[9]Although the formal breakdown of Bretton Woods took place in 1973, some countries under observation pegged their currency against the dollar during 1973 but not at the beginning of 1974.

[10]Juselius and MacDonald [2004] treat the real exchange rate as an $I(1)$ variable in their analysis. For an analysis of PPP in an $I(2)$ framework see Johansen, Juselius, Frydman and Goldberg [2010].

[11]We applied an auxiliary regression either with a constant or with a linear trend plus constant, since a graphical inspection shows that some of the series, especially the CPIs, may exhibit a time-dependent mean (see Figures 2 and 3).

[12]The results of the unit root tests and the alternative specifications are available upon request.

the nonlinear unit root test proposed by Kapetanios, Shin and Snell [2003] does not change the overall findings.

3.2 Empirical framework

In the following we use an MS-VECM to examine the relationship between price differentials, exchange rates, and world oil prices for the set of countries mentioned above. As noted in the previous section, the linkage between these variables was subject to many far reaching economic events as well as policy interventions over the last four decades, leading potentially to a state-dependent behavior of the adjustment to deviation from the long-run relations, which we expect to be captured appropriately by an MS-VECM. The development of traditional state-dependent time series models, on which the MS-VECM described here is based, goes back to the seminal works by Hamilton [1989] and Krolzig [1997]. As will be specified below, the parameters of an MS-VECM are designed to take a constant value in each regime and to shift discretely from one regime to the other with different switching probabilities. The switches between different states are not counted as deterministic occurrences, but are assumed to follow an exogenous stochastic process.

Thus the baseline framework is an M-regime vector autoregression (VAR) of order $p+1$, which allows each parameter to be state-dependent:

$$\Phi(L)(s_t)X_t = \nu(s_t) + \varepsilon_t, \quad t = 1, \ldots, T, \tag{6}$$

where $X_t = [o_t, e_t, p_t^*, p_t]'$ represents a vector of the observed time series, $\Phi(L)(s_t) = I_4 - \Phi_1(s_t)L - \ldots - \Phi_{p+1}(s_t)L^{p+1}$ is an 4×4 matrix lag polynomial of finite order $p+1$ with

$$\left|I_4 - \Phi_1(s_t)z - \ldots - \Phi_{p+1}(s_t)z^{p+1}\right| \neq 0 \tag{7}$$

for $|z| \leq 1$, $\nu(s_t)$ denominates a vector of regime-dependent intercept terms, and ε_t describes a vector of error terms with regime-dependent variance-covariance matrix $\Sigma(s_t)$, $\varepsilon_t \sim N(0, \Sigma(s_t))$. The stochastic regime-generating process s_t is assumed to follow an ergodic, homogenous, and irreducible first-order Markov chain with a finite number of regimes, $s_t \in \{1, \ldots, M\}$, and constant transition probabilities:

$$p_{ij} = Pr(s_{t+1} = j|s_t = i), \quad p_{ij} > 0, \quad \sum_{j=1}^{M} p_{ij} = 1 \quad \forall \quad i, j \in \{1, \ldots, M\}\,. \tag{8}$$

The first expression of Equation (8) gives the probability of switching from regime i to regime j at time $t+1$, which is independent of the history of the process. p_{ij} is the element in the ith row and the jth column of the $M \times M$ matrix of the transition probabilities P, which is usually not symmetric. The ergodicity assumption implies a stationary unconditional probability distribution of the regimes, the homogeneity assumption defines the transition probabilities to be constant, and, finally, the irreducibility assumption assures that any regime can be reached from any other regime [Beckmann and Czudaj, 2013a].

With the notion that $\Pi(s_t) := -[I_4 - \Phi_1(s_t) - \ldots - \Phi_{p+1}(s_t)]$ and $\Gamma_i(s_t) := -[\Phi_{i+1}(s_t) + \ldots + \Phi_{p+1}(s_t)]$ with $i = 1, \ldots, p$, Equation (6) can be rearranged to an M-regime pth order MS-VECM, which in general allows for regime shifts in the vector of intercept terms, the autoregressive part, the long-run matrix, and the variance-covariance matrix of the errors:

$$\Delta X_t = \mu(s_t) + \Gamma(L)(s_t)\Delta X_{t-1} + \Pi(s_t)X_{t-1} + \varepsilon_t, \tag{9}$$

where Δ denotes the difference operator and $\mu(s_t)$ denominates a vector of regime-dependent intercept terms. The 4×4 matrix lag polynomial $\Gamma(L)(s_t) = I_4 + \Gamma_1(s_t)L + \ldots + \Gamma_p(s_t)L^p$ of order p denotes the state-dependent short-run dynamics of the model.

The non-stationary behavior of the series is accounted for by a reduced rank ($r < 4$) restriction of the state-dependent 4×4 long-run level matrix $\Pi(s_t)$, which can be fragmented into two $4 \times r$ matrices $\alpha(s_t)$ and β such that $\Pi(s_t) = \alpha(s_t)\beta'$. β' gives the coefficients of the variables for the r long-run relations, which are assumed to be constant over the whole sample period, while $\alpha(s_t)$ contains the regime-dependent adjustment coefficients describing the reaction of each variable to disequilibria from the r long-run relations given by the r-dimensional vector $\beta' X_{t-1}$. Thus, the most interesting part of our study is the speed with which deviations from the long-run equilibria are corrected, given by $\alpha(s_t)$.

In order, firstly, to identify the rank of $\Pi(s_t)$, i.e. the number of cointegrating relations r, and to estimate the coefficients of the r cointegrating vectors in β', we rely on the standard framework developed by Johansen [1988, 1991]. Then, conditional on these cointegrating vectors, the regime-dependent adjustment parameters $\alpha(s_t)$, intercept terms $\mu(s_t)$, autoregressive coefficients $\Gamma(L)(s_t)$, and variance-covariance matrix $\Sigma(s_t)$, as well as the transition probabilities are all estimated using a Markov Chain Monte Carlo (MCMC) method, namely the multi-move iterative Gibbs sampling procedure proposed by Krolzig [1997] and described below.[13] This two-step framework is adopted from the work of Saikkonen [1992] as well as Saikkonen and Luukkonen [1997], who showed that the Johansen procedure provides consistent estimates for the cointegrating vectors, even in the presence of regime-switching.[14] See Appendix for details.

When using MS models, an important issue is the identification of the appropriate configuration of the model. To tackle this issue, we apply the bottom-up technique suggested by Krolzig [1997]. According to this approach, the most accurate characterization of the MS model can be achieved by restricting the effects of regime shifts on a particular set of parameters and testing the model against a supposedly unrestricted alternative. Therefore, we checked if the intercept terms, the autoregressive part, the variance-covariance matrix of the errors, and the adjustment coefficients of the VECM are state-dependent by applying several likelihood ratio (LR) tests, as proposed by Krolzig [1997]. The

[13]To check for robustness, we also conducted the expectation-maximization algorithm instead of the Gibbs sampler and obtained almost the same results.

[14]Francis and Owyang [2005] as well as Beckmann and Czudaj [2013a] followed this methodology in analyzing the long-run impact of the monetary policy of the US Fed and the inflation hedging function of gold, respectively. Both studies are restricted to the case of two regimes. In this study, we apply an unrestricted approach, which allows for M states.

LR test statistic is computed by $2(\ln L^* - \ln L)$, where L^* and L denote the maximum likelihood of the unrestricted and restricted configurations of the model, respectively. Hence, the resulting test statistic is asymptotically $\chi^2(k)$-distributed, where k equals the number of restrictions imposed.[15] The null of no regime dependence of the particular parameters is tested; thus, a rejection is in favor of a less restrictive model specification, allowing either some or all parameters to switch. We also test for the most suitable number of regimes, applying the same procedure. To save space, we do not report the corresponding LR test statistics. However, these suggest that the two-state MS-VECM, which allows for regime dependence in the adjustment parameters $\alpha(s_t)$ and the variance-covariance matrix $\Sigma(s_t)$ seems to be the most adequate specification of the relationship between prices, exchange rates, and world oil prices for each country under investigation in the next section.

As a check of adequacy, we also consider the regime classification measure (RCM) suggested by Ang and Bekaert [2002] in order to examine the regime classification performance of Markov-switching models of varying specifications. The statistic is computed as follows:

$$RCM(M) = 100M^2 \frac{1}{T_j} \sum_{t=1}^{T_j} \prod_{j=1}^{M} \tilde{p}_{j,t}, \tag{10}$$

where $\tilde{p}_{j,t}$ denotes the smoothed probability for regime j and provides a degree of accuracy with which a model identifies regime switching behavior over the entire sample period or a particular sub-sample period. The regime variable is Bernoulli distributed and thus, the RCM corresponds to a sample estimator of its variance. It takes values between 0 and 100, with 0 representing a perfect regime classification performance and 100 denoting that the model fails to exhibit any information about the regime-dependence.

4 Empirical results

4.1 Model selection

In the following, we summarize the empirical findings for the economies under observation in two tables: Table 2 provides the results for oil-exporting countries, while Table 3 focuses on oil-importing countries.

Tables 2 and 3 about here

To identify an adequate setting for the deterministic components in each model, we follow the methodology of Juselius [2006]. Out of the five possible configurations, preliminary exclusion tests suggest that in most cases a deterministic trend should not be included into the cointegrating space. In those cases, we allow for deterministic trends in the data, but exclude them from the long-run relationships.

[15] The regularity conditions for the test could be violated in some cases, leading to a distribution that differs from the χ^2 distribution. Nevertheless we rely on the asymptotic distribution following Sarno and Valente [2006]. Thus, the results of the LR tests should be interpreted with caution. For details related to the testing procedure and its asymptotics, see Krolzig [1997]. For a discussion of the problems regarding this technique, see Hansen [1992] and Garcia [1998].

The choice of the lag length p of each MS-VECM setting is based on the Schwarz criterion and tests for autocorrelation effects, which are presented in Panel (c) of Tables 2 and 3. According to Rahbek, Hansen and Dennis [2002], the rank test results we gain in the following are still robust under the remaining ARCH effects in some cases. While excess kurtosis does not introduce a significant bias to the estimated cointegrating vectors, the findings are sensitive to excess skewness [Juselius and MacDonald, 2004; Juselius, 2006]. Because of the high skewness and kurtosis of some variables, dummies have been included in some cases. After introducing dummy variables to account for outliers, following the methodology described in Juselius [2006], the rejection of the assumption of normality is due to excess kurtosis, so that our results are still reliable. The corresponding statistics are available upon request.

Having tested for the overall adequacy of the model framework, we proceed with the determination of the rank, that is, identifying the number of stationary long-run relationships. This is a crucial step in our analysis, since the results of restriction and validity tests as well as the reliability of the estimation depend on the right choice of the rank (r). To identify the number of cointegrating relations r, we rely on the trace test developed by Johansen [1988, 1991] as mentioned above. Since US consumer prices and the price of oil are embedded in each configuration, we have tested for cointegration between both as a first step. The results, which are available upon request, suggest that a relationship between both quantities exists. Hence, we expect at least one long-run relationship for each configuration, since the results from subsystems should, from a theoretical point of view, continue to hold in larger systems [Juselius, 2006].

The trace test statistics for the economies under observation are given in Panel (a) of Tables 2 and 3.[16] For Brazil, Mexico, Russia, and South Africa, we pay special attention to the Bartlett-corrected trace test, which implements a small sample size correction due to comparably small sample sizes. In a borderline case between two choices for the rank, we also considered the recursively generated trace test statistics, also available upon request. For Canada, India, Norway, and South Africa the findings suggest only one cointegrating relationship. For all other settings, the trace test indicates two long-run relationships as being the most adequate choice. As a next step, we put the results regarding the short- and the long-run dynamics under closer scrutiny. The estimated coefficients of each cointegrating relation are given in Panel (d), and the LR tests of the model restrictions, which imply in each case that our model cannot be rejected at the 10% level, are presented in Panel (b). The regime-dependent parameters of our MS-VECM are estimated for each country by the Gibbs sampling technique mentioned above, and the corresponding results for the adjustment coefficients and the transition probabilities are displayed in Tables 2 and 3, Panels (e) and (f), respectively.

[16]The test statistic of the corresponding likelihood test, the so-called trace test, is given by $trace(r) = -T \sum_{i=r+1}^{4} log(1-\hat{\lambda}_i)$. Under the null of $4-r$ unit roots, $\lambda_i, i = r+1, \ldots, 4$, should behave like random walks and the test statistic should be small. Starting with the hypothesis of full rank, the rank is determined by using a top-bottom procedure until the null cannot be rejected [Juselius, 2006].

4.2 Results for oil-exporting countries

Starting with the findings for the oil-exporting countries, those for Brazil and Mexico display a similar pattern, with the first long-run relationship showing the relation between the domestic exchange rate and the price of oil, and the second relationship corresponding to the dynamics between the oil price and domestic as well as US prices. The first relation for Russia also links the exchange rate and the oil price, while the second relation gives a positive relationship between the oil price and US prices. For Russia and Brazil, an increase in oil prices is associated with a nominal appreciation of the domestic currency, while the opposite seems to hold in the case of Mexico. With regard to the second relation, the price differential of the domestic economy relative to the US increases, if the oil price rises for Mexico. Although strict proportionality between prices does not hold, a similar pattern can be observed for Brazil. In terms of real exchange rates, the stronger increase in domestic prices implies a real appreciation against the dollar. However, Russia seems to experience a real depreciation as a result of the positive link between the oil price and US consumer prices.

To explain the differing character of the long-run relationships and to analyze the issue of causality, a closer examination of the time-varying adjustment pattern is necessary. In the case of Brazil, the oil price adjusts to deviations from both long-run relationships in the first regime, since both adjustment coefficients are correctly signed and significant. The adjustment to the first cointegration relation, in particular, is in line with our considerations, based on the law of one price in Section 2.2: an appreciation of the domestic currency against the dollar lowers the domestic price of oil and increases prices through increasing demand, as suggested by the 'denomination channel'. However, in the second regime the oil price does not adjust to those errors, since both coefficients are insignificant. Thus, regime 1 can be regarded as the oil-adjustment regime. In the Russian case, regime 1 emerges as the exchange rate adjustment regime, since the exchange rate adjusts significantly to deviations from the first cointegration vector. Domestic prices also show an adjustment pattern in both regimes. The findings for the adjustment coefficients of Mexico display a greater complexity, since both the price for oil and the exchange rate adjust to the first relation in one regime. Both prices do also adjust to the second relation in the first regime.

Hence, Brazil seems to experience a real appreciation against the dollar as a result of both a nominal appreciation and a rise in relative prices. The results for Mexico and Russia are not clear-cut, as the reactions of the exchange rate and the price differential have contrary effects on the real exchange rate, because a nominal depreciation and an increase in relative domestic prices are observed. In the cases of Canada and Norway, the results are again more difficult to interpret, as the nominal exchange rate, consumer prices, and the oil price all enter the same long-run relationship. The findings for Canada display a long-run relationship between the price for oil, Canadian consumer prices, and the US dollar exchange rate. Since the oil price significantly adjusts in the first regime, this relationship may be related to the argumentation of the law of one price described in Equation (5), as an appreciation of the Canadian dollar coincides with an increase in the price of oil. This pattern is plausible, as the Canadian dollar is often labeled as the prime example of a commodity currency. In addition, Canada is one of the largest oil exporters to the US. In the case of Norway, the price differential of the domestic economy relative to the US is positively related to the oil price, implying a real appreciation of the domestic economy as a result of a rising oil price. However, the oil price and the exchange rate are positively related, which indicates that an increase in oil prices coincides with a nominal depreciation of the domestic currency. Similarly to the case of Mexico and Russia, the results are not clear-cut with respect to the real exchange rate.

Summing up, the findings provide some important insights, although an unambiguous pattern cannot be identified. Firstly, the results clearly demonstrate the need to account for nonlinearities, since the adjustment pattern significantly varies for each configuration. For Canada and Brazil, the findings clearly point to a causality from exchange rates to oil prices, in accordance with the 'denomination channel', or the law of one price, with a nominal depreciation of the dollar resulting in higher oil prices. The same result holds for Russia where the exchange rate adjusts. For Mexico and Norway, an increase in oil prices is related to a depreciation against the dollar. While identifying a clear causality pattern for Mexico is difficult, the relationship for Norway possibly mirrors a link between US prices and the oil price, with both quantities strongly adjusting to long-run errors. Our findings also highlight the importance of distinguishing nominal and real effects, since the link between oil prices and the real exchange rates is in many cases not clear-cut.

4.3 Results for oil-importing countries

Turning to the results for oil importers, the findings for South Africa are disappointing in the sense that neither the nominal exchange rate nor the oil price adjusts to the single long-run relationship, with only US prices showing adjustment. The results for the Euro Area are similar to those for Brazil and Canada in the sense that a domestic appreciation against the dollar leads to an increase in the oil price, since both quantities are inversely related and the oil price adjusts in the first regime. Similarly to Brazil, the second relationship gives a positive relationship between the oil price and relative domestic prices, resulting in a real appreciation.

For the UK and Sweden, one relationship seems to link domestic prices, the oil price, and the exchange rate. While an increase in oil prices coincides with a nominal depreciation of the British pound, the opposite is true for the Swedish krona. Considering that the oil price is positively related to the relative domestic price differential according to the second relation for Sweden, the results indicate a real appreciation in this case. In the Swedish case, the oil price again significantly adjusts in the first regime, providing further evidence of an increase in oil prices as a result of a dollar depreciation, as suggested by the 'denomination channel'. With regard to the second relation, oil and consumer prices both adjust. For the UK, the exchange rate adjusts to deviations from the first relation in one regime while the second relation corresponds to both prices without including the exchange rate and the oil price. Hence, a direct link between the oil price and the price differential cannot be identified. Finally, the first long-run relationship for Japan includes the yen/dollar exchange rate, the oil price, and the price differential. The price for oil is positively related to both the price differential of the UK relative to the US and the nominal exchange rate. Hence, an increase in the oil price unambiguously coincides with a real depreciation of the yen. According to the adjustment coefficients, both CPIs and the oil price are responsible for such a pattern. The second relationship excludes oil prices and the nominal exchange rates, displaying a stationary relationship between both prices. Finally, a direct relationship between the nominal exchange rate and the oil price is observed for both India and South Korea. While an increase in the oil price coincides with a depreciation of the Indian currency, the opposite pattern occurs for South Korea. The second long-run relationship for Korea mirrors a PPP relationship. For both economies, the exchange rate adjusts to this relation in both regimes, but with different speed. Oil prices do not show a significant adjustment mechanism.

Altogether, the results again do not provide a clear pattern. The main finding that oil prices actually adjust more often than exchange rates continues to hold. The effects of oil prices on real exchange rates are in most cases again not clear-cut.

4.4 Is there an unambiguous causality pattern?

For both importers and exporters, the transition probabilities are highly significant for almost every country and show that the regimes are generally persistent in each case. The smoothed probabilities shown in Figure 4 illustrate the reconstructed incidences of the first regime over the whole sample period by inferring the probabilities of the occurrence of the unobserved states conditional on the available information in the whole dataset [Krolzig, 2003]. For regime classification, Hamilton (1989) suggested the use of a probability of 0.5. Especially in the case of Brazil and Mexico it becomes evident that the above-mentioned adjustment patterns seem to be related to times of turbulence, for example until the mid-80s or during the global financial crisis. However, in ‘normal times’, which are unaffected by major shocks from historical events there seems to be less adjustment within our model. This is consistent with the evidence that the oil price, dollar exchange rates, and the CPIs are cointegrated, since within our MS-VECM framework larger deviations from the long-run equilibrium seem to be corrected.

Figure 4 about here

Moreover, the RCM indicates a good fit for our MS-VECM for almost every country (see Panel (g) of Tables 2 and 3). For most of the countries the value of the RCM statistic is fairly low and thus our models do not seem to be misspecified. The only exception is Russia, with a value near 100, which is also indicated by the series of smoothed probabilities shown in Figure 4 (first column, third row). These lie in the range between 0.4 and 0.6 for the whole sample period and thus do not clearly identify the regime.

Although our findings do not provide a clear distinction between oil exporters and importers, some general conclusions can be drawn. Firstly, the pattern for oil exporters seems to be more complex, since both exchange rate and the oil price significantly adjust to long-run deviations in the case of Mexico. This is not the case for oil importers where only one of both quantities reacts to long-run disequilibria. We have found a causality from dollar exchange rates to oil prices, which is in line with theory in four out of twelve cases in the sense that dollar depreciations trigger an increase in oil prices and exchange rates do not adjust. In contrast, the UK, Russia, India, and South Korea represent examples of where exchange rate adjustment, but no oil price adjustment is observed. Interestingly, a domestic depreciation is now also related to a decrease in the oil price. However, the adjustment pattern for the UK shows that the exchange rate adjusts in one regime while contributing to disequilibria in the second. In addition, the adjustment seems to be also driven by prices. Nevertheless, the remaining exchange rate effect may be explained by the 'wealth channel' introduced in Section 2. For India and South Korea as oil importers, these wealth effects are also important: the different effects on the exchange rate might be explained by their relative attractiveness for oil exporters compared to other economies. Over the analyzed sample, it seems reasonable to assume that investors have favored South Korea more compared to India. When analyzing the causality pattern for Mexico, we see that the exchange rate adjustment may also mirror an adjustment to the price differential, according to purchasing power parity. However, a clear conclusion with regard to the issue of causality according to the adjustment pattern is not possible. The findings for Japan suggest that the oil price contributes to the adjustment, with the identified relationship mainly displaying a link between consumer prices and the price of oil. For South Africa, both variables seem to be weakly exogenous, with only prices driving the adjustment.

Overall, a reasonable conclusion is that the causality in nominal terms frequently runs from exchange rates to oil prices. However, a direct causality from nominal oil prices to nominal exchange rates is observed for India, Russia, and South Korea. In addition, an indirect influence from oil prices may stem from an effect on consumer prices which triggers effects on the nominal exchange rates. If the price differential changes as a result of a change in oil prices, the real exchange rate also varies by definition. Theoretically, this effect should be offset by adjustments from the nominal exchange rate. However, as mentioned in Section 2, the overall record suggests that nominal exchange rate adjustment to deviations from the price differential of two economies is mostly observed if deviations are large. Furthermore, a strict version of PPP is frequently rejected. It is important to keep in mind that the dataset under investigation does not distinguish between tradable and non-tradable goods, with PPP not expected to hold for the latter. Hence, the 'terms of trade channel', which was introduced in Section 2 and suggests an impact of oil price shocks on real exchange rates through prices, may well explain the ambiguous result regarding the real exchange rate for some economies. The argument of second round effects of the oil price on exchange rates may also be valid with regard to other macroeconomic fundamentals, which are linked to the exchange rate and react to oil price shocks. This mechanism may explain the results of Lizardo and Mollick [2010] and others in favor of incorporating the oil price into the monetary model of exchange rate determination.

An obvious question is why the results differ so much between the different currency pairs, since the nominal oil price and US prices, which are cointegrated in a bivariate system, are included in each configuration. In addition, similar results could be expected, since exchange rates are also known to display a high degree of co-movements. A straightforward explanation is provided by Siklos and Granger [1997] in their concept of regime-sensitive cointegration: common stochastic trends that are responsible for a long-run relationship may only be present during specific sub-periods, similar to the time-varying adjustment incorporated into this paper. If we consider that this issue implies a larger degree of complexity for systems including several variables, our ambiguous results are not surprising.

5 Conclusion

Although the link between oil prices and dollar exchange rates has been frequently analyzed, previous research has generally neglected three important issues in this context. Firstly, the role of prices and nominal exchange rate movements has yet to be disentangled. Secondly, the direction of the underlying causalities has not been analyzed in depth. And finally, nonlinearities have only been considered in a minor number of studies.

Based on an MS-VECM, which is able to distinguish between long-run and time-varying short-run dynamics, this study has accounted for those shortcomings. Firstly, we have demonstrated that the adjustment pattern is time-varying for all country pairs under investigation, suggesting that nonlinearities are an important issue when analyzing oil prices. In terms of causality, the results differ between the economies under observation but suggest that the causality in nominal terms, which runs from dollar exchange rates to oil prices, possibly mirrors the the 'denomination' and the 'adjustment channel'. On the one hand, changes in nominal oil prices trigger real exchange rate effects, through the price differential as well as through a direct influence on the nominal exchange rate. If both effects are observed, these often express an ambiguous pattern for the real exchange rate. Overall, the results differ within the group of oil-exporters and oil-importers, and fail to exhibit a clear pattern. However, with respect to these conclusions, a possible caveat stems from the fact that we have analyzed different periods of time for some of the countries under observation. On the other hand, the results are not surprising, since the proposed theoretical frameworks are rather simple and apply various restrictions. At the same time, empirical research has shown that both exchange rates and the oil price are notoriously difficult to model. What we can say is that the pattern for oil exporters seems to be more complex, since both the exchange rate and the oil price adjust to long-run errors for Mexico. This is not the case for oil importers where only one of both variables adjusts, mostly with a varying degree between the two regimes.

Our study has also highlighted another issue: there may be little gain from focusing solely on real exchange rates and real oil prices. It is well established that the nominal exchange rate and the oil price fluctuate much stronger compared to prices. Price linkages may of course be important over the long-run and therefore interesting for policymakers; however, nominal exchange rates and the nominal oil prices are much more important when it comes to modeling short-run linkages and adjustment behavior.

Finally, it is important to keep in mind that exchange rates exhibit a high degree of co-movements.

In the context of oil prices, a global panel analysis for each sub-group (oil importers and exporters), which accounts for the existence of cross-sectional dependence, might therefore be an interesting research topic. However, since our aim was to compare results on a country base, we leave this task for further research. Another potential extension left for future research could be seen in the inclusion of further macroeconomic variables such as industrial production, money supply, or stock prices into our framework. See Filis [2010] and Basher, Haug and Sadorsky [2012] for a discussion on this issue.

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Appendix

The intention of this Appendix is to describe the iterative Gibbs sampling procedure we followed. In doing so, we define an $(1 + 4p + r)M$-dimensional vector:

$$\begin{aligned} Z_t = [\mathbb{1}(s_t = 1) \quad \dots \quad \mathbb{1}(s_t = M) \quad \Delta X_{t-1}\mathbb{1}(s_t = 1) \quad \dots \quad \Delta X_{t-1}\mathbb{1}(s_t = M) \quad \dots \\ \Delta X_{t-p}\mathbb{1}(s_t = 1) \quad \dots \quad \Delta X_{t-p}\mathbb{1}(s_t = M) \quad \beta' X_{t-1}\mathbb{1}(s_t = 1) \quad \dots \quad \beta' X_{t-1}\mathbb{1}(s_t = M)]', \end{aligned} \tag{11}$$

where $\mathbb{1}(s_t = i)$ denotes an indicator function which equals 1 for regime i and 0 otherwise. Therefore, Equation (9) can be summarized as follows:

$$Y = \Xi Z + \varepsilon, \tag{12}$$

with

$$Y = [\Delta X_1 \quad \dots \quad \Delta X_T], \quad Z = [Z_1 \quad \dots \quad Z_T], \quad \varepsilon = [\varepsilon_1 \quad \dots \quad \varepsilon_T], \tag{13}$$

and the $4 \times (1 + 4p + r)M$ coefficient matrix:

$$\Xi = [\mu(s_t) \quad \Gamma_1(s_t) \quad \dots \quad \Gamma_p(s_t) \quad \alpha_1(s_t) \quad \dots \quad \alpha_r(s_t)]. \tag{14}$$

Conditional on Equation (12), the cointegrating matrix β, and a series of states $\tilde{s}_T = \{s_1, \dots, s_T\}$, coefficient values are drawn from the posterior normal-inverse Wishart distribution with uninformative priors ν_{01}, …, ν_{0M}, N_0, F_0, W_{01}, …, W_{0M}. By the application of uninformative priors, we model the cointegrating vectors explicitly [Francis and Owyang, 2005].

At each iteration step, Ξ and $\Sigma(s_t)$ for $s_t \in \{1, \dots, M\}$ are drawn from a distribution with ν degrees of freedom, precision matrix N, parameter means F, as well as variance-covariance matrices W_1, …, W_M, which are defined as follows for regime i:

$$\begin{aligned} \nu_i &= \nu_{0i} + \hat{T}_i, \quad N = N_0 + Z'Z, \quad F = N^{-1}\left(N_0 F_0 + Z'Z\hat{F}\right), \\ W_i &= \frac{\nu_0}{\nu} W_{0i} + \frac{\hat{T}_i}{\nu_i}\hat{\Sigma} + \frac{1}{\nu}\left(\hat{F} - F_0\right)' N_0 N^{-1} Z'Z \left(\hat{F} - F_0\right), \end{aligned} \tag{15}$$

where $\hat{F} = (Z'Z)^{-1} Z'Y$, $\hat{\Sigma} = \left(Y - Z\hat{F}\right)'\left(Y - Z\hat{F}\right)$, and $\hat{T}_i$ denominates the number of periods in state i.

Conditional on the data series $\tilde{X}_T$ and the drawn parameters Ξ and $\Sigma(s_t)\ \forall\ s_t$, the series of states $\tilde{s}_T$ is drawn from the posterior distribution $p\left(\tilde{s}_T | \tilde{X}_T, \Xi, \Sigma(s_t)\forall s_t\right)$, which is obtained from:

$$p\left(s_t | \tilde{X}_t, \Xi, \Sigma(s_t)\forall s_t\right) = \frac{f\left(X_t | \tilde{X}_{t-1}, s_t, \Xi, \Sigma(s_t)\forall s_t\right) p\left(s_t | \tilde{X}_{t-1}, \Xi, \Sigma(s_t)\forall s_t\right)}{\sum_{s_t} f\left(X_t | \tilde{X}_{t-1}, s_t, \Xi, \Sigma(s_t)\forall s_t\right) p\left(s_t | \tilde{X}_{t-1}, \Xi, \Sigma(s_t)\forall s_t\right)}, \tag{16}$$

where

$$p\left(s_t | \tilde{X}_{t-1}, \Xi, \Sigma(s_t)\forall s_t\right) = \sum_{s_{t-1}} p\left(s_t | s_{t-1}\right) p\left(s_{t-1} | \tilde{X}_{t-1}, \Xi, \Sigma(s_t)\forall s_t\right), \tag{17}$$

and $p\left(s_{t-1} | \tilde{X}_{t-1}, \Xi, \Sigma(s_t)\forall s_t\right)$ is given by each previous iteration step [Hamilton, 1989; Kim and Nelson, 1999]. The transition probabilities p_{ij} are also derived within this algorithm by drawing from posteriors formed from beta conjugate distributions [Kim and Nelson, 1999; Francis and Owyang, 2005].

Tables

Table 1: Data description

Country	Series	Unit	Sample period	Source
	WTI	$/barrel	1974:01-2011:12	Federal Reserve Bank of St. Louis
	Brent	$/barrel	1987:05-2011:12	EIA
	Dubai	$/barrel	1986:06-2011:12	Thomson Reuters
Brazil	CPI	2005=100	1979:12-2011:11	OECD
	Exchange rate vs. $	Brazil. Real	1992:12-2011:12	Thomson Reuters
Canada	CPI	2005=100	1974:01-2011:11	OECD
	Exchange rate vs. $	Canad. Dollar	1974:01-2011:12	Federal Reserve, United States
Euro Area	CPI	2005=100	1980:02-2011:11	ECB
	Exchange rate vs. $	Euro	1974:01-2011:12	WM/Reuters
India	CPI	2005=100	1974:01-2011:11	OECD
	Exchange rate vs. $	Indian Rupee	1993:01-2011:12	PACIFIC Exchange Rate Service
Japan	CPI	2005=100	1974:01-2011:11	OECD
	Exchange rate vs. $	Japanese Yen	1974:01-2011:12	Federal Reserve, United States
Mexico	CPI	2005=100	1974:01-2011:11	OECD
	Exchange rate vs. $	Mexican Peso	1986:07-2011:12	Global Treasury Information Services
Norway	CPI	2005=100	1974:01-2011:11	OECD
	Exchange rate vs. $	Norweg. Krone	1974:01-2011:12	Federal Reserve, United States
Russia	CPI	2005=100	1992:01-2011:11	OECD
	Exchange rate vs. $	Russian Ruble	1992:07-2011:12	Central Bank of the Russian Federation
South Africa	CPI	2005=100	1974:01-2011:11	OECD
	Exchange rate vs. $	South Afr. Rand	1974:01-2011:12	Federal Reserve, United States
South Korea	CPI	2005=100	1974:01-2011:11	OECD
	Exchange rate vs. $	South Kor. Won	1981:04-2011:12	PACIFIC Exchange Rate Service
Sweden	CPI	2005=100	1974:01-2011:11	OECD
	Exchange rate vs. $	Swedish Krona	1974:01-2011:12	Federal Reserve, United States
UK	CPI	2005=100	1974:01-2011:11	OECD
	Exchange rate vs. $	Pound Sterling	1974:01-2011:12	WM/Reuters
USA	CPI	2005=100	1974:01-2011:11	OECD

Note: Abbreviations: WTI: West Texas Intermediate, EIA: US Energy Information Administration, CPI: Index of Consumer Prices, OECD: Organization for Economic Co-operation and Development, ECB: European Central Bank.

Table 2: Empirical results – oil-exporting economies

Country	(a) Trace test H_0	p-v.	p^{α}-v.	(b) Test of rest. model	(c) Test for AC	(d) Long-run relations		(e) Adjustment coeff. Regime 1 α_1	Regime 1 α_2	Regime 2 α_1	Regime 2 α_2	(f) Trans. prob. Regime 1	Regime 2	(g) RCM
	r=0	.000	.000	$\chi^2(3)$	LM(1): .083	$\underset{(4.605)}{.294^{***}} o + e$	Δo	−.172*** (−2.646)	−.407*** (−2.660)	−.043 (−1.483)	−.113 (−1.638)			
Brazil	r=1	.004	1.000	=5.757	LM(2): .095		Δe	.027 (.551)	−.007 (−.061)	.005 (.263)	.008 (.154)	.648*** (3.927)	.352** (2.133)	46.220
	r=2	.214	1.000	[.124]	LM(3): .705	$o + \underset{(6.525)}{1.566^{***}} p^* \underset{(-10.517)}{-2.600^{***}} p$	Δp^*	−.004 (−1.333)	−.009 (−1.286)	.001 (1.000)	.001 (.500)	.088 (1.060)	.912*** (10.988)	
	r=3	.345	.988		LM(4): .116		Δp	.006 (1.200)	.005 (.455)	.002 (.667)	.001 (.167)			
	r=0	.000	.000	$\chi^2(2)$	LM(1): .000	$\underset{(-6.553)}{-.273^{***}} o \underset{(-1.831)}{-.349^{*}} e + p$	Δo	.001 (1.000)		−.002 (−1.000)				
Canada	r=1	.723	.996	=3.636	LM(2): .572		Δe	.000 (1.000)		.001 (1.000)		.977*** (97.700)	.075** (2.500)	8.226
	r=2	.797	.987	[.162]	LM(3): .760		Δp^*	.000 (1.000)		.001** (2.000)		.023** (2.300)	.925*** (30.833)	
	r=3	.857	.921		LM(4): .060		Δp	.000 (1.000)		.002*** (4.000)				
	r=0	.000	.000	$\chi^2(3)$	LM(1): .083	$o \underset{(-5.380)}{-5.144^{***}} e + \underset{(4.631)}{10.626^{***}} \mu$	Δo	−.119*** (−2.833)	.045*** (3.214)	−.001 (−.100)	.006 (1.200)			
Mexico	r=1	.004	.705	=4.581	LM(2): .104		Δe	.046** (2.000)	−.007 (−.875)	−.002 (−.667)	.002 (1.000)	.711*** (7.813)	.049** (2.450)	7.952
	r=2	.088	.736	[.205]	LM(3): .547	$o + \underset{(9.183)}{11.676^{***}} p^* \underset{(-9.183)}{-11.676^{***}} p$	Δp^*	−.005*** (−5.000)	.002*** (−4.000)	.001** (2.000)	−.000 (−1.000)	.289*** (3.176)	.951*** (47.550)	
	r=3	.093	.585		LM(4): .101		Δp	−.002 (−.667)	.003*** (3.000)	−.001** (2.000)	.001** (2.000)			
	r=0	.000	.000	$\chi^2(1)$	LM(1): .002	$\underset{(-2.795)}{-.141^{***}} o + \underset{(2.376)}{.337^{**}} e - p^* + p + \underset{(7.886)}{.002^{***}} t$	Δo	.009* (1.800)		−.000 (−1.000)				
Norway	r=1	.075	.949	=.832	LM(2): .318		Δe	−.003 (−1.500)		−.030 (−1.000)		.970*** (107.778)	.166*** (3.952)	.000
	r=2	.122	.782	[.362]	LM(3): .629		Δp^*	.001** (2.000)		.018*** (6.000)		.030** (3.333)	.834*** (19.857)	
	r=3	.165	.291		LM(4): .001		Δp	.001** (2.000)		.014** (2.000)				
	r=0	.000	.000	$\chi^2(2)$	LM(1): .349	$\underset{(-6.884)}{-.042^{***}} o \underset{(-32.348)}{-.239^{***}} e + \mu$	Δo	−.813 (−.639)	.039 (.661)	−.095 (−.097)	.071 (1.340)			
Russia	r=1	.004	.963	=2.212	LM(2): .025		Δe	.619** (1.984)	.010 (.500)	.538* (1.843)	.011 (.611)	.904*** (13.101)	.082 (1.224)	98.098
	r=2	.009	.885	[.331]	LM(3): .291	$\underset{(-34.217)}{-.411^{***}} o \underset{(-11.045)}{-.994^{***}} p^*$	Δp^*	−.039* (−1.857)	−.001 (−.333)	−.015 (−.750)	.000 (.000)	.096 (1.391)	.918*** (13.702)	
	r=3	.079	.633		LM(4): .488		Δp	−.032 (−.941)	−.008*** (−2.667)	−.043 (−1.229)	−.007** (−2.333)			

Note: Panel (a) reports Johansen [1988, 1991] cointegration tests at which p^{α}-value refers to a simulation with T=400 and 2,500 replications. Panel (b) shows the test of the restricted model [p-value]. Panel (c) displays p-values of Lagrange multiplier (LM) tests for autocorrelation (AC). Panel (d) shows the estimates of the cointegration vector (t-stat.). Panel (e) reports the switching adjustment coefficients towards the long-run equilibrium for both regimes (t-stat.). Panel (f) reports the transition probabilities, while (g) gives the regime classification measure (RCM). */**/*** rejection of the null hypothesis at the 10%/5%/1% significance levels. See Section 4.2 for further explanations.

Table 3: Empirical results – oil-importing economies

Country	(a) Trace test H_0	p-v.	p^{α}-v.	(b) Test of rest. model	(c) Test for AC	(d) Long-run relations		(e) Adjustment coeff. Regime 1 α_1	Regime 1 α_2	Regime 2 α_1	Regime 2 α_2	(f) Trans. prob. Regime 1	Regime 2	(g) RCM
	r=0	.000	.000	$\chi^2(1)$	LM(1): .044	$\underset{(3.253)}{.020^{***}}o + \underset{(6.180)}{.486^{***}}p^* \underset{(-9.629)}{-.691^{***}}p + \mu$	Δo	$-.104$ $(-.623)$	$.061^{**}$ (2.033)	.100 (.855)	$-.004$ $(-.308)$			
Euro Area	r=1	.005	.232	=.023	LM(2): .263		Δe	.019 (.292)	$-.005$ $(-.455)$	.042 (.737)	$-.001$ $(-.125)$	$.932^{***}$ (35.519)	$.075^{***}$ (2.678)	26.593
	r=2	.059	.681	[.879]	LM(3): .460	$\underset{(-7.088)}{-.227^{***}}o \underset{(-3.066)}{-1.095^{***}}e + \mu$	Δp^*	.006 (1.200)	$.003^{***}$ (3.000)	$.019^{***}$ (6.333)	$.001^{**}$ (2.000)	$.068^{**}$ (2.519)	$.925^{***}$ (33.036)	
	r=3	.177	.609		LM(4): .005		Δp	$.034^{***}$ (11.333)	$-.000$ $(-.000)$	$.030^{***}$ (15.000)	.000 (.000)			
	r=0	.000	.000	$\chi^2(2)$	LM(1): .001	$e \underset{(-1.912)}{-.149^{**}}o \underset{(-38.482)}{-10.783^{***}}\mu$	Δo	$-.021$ (-1.414)		$-.019$ (-1.428)				
India	r=1	.159	.147	=.128	LM(2): .511		Δe	$-.005^{*}$ (-1.739)		.000 (.277)		$.912^{***}$ (26.824)	$.088^{***}$ (2.588)	31.686
	r=2	.793	.769	[.938]	LM(3): .001		Δp^*	$-.003^{***}$ (-4.966)		$-.003^{***}$ (-7.222)		$.088^{***}$ (2.667)	$.912^{***}$ (27.636)	
	r=3	.618	.599		LM(4): .114		Δp	$-.005^{***}$ (-3.481)		$-.004^{***}$ (-3.117)				
	r=0	.000	.000	$\chi^2(1)$	LM(1): .001	$\underset{(6.327)}{.021^{***}}o \underset{(-4.181)}{-.196^{***}}e - p^* + p + \underset{(4.951)}{.001^{***}}t$	Δo	$-.403^{**}$ (-2.277)	$-.009^{**}$ (-2.250)	$-.025$ $(-.253)$	$-.001$ $(-.333)$			
Japan	r=1	.001	.055	=.328	LM(2): .123		Δe	.039 (.639)	$-.001$ $(-.500)$	$-.026$ $(-.482)$	$-.002$ (-1.000)	$.937^{***}$ (54.056)	$.060^{**}$ (2.222)	35.128
	r=2	.493	.839	[.567]	LM(3): .535	$\underset{(-9.306)}{-.514^{***}}p^* + p + \underset{(6.084)}{.001^{***}}t$	Δp^*	.000 (.000)	$.001^{**}$ (2.000)	$.010^{**}$ (2.500)	$.001^{**}$ (2.000)	.027 (1.500)	$.940^{***}$ (34.815)	
	r=3	.402	.598		LM(4): .024		Δp	$-.006$ $(-.600)$	$-.000$ $(-.000)$	$-.031^{***}$ (-2.385)	$.001^{**}$ (2.000)			
	r=0	.000	.000	$\chi^2(2)$	LM(1): .031	$\underset{(-6.884)}{-1.371^{***}}o + e \underset{(-2.578)}{-4.729^{**}}p^* + \underset{(2.578)}{4.729^{**}}p$	Δo	$-.004$ $(-.800)$		.041 (1.414)				
South Africa	r=1	.137	1.000	=3.581	LM(2): .008		Δe	$-.002$ $(-.667)$		$-.012$ $(-.800)$		$.975^{***}$ (31.452)	$.221^{**}$ (2.105)	12.365
	r=2	.452	1.000	[.167]	LM(3): .190		Δp^*	$-.001^{**}$ (-2.000)		.001 (1.000)		.025 (.806)	$.779^{***}$ (7.419)	
	r=3	.437	.NA		LM(4): .391		Δp	$-.001^{**}$ (-2.000)		$-.001$ (-1.000)				
	r=0	.000	.000	$\chi^2(3)$	LM(1): .000	$e \underset{(-2.096)}{-.118^{***}}o \underset{(-.699)}{-.700}D97 \underset{(-34.472)}{-6.745^{***}}\mu$	Δo	$-.017$ $(-.472)$	.015 (.714)	.003 (.077)	.013 (.382)			
South Korea	r=1	.002	.002	=.933	LM(2): .016		Δe	$-.046^{***}$ (-3.067)	$-.013^{*}$ (-1.625)	$-.010^{***}$ (-3.333)	$-.005^{*}$ (-1.667)	$.956^{***}$ (39.833)	$.044^{*}$ (1.833)	15.118
	r=2	.103	.090	[.817]	LM(3): .026	$p^* - p + \underset{(2.713)}{.456^{***}}e \underset{(-2.401)}{-2.782^{***}}\mu$	Δp^*	$-.002^{**}$ (-2.000)	$.002^{**}$ (2.000)	-0.002^{**} (-2.000)	$.002^{***}$ (2.000)	$.036^{***}$ (2.118)	$.964^{***}$ (56.706)	
	r=3	.452	.423		LM(4): .081		Δp	$-.003$ (-1.500)	$.004^{***}$ (4.000)	$-.006^{**}$ (-2.000)	.002 (1.000)			
	r=0	.000	.000	$\chi^2(1)$	LM(1): .014	$\underset{(-6.363)}{-3.475^{***}}o \underset{(-5.677)}{-.509^{***}}e + p + \underset{(9.919)}{.029^{***}}t$	Δo	$.006^{**}$ (2.000)	$-.011^{**}$ (-2.200)	.000 (1.000)	.000 (1.000)			
Sweden	r=1	.003	.721	=2.446	LM(2): .465		Δe	$-.002$ (-1.000)	.002 (.667)	$-.011$ $(-.846)$	.022 (.957)	$.969^{***}$ (60.563)	$.188^{***}$ (3.917)	2.938
	r=2	.048	.791	[.118]	LM(3): .804	$\underset{(-6.049)}{-1.267^{***}}o - p^* + p + \underset{(10.645)}{.012^{***}}t$	Δp^*	$.001^{**}$ (2.000)	$-.002^{***}$ (-4.000)	$.004^{***}$ (4.000)	$-.009^{***}$ (-3.000)	$.031^{*}$ (1.938)	$.812^{***}$ (16.917)	
	r=3	.102	.186		LM(4): .034		Δp	.000 (1.000)	$-.002^{**}$ (-2.000)	.004 (1.333)	$-.009^{***}$ (-1.800)			
	r=0	.000	.000	$\chi^2(1)$	LM(1): .044	$\underset{(7.534)}{1.135^{***}}p^* \underset{(-8.944)}{-1.324^{***}}p + \mu$	Δo	.027 (.122)	$-.027$ $(-.091)$	.049 (.383)	$-.038$ $(-.216)$			
UK	r=1	.002	.284	=.084	LM(2): .263		Δe	$-.236^{*}$ (-1.873)	$.349^{**}$ (2.029)	$-.028$ $(-.239)$	.061 (.386)	$.514^{***}$ (2.888)	.119 (.783)	49.340
	r=2	.217	.746	[.771]	LM(3): .460	$\underset{(4.095)}{.045^{***}}o \underset{(-6.301)}{-.254^{***}}e \underset{(-29.555)}{-.259^{***}}p + \mu$	Δp^*	$-.011^{***}$ (-2.750)	$.021^{***}$ (4.200)	$-.005^{**}$ (-2.500)	$.016^{***}$ (5.333)	$.486^{***}$ (2.730)	$.881^{***}$ (5.796)	
	r=3	.153	.428		LM(4): .005		Δp	.014 (1.167)	.022 (1.158)	$.010^{**}$ (1.667)	.005 (.625)			

Note: See Table 2 for details.

Figures

Figure 1: Logarithms of different oil price series

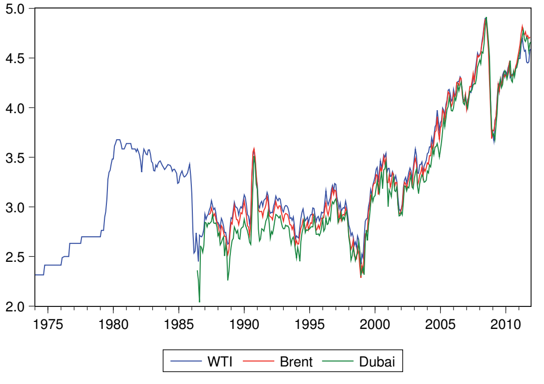
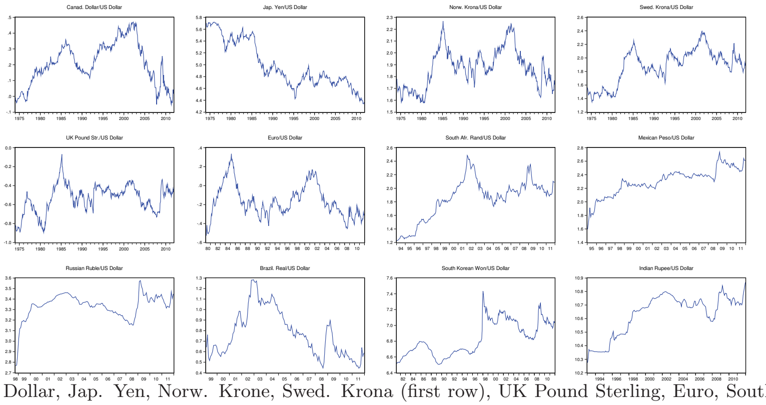
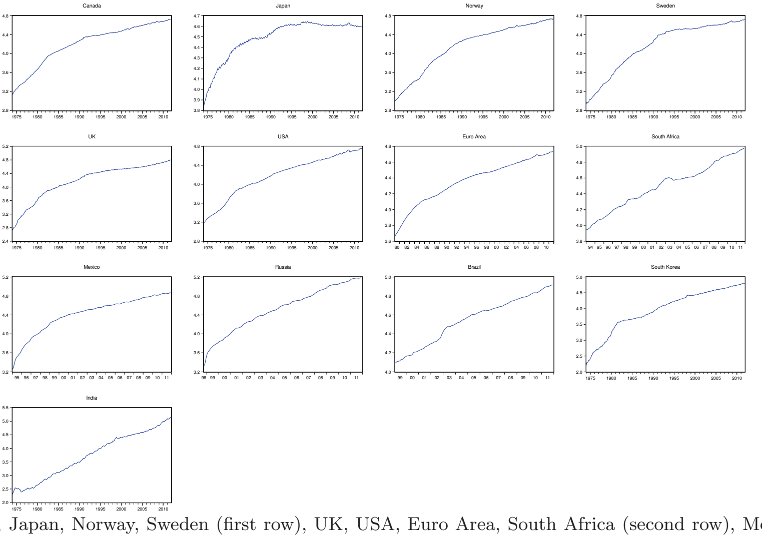


Figure 2: Series of logarithms of 12 currencies against the US dollar



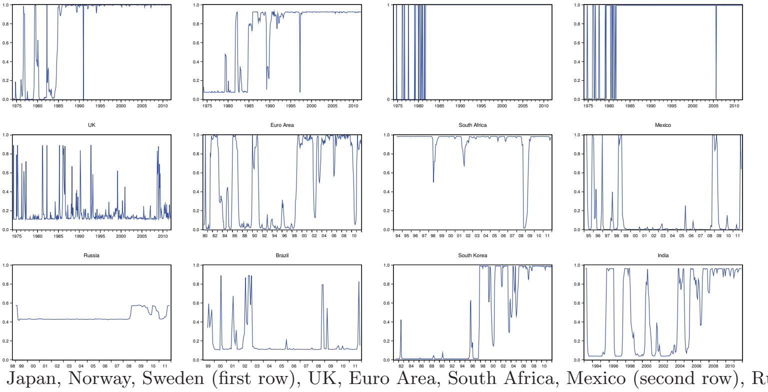
Note: Canad. Dollar, Jap. Yen, Norw. Krone, Swed. Krona (first row), UK Pound Sterling, Euro, South Afr. Rand, Mexican Peso (second row), Russian Ruble, Brazil. Real, South Kor. Won, Indian Rupee (third row).

Figure 3: Logarithms of 13 CPI series



Note: Canada, Japan, Norway, Sweden (first row), UK, USA, Euro Area, South Africa (second row), Mexico, Russia, Brazil, South Korea (third row), India (fourth row).

Figure 4: Smoothed probabilities for regime 1 for each country



Note: Canada, Japan, Norway, Sweden (first row), UK, Euro Area, South Africa, Mexico (second row), Russia, Brazil, South Korea, India (third row).