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Article

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Direct and Indirect Impacts of Oil Price Shocks on Ecuador's Economic Cycles (2000:01-2020:01)*

Impactos directos e indirectos de las perturbaciones del precio del petróleo en los ciclos económicos de Ecuador (2000:01-2020:01)

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Abstract

We analyse the non-linear relationship between oil price shocks and the real business cycle in Ecuador, a dollarized economy where oil exports are the country's main source of foreign exchange. We estimate several autoregressive Markov switching models for the period 2000:01-2020:01 to identify the differentiated impact of nominal oil price shocks on real GDP in expansion and slowdown regimes. We find evidence that oil price shocks have an asymmetric effect on Ecuador's economic growth, with a larger impact during slowdowns. They also affect all components of aggregate demand differently in each regime, with a larger impact on investment during expansions.

Key words: *Business cycle; Oil prices; Nonlinear models; Markov Regime-Switching Model; Ecuador.*

JEL Classification: *E32, Q02, F63, O11.*

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Resumen

En este trabajo se analiza la relación no lineal entre las perturbaciones del precio del petróleo y el ciclo económico real en Ecuador, una economía dolarizada en la que las exportaciones de petróleo son la principal fuente de divisas del país. Para ello se estiman varios modelos autorregresivos de conmutación de Markov para el período 2000:01-2020:01 con el fin de identificar el impacto diferenciado de las perturbaciones nominales del precio del petróleo sobre el PIB real en los regímenes de expansión y desaceleración económica. Los principales resultados muestran que las perturbaciones de los precios del petróleo tienen un efecto asimétrico en el crecimiento económico del país, con un impacto mayor durante períodos de desaceleración. También se observa que afectan a todos los componentes de la demanda agregada de forma diferente en cada régimen, con un mayor impacto en la inversión durante fases de expansión del ciclo económico.

Palabras clave: *Ciclo económico; precios del petróleo, modelos no lineales, modelo de cambio de régimen de Markov; Ecuador.*

Clasificación JEL: *E32, Q02, F63, O11.*

1. INTRODUCTION

For most developing countries, commodity trade remains the main source of foreign exchange (Bowman and Husain 2004). This structure prevails today in most Latin American countries, exposing them to exogenous price shock in primary products (Acosta 1998). Commodity price fluctuations have a significant impact on economic and social aggregates in developing countries and are responsible for deeper growth cycles.

In the case of Ecuador, it is a small oil exporter with no influence on international oil prices but whose dependence is seen as a major factor in the country's economic cycles (Martinez Valle 2008). The weight of the oil sector in the economy has been around 6% of GDP since 2000, reaching a peak of 14% in 2010-2011 (CepalSTAT¹). Domestic prices of oil derivatives have been intervened since 1973, always below international prices, with a staggered upward trend that has occurred every few years until 2018, and on an annual basis thereafter. This feature has largely insulated Ecuadorian demand from

¹ CEPALSTAT Bases de Datos y Publicaciones Estadísticas (<https://statistics.cepal.org/portal/cepalstat/index.html?lang=es>)

short-term oil price shocks, creating certainty and controlling inflation. However, investment is still affected by the oil shock on imported raw materials and capital goods. At the same time, the government allocates a high share of public expenditure to fuel subsidies, fluctuating between 32.1% in 2007 and 8.4% in 2017. (Muñoz-Miño 2018). However, since 2000, Ecuador has adopted the US dollar as its national currency, thereby losing its exchange rate policy and restricting its monetary policy. This has made the country more vulnerable to external shocks (Córdova Zambrano 2016). The country is currently embarking on large-scale mineral production, which will increase its dependence on commodities.

Several authors have analysed the transmission mechanisms of oil price changes on real variables and, ultimately, on the business cycle. Kilian (2008) points out that high oil prices cause economic crises in developed countries, most of which are net importers of this commodity. The direct or indirect transmission channels, ultimately affect aggregate demand and supply, either through changes in relative prices or through increases in production costs and uncertainty (Gonzalez and Hernandez 2016). In the case of primary exporters such as Ecuador, where oil exports are an important source of revenue for the public budget, this relationship should be the opposite.

The methodological strategy of most studies analysing the relationship between oil price shocks and economic growth was based on the application of linear models. However, several scholars find that commodity price shocks have different effects on output depending the business cycle (Raymond and Rich 1997; Clements and Krolzig 2002; Cologni and Manera 2009; Bjørnland *et al.* 2018; Cross *et al.* 2021). All of them agree that these dynamics are best characterized by Markov switching models.

With these considerations in mind, the aim of this paper is twofold. First, we analyse the business cycles of the Ecuadorian economy during the period 2000 to 2020 (quarterly series).² Second, we study the impact of oil price shocks on the business cycle in Ecuador, taking into account the presence of asymmetries or non-linearities in their relationship. A Markov Regime-Switching Autoregressive (MSAR) model is used to identify the regime shifts between expansions and contractions and how oil price shocks may have contributed to them. Instead of considering growth rates of both variables, we use trend deviations of both series to ensure stationarity of the underlying series.

The rest of the paper is structured as follows. The next section reviews the relevant literature on business cycles and their relationship with commodity price shocks in developing countries, with a particular focus on oil price fluctu-

² The sample period starts from 2000:01 due to the unavailability of previous data for quarterly GDP and its components at the Central Bank of Ecuador. The latest quarter analysed is 2020:01 in order to avoid the impact of COVID-19.

ations. This is followed by a description of the Markov switching (MS) model first applied by Hamilton (1989), which provides a better fit to time series with important structural changes. We then include in the MS model the impact of exogenous oil shocks on real GDP cycle, following Raymond and Rich (1997), Clements and Krolzig (2002), Holmes and Wang (2003), Cologni and Manera (2009) or Balcilar (2017). This is followed by a discussion of the different alternatives used with respect to the variables considered in the empirical analyses (nominal or real, in levels, differences or deviations from trend) and introduce the data sources. We then present the main results obtained for the Ecuadorian economy during the period 2000:01-2020:01 to reflect the expansionary and contractionary states of the country and its relationship with oil price shocks. The last section concludes.

2. ECONOMIC EFFECTS OF OIL PRICE SHOCKS

Following the oil shocks of the 1970s, a large literature on oil price shocks and their impact on macroeconomic aggregates emerged (Mork 1989), especially in the United States, a country with a significant and secular oil trade deficit. Most of them found an inverse relationship between oil price growth and output growth (Hamilton 1983, 1996b 2011; Mork 1989; Lee *et al.* 1995; Raymond and Rich 1997; Clements 2002; Kilian 2010 2014; Kilian and Vigfusson 2011 among others). Blanchard and Gali (2007), Peersman and Robays (2012), Calvacanti and Jalles (2013), Taghizadeh-Hesary *et al.* (2016), find the same result for net oil importers.

Most studies have focused on investigating the possible non-linear and asymmetric relationship between oil price shocks and economic growth. (Balcilar *et al.* 2017) and its transmission mechanism. Hamilton (1983) finds that major oil shocks (1973-1974, 1979, and 1980-1981) have been followed by major recessions in the US. Mork (1989) observes a negative effect on US output growth when oil prices rise and no correlation when oil prices fall. Hamilton (1996) obtains similar results. Raymond and Richard (1997) find that oil price shocks are responsible for shifts in the mean of some low-growth periods of output rather than the transition probabilities between growth states. Hamilton (2003) notes that oil price increases have a larger effect on GDP growth than oil price decreases. Cologni and Manera (2009) analyse the impact of seven different definitions of oil shocks (all in differences) on business cycle measured as the output growth. They find asymmetric effects of oil price shocks depending on the phase of the cycle for the G-7 countries and that their ability to explain recessionary episodes has declined over time due to improvements in energy efficiency and a better management of external supply and demand

shocks by monetary authorities. Herrera *et al.* (2011) observe that the results are sensitive to the estimation period and the aggregation level. Non-linear models have stronger supports for samples up to 1973, but samples with data after 1973 became much weaker. On the other hand, Kilian (2008) finds no evidence of asymmetries in the response of US demand to increases and decreases in energy prices.

Other studies have focused on the transmission channel of oil price shocks to real GDP and other relevant macroeconomic variables. On demand-side Hamilton (1988), Bresnahan and Ramey (1993), Kilian (2008) consider that oil shocks negatively affect US real GDP through consumer spending and business investment. On the supply side, McCallum (1989) finds that oil price increases are a prominent disruptor for industries, that have to pay for imported raw materials, especially energy. Hamilton (1983), Davis (1985), Loungani (1986) and Mork (1989) find reallocation effects of energy price shocks on capital and labour across sectors. However, Barsky and Kilian (2004) argue that energy price shocks should not be considered as aggregate supply shocks because they cannot be interpreted as productivity shocks to real GDP.

Over the past decade, a growing number of studies have also focused on emerging markets. Lescaroux and Mignon (2008) examine three groups of countries: oil importers, oil exporters and OPEC oil producers for different periods. They find that oil prices have a causal effect on GDP for oil importers and OPEC countries, but not for other oil exporters. Berument *et al.* (2010), using a set of small oil exporters and importers in the Middle East and North Africa between 1952 and 2005, find that a shock in oil prices (demand or supply driven) has a positive and significant effect on the growth of net oil exporting economies. In the case of oil-importing economies, output is found to fall for positive oil supply shocks, but rise with positive demand shocks. Alley *et al.* (2014), find a positive impact in oil exporting countries such as Nigeria. However, these shocks create uncertainty and undermine effective fiscal management of oil revenues.

Ahmadia and Manera (2021) find that the impact of oil shocks on the output of oil exporters varies across countries and depends strongly on the underlying cause of the oil shocks (demand or supply driven), as well as the economic health of each country. They also find no evidence of an asymmetric response of output to oil price rises or falls. Babuga and Ahmad (2022), for net oil exporters in Sub-Saharan Africa, whose economies are largely dependent on oil revenues for saving, investment and economic diversification, find a non-linear, inverted U-shaped relationship between the increase in oil prices above a certain threshold and real GDP.

For Latin American countries, Perilla (2010) observes a positive relationship between oil price shocks and the growth of the Colombian economy (a

net oil exporter) for the period 1990-2009. Gonzalez and Hernandez (2016) confirm this result for the period 1982-2013, that last 4 to 5 quarters after the shock. They suggest that private consumption serves as an indirect transmission channel of oil price shocks to GDP, especially in the period 2000-2013. Alarcón *et al.* (2016), find for Brazil, a net oil importer, a strong significant negative effect on economic growth over the period 1991:01-2014:01. For Colombia and Peru, the result is less significant perhaps due to domestic oil price controls that allow industries to be less sensitive by oil price shocks, in line with Blanchard and Gali (2007) and Uribe and Ulloa (2011). For Ecuador, Paladines (2017) and Paladines and Paladines (2017) and Peralta (2020), using annual data, find a positive impact of oil price shocks on output per capita in the following two years, before returning to the initial level.

3. OIL SHOCKS EFFECTS ON ECONOMIC CYCLES: A MARKOV-SWITCHING ANALYSIS

The study of business cycles has increased significantly in developed countries, especially in the United States, thanks to the efforts of the National Bureau of Economic Research (NBER), which have abundant information on business cycles and their impact on the different economic variables, especially employment (Mejía-Reyes 2003). Time series analysis has shown that the regression parameters are not constant over time and that there are structural changes that divide the time series into different regimes with different dynamic patterns over time. Nelson and Plosser (1982), Neftci (1984), De Long and Summers (1984), Watson (1986), Hamilton (1989) are among the first to note the existence of nonlinearities or asymmetries in economic variables and business cycles.

Since Hamilton (1989), a growing number of researchers have analysed these asymmetries using Markov Switching regression (Filardo 1994, Durland and McCurdy 1994; Hansen 1996). For Latin America, we find Mejía-Reyes (2000), Salamanca Lugo (2012) or Bayancela (2016) for Ecuador.³ However, these studies only consider univariate autoregressive models in which the business cycle is explained by GDP growth. As Blanchard and Quah (1989) point out, the analysis of GDP alone is not sufficient to characterise the effects of both supply and demand shocks (Kuan 2002).

³ Mejía-Reyes (2000) uses multiple univariate Markov Switching autoregressive (MSAR) models in eight economies, finding asymmetries in their business cycles, with recessions being deeper in absolute terms, less persistent and more volatile than expansions. Salamanca Lugo (2012), uses a Markov-Switching vector autoregressive regression model (MSVAR), to analyse the presence of a common cycle between Colombia, Venezuela and Ecuador, in which the fluctuations of each economy are characterized by similar movements of productive activity, with marked asymmetries between expansion and contraction phases. Bayancela (2016), applies a Markov regime-switching model, to explain the economic cycles in Ecuador for the period 1997-2015.

Hamilton (1996) is the first to include a dynamic specification of Markov switching models that depend on a vector of observable exogenous variables. However, he does not analyse their impact on the variable of interest. Raymond and Rich (1997) use a generalised Markov switching model to examine the influence of net real oil price increases on post-war US business cycle fluctuations (GDP growth) and whether they help to predict transitions between periods of positive and negative growth. Clements and Krolzig (2002) use a cointegrated Markov-switching vector autoregressive model (MS-VAR) with three-states and note that business cycle asymmetries do not appear to be explained by oil prices. Cologni and Manera (2009) examine the impact of oil shocks on the G-7 business cycle and find an asymmetric effect of oil price growth on output growth using different MS-VAR models. They find that models with exogenous oil variables generally outperform the corresponding univariate specifications. Balcilar *et al.* (2017) use a Bayesian Markov switching vector autoregressive model and found that oil price shocks affect South African real output growth under the low growth regimes. Bjørnland *et al.* (2018) take a different approach to analysing the role of oil price volatility in US macroeconomic variables and monetary policy. Based on Liu, Waggoner, and Zha (2011) and Bianchi (2013) they use a New Keynesian Markov switching rational expectations model in a DSGE framework. They find that the decline in oil volatility since 1985 is the most important factor reducing macroeconomic variability in the US. Živkov and Đuraškovi (2023) use MS-GARCH models to investigate how oil price uncertainty affects real GDP and industrial production in eight Central and Eastern European countries (CEEC). They find that oil price uncertainty has a small effect on output in moderate market conditions in the selected countries. On the other hand, in periods of deep economic crisis, an increase in oil price uncertainty reduces output, thereby adding to recessionary pressures in the economy. Conversely, when the economy is in expansion, oil price uncertainty has no effect on output.

Regime switching models were introduced into the literature by Quandt (1972), who examined time series processes that can exhibit random structural changes in which the switching events are independent over time (Kuan 2002). Subsequently, Goldfeld and Quandt (1973), Miron *et al.* (1987), Hamilton (1989) proposed a Markov regime switching model to analyse the US business cycle where switching events depend on the immediate past state.

The Markov switching model contains multiple structures that can capture nonlinear dynamics and sudden changes in the variability of a stationary time series autoregression (Hamilton 1996).⁴ A general extension of the Markov

⁴ Markov switching models require stationary data with zero mean. If the series have a unit root, the switching intercept results in a deterministic trend with breaks in that series. One solution is to transform the by applying first differences (Kuan 2002).

switching vector autoregressions of order p and s regimes mean adjusted model [MSM(s)-VAR(p)] is presented in Krolzig (1998) and Clements and Krolzig (2002):

$$(1) \quad y_t - \mu_{s_t} = \sum_{k=1}^p \alpha_{ks_t} (y_{t-k} - \mu_{s_{t-k}}) + \varepsilon_{s_t},$$

where y_t is a stationary vector, $\varepsilon_{s_t} \approx i.i.d.N(0, \sigma_{s_t}^2)$ and all parameters (μ_{s_t} , α_{ks_t} , $\sigma_{s_t}^2$) depend on the realised regime, a latent variable s_t which is called a regime or state. In this model, there is an immediate one-off jump in the process mean after a regime change (μ_1 in regime s_1 , jumps immediately to μ_2 when regime changes to s_2). Krolzig (1998) also presents a model with smooth adjustment of the mean after the transition from one regime to another. In this case we can use a model with a regime-dependent intercept [MSI(s)-VAR(p)]:

$$(2) \quad y_t = v_{s_t} + \sum_{k=1}^p \alpha_{ks_t} (y_{t-k} - \mu_{s_{t-k}}) + \varepsilon_{s_t},$$

The Markov switching models are defined by the transition probabilities that determine the persistence of each regime (Kuan 2002). If we consider only two regimes, $s_t = i, j$ are the unobserved first-order Markovian state variables governing the transition between the two distributions of y_t which can be summarised in the following transition probability matrix (P):

$$(3) \quad \begin{bmatrix} p_{00} & p_{01} \\ p_{10} & p_{11} \end{bmatrix},$$

where $p_{ij} \equiv \Pr ob[s_t = j | s_{t-1} = i]$ ($i, j=0,1$) denotes the transition probabilities from state $s_{t-1} = i$ to state $s_t = j$, that satisfies $p_{00} + p_{01} = 1$ and $p_{01} + p_{11} = 1$. For example, p_{11} is the probability of being in state 1 in period t if the economy was in state 1 in $t-1$. Clements and Krolzig (2002) and Cologni and Manera (2008) support Raymond and Rich (1997)'s assumption that transition probabilities are time-invariant, i.e., the likelihood of transitioning between different states remain constant over the entire time period under consideration. This assumption is based in the ergodic property of the MS model.⁵

Other MS models have focused on state dependence in the variance of the error term. The Markov switching autoregressive conditional heteroscedasticity (MSARCH) model and a Markov switching generalised autoregressive conditional heteroscedasticity (MS-GARCH) model allow these differenc-

⁵ Other authors have considered endogenous switching models where the probability of switching regime can vary over time depending of the state of the economy (Chang *et al.* 2017; Bazzi *et al.* 2017; Benigno *et al.* 2020; Hubrich and Waggoner 2021). However, these models assume endogeneity of the switching process where there are structural breaks (Bhar and Hamori 2007).

es to be analysed. The first one assumes a conditional mean of the residuals for each state and the conditional variance as a function of the lagged squared residuals ($\sigma_{s_t} = \alpha_{0t} + \alpha_{1t}\varepsilon_{t-1}^2$). Here conditional variance captures recent shocks through the squared residuals. The MS-GARCH model extends the previous one by including the lagged values of the conditional variance ($\sigma_{s_t} = \alpha_{0t} + \alpha_{1s_t}\varepsilon_{t-1}^2 + \beta_{1s_t}\sigma_{t-1s_t}$). It additionally captures the persistence and asymmetry in the volatility of the residuals by combining recent shocks with past volatility, through the lagged conditional variance (Bauwens *et al.* 2018). When $\alpha_{1t} + \beta_{1s_t}$ is statistically significant there will be conditional heteroscedasticity in the dispersion of the error term (Ardia *et al.* 2019).

4. VARIABLE SPECIFICATIONS

The effect of oil price shocks on business cycles has been analyzed using different models and variable specifications. There is consensus in the literature on the use of GDP in real terms, but there are different criteria on whether to include nominal or real oil prices. Some authors use nominal prices (Hamilton 1983, 2008; Jimenez-Rodriguez 2009; Alquist *et al.* 2013; Abdulkareem and Abdulhakeem 2016; Balcilar *et al.* 2017; Karaki 2017; Majidli 2020; Dwipa and Wicaksono 2021, just to cite a few). Others use the real price of oil by deflating the nominal price with the US consumer price index (Mork 1989; Ferderer 1992; Hooker 1996; Raymond and Rich 1997; Clements and Krolík 2002; Holmes and Wang 2003; Kilian 2006; Cologne and Manera 2009; Berument *et al.* 2010; Cross *et al.* 2021). At this respect, Hamilton (1993:238) gives two reasons in favour of using nominal oil prices: “(1) the institutional argument is that nominal, not real oil prices track the historical petroleum shocks and are the exogenous variable belonging in a reduced-form regression, and (2) it is naive to assume that the expected change in the relative shadow price of oil equals the (possibly disequilibrium) market price divided by a contemporaneous price index”. We ran all the models with both real and nominal oil prices. The results were similar, although the fit was lower in the second case.

Another issue is whether these variables should be included in levels or in differences (change in the natural logarithm). In order to represent the business cycle, most of the literature cited has used GDP growth rates, measured as the percentage change in real GDP from one period to another. Positive growth rates indicate economic expansion, while negative growth rates denote contraction. The first-differencing method eliminates the trend component, but it exacerbates the effect of high frequency noise (Stock and Watson 1999). Alternatively, the business cycles can be measured as deviations of actual GDP from its long-term trend. Positive deviations from the trend indicate above-average

economic activity, while negative deviations suggest below-average activity. Various statistical techniques are used to estimate the trend component of GDP, such as moving averages, Hodrick-Prescott (HP) or Baxter King (BK) filters or other time-series decomposition methods (Baxter and King 1994; Hodrick and Prescott 1997; Stock and Watson 1999; Orphanides and Van Norden 2002). The Central Banks of Ecuador, Chile, Mexico or Brazil calculates business cycles using GDP trend deviations based on OECD (1987).⁶ However, using GDP trend deviations alone does not assure stationarity of GDP (Stock and Watson 1999). We need to analyse the properties of the GDP series using unit root tests (Augmented Dickey-Fuller, Phillips-Perron). We will use this approach.

Another concern is to control for the seasonality of the series. Stock and Watson 1988; Finn 1991; Artis *et al.* 1997; Xiong 2015) provide evidence of quarterly GDP seasonality. In order to deseasonalise the quarterly series, structural time series models may be used when needed which include seasonal dummies in the regression (Baum 2006).

With respect to oil price shocks, most studies have used the growth rate (Hamilton 1983, 1996b, 2003; Gisser and Goodwin 1986; Mork 1989; Dotsey and Reid 1992; Hooker 1996; Kilian 2008; Lescaroux and Mignon 2008; Berument *et al.* 2010; Bergman 2019; Maheu *et al.* 2020). Hamilton (1996b) recommends to use an annual net oil price increase over the previous year. Raymond and Rich (1997) and Clements and Krolzig (2002) use the same variable. Just a few have estimated oil prices at levels (Huntington 2005; Gronwald 2008; Gozali 2010; ThankGod and Maxwell 2013). Hooker (1996) analyses oil prices in nominal log-differences and in real log-levels, however he gives three reasons for using levels: the price of oil appears to be bounded up and down. Carruth, Hooker and Oswald (1995) and Phelps (1994) have developed theoretical models which imply that firms' input prices are affected by the level rather than the first difference. Finally, the real price of oil is now roughly at the level of the 1950s and 1960s, which is consistent with stationarity. Since Markov switching models require stationary series, we will use trend deviations for all variables.

On the other hand, oil price shocks can be assumed a state-invariant covariate as in Raymond and Rich (1997), based on evidence from Hamilton (1983) and Cochrane (1994) that oil price changes are exogenous to the state of US economy. Clements and Krolzig (2002) use the same approach for a three-regime model. Others consider state dependent mean effects of oil price shocks (Mork 1989, Holmes and Wang 2003, Balcilar *et al.* 2017 or Živkov and Đurašković 2022). As in Cologni and Manera (2009) we will consider both

⁶ Central Bank of Ecuador uses the HP filter (Erraez 2014).

cases using different MS-VAR models.

The next step is to determine whether AR(p) or MA(q) terms are needed to correct for any remaining autocorrelation in the series (Becketti 2020). We should follow the principle of parsimony suggested by Box and Jenkins (1976), which implies that the simpler model (with fewer parameters) should be chosen. There are different approaches for a correct modelling of time series. Box-Jenkins (1970) propose an iterative process that involves four stages: identification, estimation, diagnostic checking and forecasting of time series (Wabomba *et al.* 2016). The identification process includes the analysis of the Auto Correlation Function (ACF) (for MA) and Partial Auto Correlation Function (PACF) (for AR). They can be complemented with Akaike’s (1974) information criterion (AIC) and Schwarz (1978) BIC, or the maximization of the mean log-likelihood. Muma and Karoki (2022) also propose to check autocorrelation with Ljung-Box Q-statistic, and Jarque-Bera (JB) test for normality of the residuals.

The analysis of Ecuador’s business cycle begins testing Hamilton’s (1989) MSI-AR model, a univariate autoregressive Markov Switching model for real GDP to a two-state process (expansion and slowdown), which permits for gradual adjustment of the series after the change in the state of the GDP cycle. Then we allow the AR coefficients (α_{ks_t}) and/or the variance ($\sigma_{s_t}^2$) to be function of each regime (MSIAH-AR model).⁷

Next, we add the non-linear effects of oil price shocks in the MS model. We will assume the exogeneity of oil prices with respect to Ecuadorian output, in line with Killian (2005, 2006), Raymond and Rich (1997), Clements and Krolzig (2002), Cologni and Manera (2009) or Berument *et al.* (2010). In order to deseasonalise the quarterly series, we included state dependent quarterly dummies. From equation (2) we obtain:

$$(4) \quad y_t = v_{s_t} + \sum_{k=1}^p \alpha_{ks_t} (y_{t-k} - \mu_{s_{t-k}}) + \sum_{d=1}^4 \gamma_{ds_t} + \sum_{m=1}^q \beta_{ms_t} w_{t-m} + \varepsilon_{s_t} \quad 8$$

$$\varepsilon_{s_t} \sim \text{i.i.d. } N(0, \sigma_{s_t}^2) \quad t=1,2,\dots,T$$

where (y_t) is the quarterly trend deviation of Ecuador’s real GDP; (y_{t-j}) is the autoregressive term, whose coefficients can be assumed to be state-independent (α_k) or state-dependent (α_{ks_t}) of the latent variable s_t which in-

⁷ The error terms will be heterocedastic if ($\sigma_{s_t}^2$) differ between regimes.

⁸ The general model can be also expressed as follows for p=q:

$$(5) \quad y_t = \mu_{s_t} + y_{t-1} \alpha_{s_t} + w_t \beta_{s_t} + \sum_{k=1}^p \theta_{ks_t} (y_{t-k} - \mu_{s_{t-k}} - y_{t-2k} \alpha_{s_{t-2k}} - w_{t-k} \beta_{ks_{t-1}}) + \varepsilon_{s_t}$$

$$\varepsilon_{s_t} \sim \text{i.i.d. } N(0, \sigma_{s_t}^2) \quad t=1,2,\dots,T$$

The state-dependent AR terms (θ_{ks_t}), corresponds to the lagged value of the residuals, and represents a moving average process (the current value of y_t depends linearly on the current and past error terms). (tsmswitch.pdf (stata.com))

dicates the unobservable regimes (“expansion” and “contraction”). v_{s_t} is the state dependent intercept and the exogenous variable and γ_{ds_t} is a vector of seasonal dummy variables where $\gamma_{ds_t} = 1$ if t is in quarter d and 0 otherwise. (w_{t-m}) corresponds to the quarterly trend deviation of nominal WTI prices, where the coefficient (β_m) will initially be assumed to be state invariant. This assumption is subsequently relaxed and considered to be state-dependent (β_{ms_t}) , to check whether oil prices shocks have an asymmetric or non-linear effect on economic growth (i.e. they can have either a positive or negative effect on economic growth depending on the state of the economy). (y_t) follows a p -th and (w_{t-m}) follows a q -th order autoregressive process. (ε_{s_t}) corresponds to the normally distributed errors with zero mean and state-independent (σ^2) or state-dependent $(\sigma_{s_t}^2)$ variance. The number of lags (p, q) included for each variable will be determined using information criterion and likelihood ratio (LR) tests (Cognigni and Manera 2009). The optimal specification (with lower information criterion and LR ratio) will be presented in the tables below. Unlike Hamilton (1996), Raymond and Rich (1997) or Clements and Krolzig (2002), we will also test the effects of oil price declines on real GDP cycles, given its direct dependence on the oil-exporting Ecuadorian economy.

The estimation of the parameters was based on the resolution of the expectation maximisation (EM) algorithm developed by Dempster *et al.* (1977) to find maximum likelihood estimators in probabilistic models that depend on unobservable variables. In addition, the inference of the probability of occurrence of each regime was performed using nonlinear filters and smoothers proposed by Hamilton (1989).

Finally, we test for the presence of volatility clustering in the residuals of the selected MS-AR-X model by assuming that the variance of the error term follows an MS-ARCH or an MS-GARCH process. We check whether they differ across regimes due to the effect of the explanatory variable, the oil price shocks.

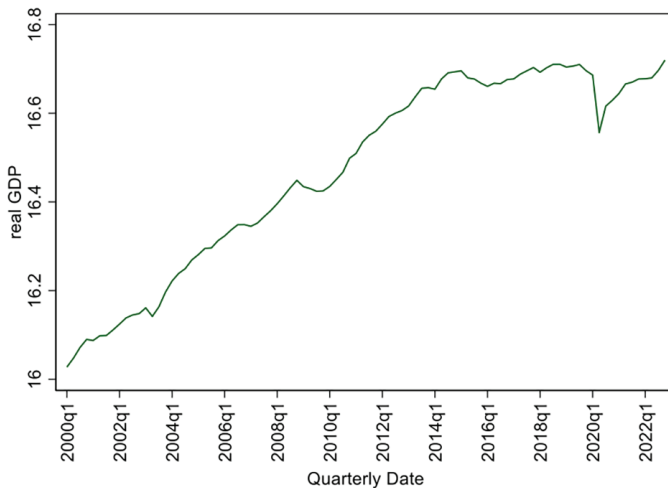
5. DATA SOURCES AND EMPIRICAL RESULTS

We examine two quarterly time series for the period 2000:01-2022:04⁹: (1) GDP in constant 2007 dollars, obtained from the Central Bank of Ecuador; (2) West Texas Intermediate (WTI) oil prices in nominal terms obtained from the FRED economic data on the St. Louis FED, which serve as a proxy for the price of Ecuadorian crude oil on international markets. Figure 1 displays the time paths of these series in levels and trend deviations.

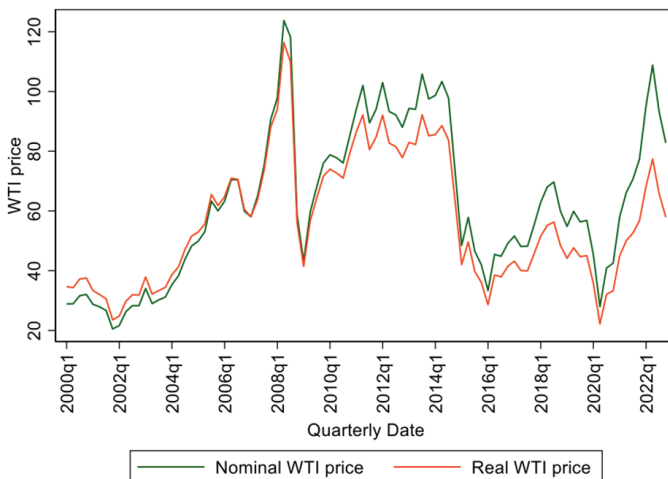
⁹ Central Bank of Ecuador has historical series of quarterly GDP from 2000:01 to 2022:04.

FIGURE 1
LEVELS AND DIFFERENCES OF QUARTERLY SERIES.
PERIOD 2000:01 TO 2022:04

A) Real GDP (US Dollar, 2007) (LOGS)

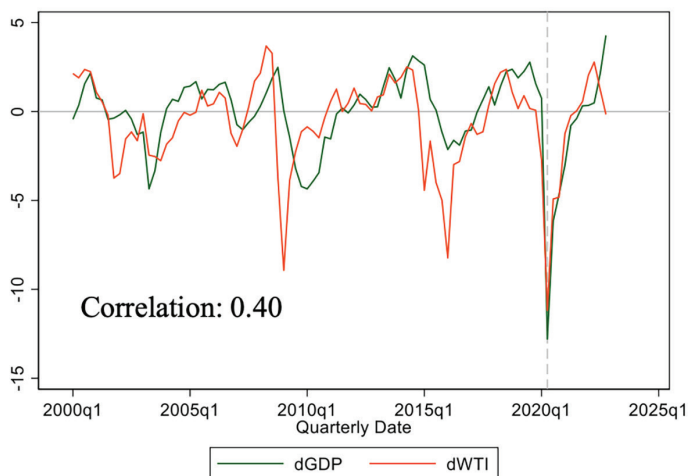


B) Nominal and Real (US Dollar, 2007) (LOGS)



The quarterly series of real GDP in logarithms (Figure 1a) are clearly non-stationary and could be characterized by a trend stationary process from 2000:01 to 2020:02. At this point Ecuador reaches its lowest real GDP growth

C) GDP and WTI Business Cycles (Trend Deviations)



Source: Central Bank of Ecuador for quarterly real GDP, FRED economic data from the St Louis FED for nominal and real oil prices.

of -12.8% (vertical dotted line) due to the Covid-19 pandemic, followed by a period of recovery. The quarterly series for nominal and real WTI in logarithms (Figure 1b) show a less clear pattern with significant fluctuations until 2015, when the price seems to stabilise. The nominal and real series also show similar trends, which start to diverge from 2009.01 onwards. We then present the trend deviation of both variables obtained after applying the Hodrick-Prescot filter, taking into account nominal oil prices. (Figure 1c). The series appear to be stationary with constant mean and variance. To guarantee stationarity of the underlying series, we conduct unit root tests (Augmented Dickey Fuller-ADF and Philip-Perron-PP). Table 1 confirms that the series are not stationary in levels but are stationary in trend deviations. Figure 1c also shows that both processes have a significant correlation (0.40), when the WTI price falls/rises, GDP also falls/rises, most often in the following quarter.

The next step is to examine the generating process of both series. The analysis of the Auto Correlation Function (ACF) and Partial Auto Correlation Function (PACF) indicate that trend deviations of the series are ARMA(2,2) (see Figure A1). Also the lower information criteria (AIC, AICC, BIC, HQIC) is found for two lags in both series (see Table A1). With regard to other statistics presented in Table A1, trend deviations for real GDP series (in logs) and nominal WTI prices appear to be normally distributed according to the Jarque-Bera Test. The Ljung-Box Q-test fails to reject the null hypothesis of a white noise, indicating that the series are not autocorrelated.

TABLE 1
UNIT ROOT TEST FOR STATIONARITY

			ADF		PP	
Variable			t-statistics	pvalue	t-statistics	pvalue
LnGDP	Level	Intercept	-2.193	0.2088	-2.434	0.1324
		Intercept, Trend	-0.305	0.9894	-0.694	0.9735
	Trend Deviation	Intercept	-3.486**	0.0083	-3.037*	0.0316
		Intercept, Trend	-3.444*	0.0458	-3.022*	0.0126
LnWTI	Level	Intercept	-2.037	0.2707	-2.076	0.2542
		Intercept, Trend	-1.966	0.6199	-2.116	0.5374
	Trend Deviation	Intercept	-3.907**	0.0020	-3.785**	0.0031
		Intercept, Trend	-3.880*	0.0129	-3.755*	0.0190

Notes: * p<0.05, ** p<0.01, *** p<0.001.

Next, we test the Markov–Switching (MS) autoregressive time series models. We considered two regimes as proposed by Hamilton (1989), Raymond and Rich (1997), Cologni and Manera (2009), where the economy can be in a contraction state of the business cycle, represented by $s_t = 1$, or in a phase of expansion, represented by $s_t = 2$.¹⁰ For both series, we applied a filtering process where factors such as seasonal patterns, outliers and trend, which may obscure the cyclical component of the series, are removed.

Table 2 presents the results of four MSAR models for the period 2000:01–2020:01 (quarterly series) in order to avoid the Covid-19 shock where the causes of the slowdown are linked to the pandemic and not so directly to changes in oil prices¹¹. The standard deviations of the estimators are shown in parentheses. All MS test have been carried out with the EM algorithm.

The first column (1) replicates the Hamilton (1989) univariate two-state Markov switching model for GDP, taking into account two lags in the autoregressive term (MSI(2)-AR(2) model), as suggested by the PACF and the information criterion. This estimation is used as a benchmark for the rest of the models. We observe that output has two clearly distinct growth regimes, state 1 being the slowdown regime and state 2 being the expansion regime, with an

¹⁰ For Estimations and post-estimations of the Markov switching AR regression, the STATA *mswitch* package was used.

¹¹ The fitted model would then have an inflated value of the variance for the stochastic level (Atkinson et al., 1997).

TABLE 2
EM ESTIMATORS FOR THE MARKOV REGIME-SWITCHING MODEL OF REAL GDP
FOR ECUADOR, 2000:01-2020:01

	(1)	(2)	(3)	(4)
STATE INVARIANT				
AR_{t-1}	1.421*** (3.89)		0.787*** (6.20)	
AR_{t-2}	-0.702***		-0.0526 (-0.42)	
w_{t-1}			0.138** (2.93)	
w_{t-2}			0.237*** (4.79)	
STATE 1 (Slowdown)				
Average growth rate in recession state(μ_{s_1})	-0.267 (-0.72)	-0.246 (0.353)	-2.695*** (-5.36)	-0.802*** (0.140)
Decreasing oil price dummy				
AR_{1t-1}		1.261*** (0.106)		-1.341*** (0.0474)
AR_{1t-2}		-0.567*** (0.104)		
w_{1t-1}				0.175** (0.0633)
w_{1t-2}				0.973*** (0.0509)
STATE 2 (Expansion)				
Average growth rate in expansion state(μ_{s_2})	0.9753** (2.18)	0.756* (0.412)	0.183 (0.61)	0.699*** (0.0882)
Decreasing oil price dummy				
AR_{2t-1}		1.894*** (0.289)		0.673*** (0.0783)
AR_{2t-2}		-1.123*** (0.253)		
w_{2t-1}				0.209*** (0.0415)
w_{2t-2}				0.316*** (0.0231)

Standard Deviation (σ)	0.8626	0.83176	0.6314	
Standard Deviation State 1 (σ_1)				0.0961
Standard Deviation State 2 (σ_2)				0.5746
p_{11}	0.8179	0.7877	0.6633	0.5851
p_{21}	0.0569	0.0693	0.0282	0.0579
Log-Likelihood	-115.85	-114.37	-86.81	-69.74
AIC	3.2761	3.2898	2.5664	2.2499
HQIC	3.3625	3.4008	2.7125	2.4676
SBIC	3.4924	3.5679	2.9317	2.7937
Duration s_1	1.060	4.710	2.970	2.410
Duration s_2	5.493	14.43	35.388	17.263
Jarque-Bera test on residuals	2.771	2.829	74***	3974***
Ljung-Box Q-test on residuals	106.1417 (0.0000)	103.0371 (0.0000)	35.7911 (0.4785)	9.9041 (1.0000)
Skewness on residuals	-0.3765	-0.3774	-1.1627	-5.0164
Kurtosis on residuals	3.565	3.579	7.202	36.266
Obs.	75	75	77	79

Notes: Seasonal dummies are not presented in the table. Standard errors in parentheses * p<0.05, ** p<0.01, *** p<0.001. i=0,1

average growth rate of -0.27% and 0.97% respectively. The dynamics of real GDP series are captured by the autoregressive coefficients (AR_i). They are considered to be state-independent and indicate that shocks have a significant inertia in the next quarter, followed by an opposite effect. This means that there is a tendency to restore equilibrium. The estimated probability of remaining in state 1 in the next period, is 0.81 while the estimated probability of transitioning to state 2 is 0.19 (1-0.81). On the other hand, the probability of remaining in state 2 in the next period is 0.94 and to transition to state 1 is 0.06 (1-0.94). This implies that both states are highly persistent, although slightly higher for expansionary states. Finally, the average duration of state 1 $\left[(1 - p_{11})^{-1} \right]$ is one quarter and for state 2 $\left[(1 - p_{22})^{-1} \right]$ is 5.5 quarters.

The second column (2) shows an univariate MSIA(2)-AR(2) model where intercepts, and AR terms are allowed to vary across regimes. The fit of the model is similar to previous one (similar LL, AIC, HQIC and SIC). The mean growth in state 1 is -0.24% while for state 2 it is 0.76%. The slope estimates represented by the first-order autoregressive (AR_{is}). coefficients also differ

across state, with higher values in expansions than in slowdowns, and with a positive impact in the first quarter followed by a shorter negative impact in the second quarter. The average duration of state 1 $\left[(1 - p_{11})^{-1} \right]$ is 1.14 quarters and that of state 2 $\left[(1 - p_{22})^{-1} \right]$ is 3.7 quarters, similar to those in the first model. In both cases, the residuals are normally distributed (we fail to reject the null hypothesis of the Jarque-Bera test), with skewness and kurtosis parameters close to zero. However, the Ljung-Box Q test for white noise in the residuals is rejected, suggesting that there is an autocorrelation problem in the error term.

We then include the nominal oil price shocks as an explanatory variable. The third column (3) presents the same specification as column (1) and includes two lags of nominal oil price deviations from trend as an exogenous state-invariant variable. This is the MSI(2)-AR(2)-X(2) model. The fit of the model improves with respect to the previous specifications (lower LL, AIC, HQIC and SIC). States 1 and 2 continue to represent the slowdown and expansion regimes, with an average growth rate of -2.7% and 0.18% respectively. The autoregressive coefficients (AR_t) have the same structural behaviour as in the previous cases, although only the first lag is statistically significant. With respect to nominal WTI price shocks, all coefficients are significant at 5%; an exogenous oil price shock has a direct effect on the business cycle of real GDP, which increases in the following quarter (0.14 and 0.23 respectively). The probability of remaining in the same state in the next quarter is higher in state 2 (0.97) than in state 1 (0.78). The average duration is also longer for state 2.

The fourth specification follows the structure of model (2) but includes oil price shocks as an exogenous state-dependent variable. After testing different specifications, the best fit was obtained with the model MSIAH(2)-AR(1)-X(2), where all parameters are allowed to be state-dependent. States 1 and 2 continue to represent contractionary and expansionary regimes (-0.8% and 0.7% respectively, both statistically significant). The autoregressive coefficients ($AR_{i,t}$) are significant at 99%, with a negative impact in state 1 (-1.3) and a positive impact in state 2 (0.7), implying that the lagged terms of y_t are better predictors in contractionary states than in expansionary states.

Oil price shocks also have a clearly asymmetric effect depending on the phase of the business cycle. In both states it has a positive and statistically significant effect on GDP, which increases in the second quarter after the shock. In expansionary states, this effect increases from 0.21 in the first quarter to 0.32 in the second one while in contractionary states the effect increases from 0.17 to 0.97. A 1% increase in the oil price has a cumulative positive effect on real GDP of 0.53% in slowdown states, while it reaches up to 1.14% over the following two quarters in expansionary states. The probabilities of remaining in

states 1 and 2 are similar to models (1) and (2). Finally, the regime-dependent standard deviation of the residuals (σ_i) is much higher in state 2. This shows that expansions are more volatile than recessions suggesting the presence of a conditional heteroskedasticity process in the error terms.

With regard of the rest of the estimates, the probability of remaining in the same state in the next quarter is also lower in contraction states than in expansion states (0.79 and 0.96 respectively). The average duration of expansions is still higher than that of contractions (22.7 versus 4.9 quarters). The regime dependent standard deviation of the residuals (σ_i) is slightly higher in state 2.

Looking at the rest of the statistical properties of the estimated residuals for the oil price models (3 and 4), none of them are normally distributed according to the Jarque-Bera test, and although the skewness is close to zero (models 4 and 5), the kurtosis is significantly positive. This result together with the state-dependent variance, indicates that the residuals are not i.i.d. and that there may be a conditional heteroscedasticity process. However, Ljung-Box Q-tests for white noise in the residuals reject the null hypothesis of white noise, suggesting that the residuals are not autocorrelated. The former condition is necessary for conditional heteroscedasticity in regime-dependent variance MSAR models (Krolzig 1997). Later we test for the presence of ARCH and GARCH processes in the error terms.

To confirm the results of the MSIAH(2)-AR(2)-X(2) model (4), we compare the predicted values with the actual values of real GDP growth (Figure 2). We find that both series are very similar, which means that the predicted values account for a large part of the variation in the dependent variable. Next, we analyse the probability that the output process is in state 1 compared with the official data. Figure 3 compares the smooth and filtered predicted probabilities (Kim 1994) of being in state 1 with the Ecuadorian business cycles calculated by the CBE (shaded areas).

The MSIAH(2)-AR(2)-X(2) model appears to correctly predict the probability of being in contractionary (1) and expansionary (0) states in most periods. It also confirms that Ecuador is more likely to be in expansionary, durable and recurrent states than in contractionary states, in line with Balcilar *et al.* (2017) for South Africa. This result differs from those observed by Neftci (1982), Hamilton (1989) or Raymond and Rich (1997), who found for developed oil importing countries that growth periods are less durable and recurrent than recessionary periods. However, since 2013:04, the frequency of slowdowns has increased significantly in Ecuador.

We also include an impulse response (IRF) analysis under the linear VAR model, considering the whole period, and then we emulate an MS-IRF test using the series below trend for state 1 and the series above trend for state 2. Comparing the results for the aggregate model (VAR) with those for state 1

FIGURE 2
 ACTUAL AND FITTED VALUES OF GDP TREND DEVIATIONS OF THE ECUADORIAN ECONOMY (2000:01-2020:1)

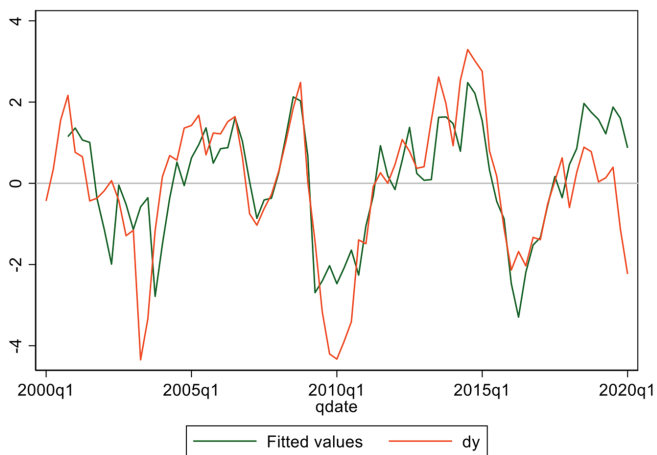
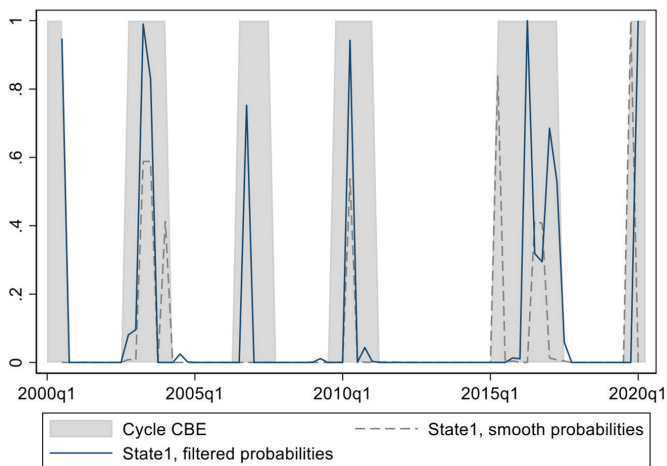


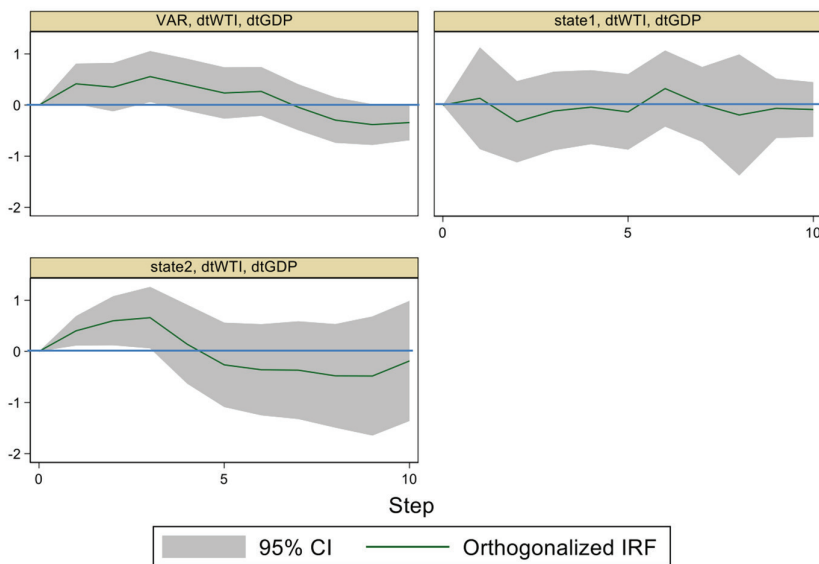
FIGURE 3
 STATES OF THE ECUADORIAN ECONOMY OBTAINED WITH THE MARKOV REGIME SHIFT MODEL (2000:01-2020:01)



Notes: The shaded areas represent the periods of deceleration below trend calculated by the Central Bank of Ecuador. Filtered and smooth probabilities estimate the state in each period using previous and contemporaneous data in the first case and the smoothing algorithm in the second.

and 2, we observe the asymmetries in the impact of oil price shocks on the business cycle in Ecuador. In expansionary phases of the cycle (state 2), there is a higher positive effect in the first four quarters and then a negative effect in the following quarters. In contractionary phases (state 1), the effect is more discrete, with alternating periods of positive and negative impulses. The forecasting error variance decomposition (FEVD) gives us the total contribution of oil price shocks in explaining the forecast uncertainty of real GDP. In the case of the linear VAR model, we obtain a cumulative effect of 0.18% over 10 quarters. After decomposing the series, we find a deeper impact in expansionary states (0.64%) than in contractionary states (0.034%), which also confirms the asymmetric behaviour of oil price shocks over the GDP cycle.

FIGURE 4
 IMPULSE RESPONSE OF REAL GDP TREND DEVIATION TO OIL PRICE SHOCKS IN
 LINEAR VAR AND MS-VAR MODELS (STATES 1 AND 2) (10 LAGS)



Graphs by irfname, impulse variable, and response variable.

To conclude this first exercise, we will analyse the conditional dependence of the error term obtained from our model (4). The residuals of the MSI-AH(2)-AR(2)-X(2) model were fitted with an MS-GARCH (1,1) model for

each regime period¹², which allows us to estimate the conditional mean and variance parameters (Table 3).

TABLE 3
CONDITIONAL DEPENDENCE OF THE ERROR TERMS.
FITTED PARAMETERS

	Estimate	Std. Error	t value	Pr(> t)
α_{01}	2.7117	0.6268	4.3259	0.000
α_{11}	0.4236	0.0348	12.1608	0.000
β_1	0.3638	0.0159	22.9207	0.000
α_{02}	2.7119	0.6931	3.9126	0.000
α_{12}	0.4236	0.0362	11.7151	0.000
β_2	0.3638	0.0164	22.1326	<0.000

All coefficients are statistically significant, indicating that there are conditional variance effects. The conditional mean parameters (α_{0k}) and the conditional variance ARCH (α_{1k}) and GARCH parameters (β_k) are similar in the two regimes, implying that there are no statistically significant asymmetries across regimes with respect to the volatility process of the error terms. The conditional mean volatility is much higher than the conditional variance volatility, indicating that the mean of the residuals is highly volatile over the period analysed and that the variability around this mean is relatively stable. The ARCH term (α_{1k}) captures recent volatility via the squared residuals while (β_k) captures past volatility, through the lagged conditional variance. This second estimate is a bit higher. The persistence of volatility is also the same in both regimes ($\alpha_{1k} + \beta_k = 0.78$), showing that the positive oil shocks cause more volatility and vice versa in both regimes. However, further analysis should be carried out in this respect.

This analysis would be incomplete if we did not examine the indirect transmission channels of oil price shocks into GDP by recognising the deep interdependence between aggregate demand and aggregate supply. For simplicity, we

¹² We used *msgarch* package from *RStudio*.

focus on the impact of oil price changes on the main components of aggregate demand. We estimate the MSIAH(2)-AR(2)-X(2) (model 4), where the trend deviation of each component of demand is the dependent variable. In all cases, state 1 represents periods of contraction, while state 2 represents periods of expansion in demand aggregates (Table 4).

Nominal WTI price shocks show regime asymmetries in all components of aggregate demand. The effects are positive and larger in expansionary states (state 2). In the case of investment, an oil price shock has a strong positive and statistically significant effect in expansionary states (of investment) in the next quarter (3.157), followed by smaller effects in the following quarter (0.31). The cumulative effect over the two quarters is 3.4% for each 1% increase in oil prices. In the case of investment slowdowns, the effect of oil price shocks is smaller (a cumulative effect of 1.07%). These results demonstrate the procyclical behaviour of Ecuadorian agents' investment decisions to international oil price shocks; when they rise, not only does the public sector have more revenue to invest, but the expectations of the private sector are higher, encouraging it to invest more. The opposite happens when oil prices fall, as the agents anticipate a crowding-out effect due to an increase in government borrowing to finance the public budget. This means that a rise in oil prices helps to restore the growth path of investment, but the opposite happens when oil prices fall, exposing the economy to a deeper cycle.

TABLE 4

EM ESTIMATORS FOR THE MARKOV REGIME-SWITCHING MODEL OF AGGREGATE DEMAND COMPONENTS FOR ECUADOR, 2000:01-2020:01

	Investment	Openness	Public expenditure	Private consumption
w_{t-1s_1}	0.574*** (0.170)	0.843*** (0.0442)	0.234 (0.152)	-0.0606 (0.101)
w_{t-2s_1}	0.473*** (0.107)	0.0699* (0.0371)	0.877*** (0.113)	0.300** (0.151)
Accumulated	1.047	0.9129	0.877	0.300
w_{t-1s_2}	3.157*** (0.109)	0.190** (0.0900)	0.509*** (0.100)	0.403*** (0.148)
w_{t-2s_2}	0.308* (0.176)	0.989*** (0.168)	0.237** (0.0779)	0.202** (0.0992)
Accumulated	3.465	1.179	0.746	0.605

Notes: Only the w_{t-m} effect is shown.

For openness, measured as the simple average of imports and exports, oil price shocks have a positive effect in both regimes. During cyclical slowdowns, the main effect is observed in the following quarter (0.84) and then declines to 0.07, with an aggregate effect of 1.18. In expansionary states, the main effect is observed in the second quarter with an aggregate effect of 0.91. For public expenditure, oil price shocks have a positive effect in both states (0.35 and 0.75 respectively). It increases in the second quarter in contractionary states, and decreases in expansionary states, confirming the pro-cyclical response of public policy to oil price shocks in expansionary states and the counter-cyclical response in slowing states. Private consumption seems to behave similarly, being positively affected by oil prices in both states.

6. CONCLUSIONS

This paper analyses the asymmetric effect of oil price shock on the business cycle of Ecuador, a highly oil-dependent and oil-exporting developing country since 1970. We applied a Markov switching autoregressive (MS-AR) regime model with two states, slowdown and expansion. We used two quarterly time series, real gross domestic product and the international price of WTI in nominal terms, during an observation period from 2000:01 to 2020:01, in order to avoid the effects of the COVID-19 pandemic on both series. Contrary to mainstream research that uses economic growth as a proxy for the business cycle, we use deviations from linear trend based on the methodology of the Central Bank of Ecuador (CBE) business cycle indicators, because although first differencing filters eliminate the trend component, they exacerbate the effect of high-frequency noise (Stock and Watson 1999). It also allows us to compare the results of our model with the Ecuadorian business cycles calculated by the CBE.

The oil price shocks are included as an exogenous variable (state-independent and state-dependent) in the Markov regime switching model with two lags. We find that exogenous oil price shocks have an asymmetric effect on Ecuador's business cycles: they have a more positive and persistent effect in expansions than in contractions; since GDP is in a slowdown process in contractionary states, an increase in oil prices would have a dampening effect but the opposite would occur when oil prices fall.

Using regime-dependent IRFs, we find that the cumulative impact of oil price shocks on real output is higher during expansions than in linear VAR models, and the opposite is true during slowdowns. The high aggregate impulse found between the two variables in expansionary states after 10 quarters (0.64) shows that Ecuador's economic specialisation in oil extraction has

helped the country to generate further expansion thanks to oil price increases, but neither have oil price falls been determinants of contractions, as the aggregate impulse found in contractionary states is very low (0.034), perhaps as a result of the oil price controls that exist in the country.

We also observe that oil price volatility plays an important role in determining the volatility of GDP growth. However, we do not find asymmetries in the conditional variance of the error terms across regimes. Conditional mean volatility is higher than conditional variance volatility, the latter being similarly driven by recent volatility (via the squared residuals obtained from the ARCH model) and past volatility (via the lagged conditional variance obtained from the GARCH model). As in Abdulkareem and Abdulkareem (2016) for Nigeria, the residuals show important persistence, with positive oil shocks causing more volatility and vice versa in both regimes.

The propagation mechanisms of oil price shocks on output have also been analysed through the components of aggregate demand (investment, private consumption, public spending and trade openness). We find that oil price shocks have significant and differentiated effects on these aggregates, demonstrating their indirect relationship with the business cycle and the importance of including them for a better understanding of the long-term evolution of the Ecuadorian economy. Oil price shocks have a strong effect in the same direction on the investment rate, which is higher in expansionary periods, demonstrating the procyclical behaviour of both variables. The same is true for the other components of demand, except for public spending, where the effect is higher in slowdown periods. This illustrates the complexity of the transmission mechanisms and the importance of a more detailed analysis of these variables.

These results suggest that Ecuador has a clear link with its natural resource specialisation, as Ocampo (2017) finds for South American countries. Exogenous (for the Ecuadorian economy) fluctuations in international oil prices have an important impact on its business cycle, especially in the case of slowdowns, exacerbating the fluctuations. This is a clear signal that the country should continue to seek new sources of income not linked to oil production and export. Public policies should strengthen and prolong the growth phases of the economy by stimulating private investment, reducing interest rates or raising total factor productivity, in order to insulate it from negative oil price shocks. However, oil revenues could continue to play an important role, especially in contractionary growth regimes, where the country has shown a weakness in its fiscal policy to stimulate growth through public spending (consumption and investment). So far, oil revenues have been used by the government to support economic growth through pro-cyclical and short-sighted fiscal policies, which have tended to exacerbate economic cycles. Instead, oil export revenues should be used to create stabilisation funds that allow fiscal policy to be counter-cyclical.

Further analysis should be done in relation to the conditional variance found in the MS-GARCH models. We could also allow for time-varying probabilities using endogenous switching models along the lines of Chang et al. (2017), Bazzi et al. (2017), Benigno et al. (2020) or Hubrich and Waggoner (2021), where the probability of switching regimes may vary over time depending on the state of the economy. It would also be interesting to investigate the out-of-sample forecasting ability of the models.

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APPENDIX A1

FIGURE A1

ACF AND PACF OF THE ARIMA PROCESS FOR GDP TREND DEVIATIONS

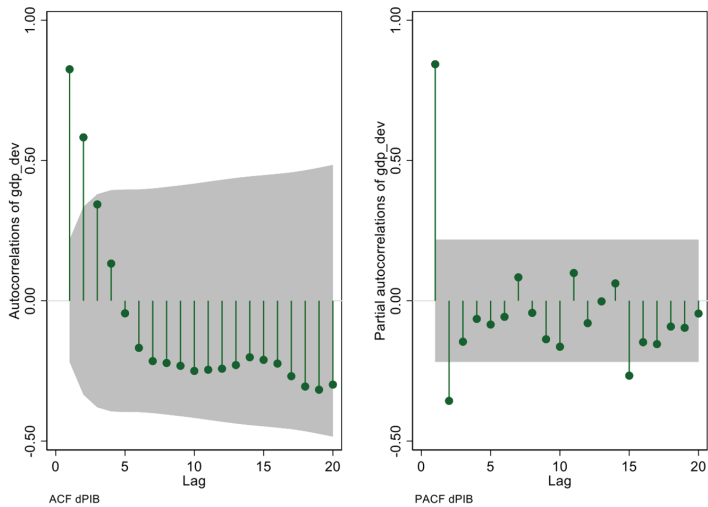


TABLE A1
LAG-ORDER SELECTION CRITERIA AND OTHER STATISTICS

	Lags	LR	AIC	HQIC	SBIC	J-B test	Skew	Kurtosis	Q-Test
dGDP	2	3.889	3.8901	3.9266	3.9814	5.828*	-0.633	3.351	261.92***
dWTI	2		-.3837	-.3720	-.354	34.66***	-1.180	5.166	194.79***

Notes: Likelihood ratio (LR), final prediction error (FPE), Akaike's information criterion (AIC), Schwarz's Bayesian information criterion (SBIC), and the Hannan and Quinn information criterion (HQIC). Only presented the Optimal lag. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Jarque-Bera normality test: Ho: normality. Ljung-Box Q-test: Ho: white noise (absence of autocorrelation).