

Oil Price Shocks and Japanese Macroeconomic Developments

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Abstract

This paper shows that oil price shocks lead to a fall in industrial production and higher inflation using quarterly data for Japan over the period 1976:I – 2008:II. We find evidence of non-linear effects of oil prices on both industrial output and inflation. Our modelling strategy incorporates information about structural breaks in the variables included to represent the macroeconomic transmission channels. Historical decompositions reveal that oil shocks contributed to both lowering industrial activity and raising inflation in the late 1970s and early 1980s. In more recent episodes of oil price increases, inflationary effects are little visible and there is very limited evidence of oil-induced industrial slowdowns.

Keywords: oil price shock; Japan; non-linear specifications

JEL classification: E32, Q43

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1. Introduction

A vast body of work has developed over the last three decades on how oil shocks affect macroeconomic developments. From a theoretical point of view, hikes in oil prices are seen to induce supply-side consequences including higher inflation and lower real output. Furthermore, terms-of-trade effects are expected to support aggregate demand in oil exporting countries, and lower it in oil importing countries. The empirical analyses have broadly corroborated these predictions.¹ After an early stage of studies involving linear models (see, *e.g.*, Rasche and Tatom, 1981; Hamilton, 1983; Burbidge and Harrison, 1984; among others), the trend has in recent years shifted to the use of non-linear approaches which seem to be better suited to capture the actual macroeconomic consequences of oil shocks. This trend goes back to the seminal work of Mork (1989), Lee *et al.* (1995), and Hamilton (1996 and 2003), who applied their models to the US economy.²

In the case of Japan, the number of studies applying non-linear methods to investigate the effects of oil price shocks on economic activity remains scanty, despite the fact that this country is the world's second largest economy and a major oil consumer and importer.³ One early non-linear study is Mork *et al.* (1994), who documented the existence of asymmetry in the inverse relationship between oil prices

¹ For recent accounts of the empirical work on this subject, see Hamilton (2008) and Jiménez-Rodríguez and Sánchez (2005), as well as the references therein.

² Arguably the most widely accepted explanation for a non-linear response of real activity to oil prices is given by Lilien's (1982) so-called *dispersion hypothesis* and its variants (see, *e.g.*, Loungani, 1986, and Hamilton, 1988). According to this view, inter-sectoral frictions may lead to higher output losses when oil prices go up and smaller gains when oil prices diminish. An overview of different economic interpretations of non-linear models is provided by Brown and Yücel (2002).

³ Japan is at present the third most important oil consuming country globally (behind the United States and China), representing some 6% of the world's total annual oil consumption. Considering the whole period 1975-2007 – which roughly corresponds to the estimation period used in this paper – Japan was comfortably the world's second largest oil consumer, accounting on average for about 8% of global oil consumption. In light of the country's limited domestic production, Japan currently imports over 4 million barrels a day of oil, which makes it the world's second largest oil importer (behind the United States) with a share of some 10% of this commodity's overall international trade (see British Petroleum data at <http://www.bp.com/sectiongenericarticle.do?categoryId=9023771&contentId=7044470>). Finally, oil use covers approximately half of Japan's primary energy needs (see, *e.g.*, Zhang, 2008).

and Japanese economic activity within a multivariate VAR model. More recently, other related non-linear works on Japan are Lee *et al.* (2001) and Jiménez-Rodríguez and Sánchez (2005), who both use multiequational specifications to analyse the link between oil prices and the macroeconomy, and Cuñado and Pérez de Gracia (2005) and Zhang (2008), who consider bivariate models. Lee *et al.* (2001) study the role of monetary policy in the transmission of oil price shocks for the Japanese economy. Using monthly data from 1960:1 to 1996:5, they find that between 30 and 50 percent of the negative impact of oil price shocks on Japanese output is attributable to monetary tightening induced by oil price shocks. Jiménez-Rodríguez and Sánchez (2005) study the effects of oil price shocks on the real economic activity of Japan (among other OECD countries) employing quarterly data from 1972:I to 2001:IV. They find a negative association between oil prices and Japanese real GDP growth when a second-order vector autoregression is used. Cuñado and Pérez de Gracia (2005) use quarterly data from 1975:I to 2002:II to show that there are no cointegrating long-run relationships between oil prices and industrial production and between oil prices and CPI, with the impact of oil shocks on these variables thus being limited to the short run. Moreover, they find that oil price shocks Granger-cause both output growth rates and inflation rates. Zhang (2008) applies Hamilton's (2001) approach to investigate the relationship between oil price shock and Japanese industrial production growth using quarterly data from 1957:I to 2006:IV. He finds that the oil price changes and macroeconomic activity in Japan appear to be affected by a non-linear relationship.

The present paper extends the related literature on Japan in two ways. First, we uncover the occurrence of structural breaks in some of the model's macroeconomic variables, which is then taken into account for the econometric approach used,

including the analysis of (non)stationarity of the model's variables, as well as the choice of the sample period and the introduction of dummies to control for regime changes. The attention paid to structural shifts in the Japanese economy is justified by the consideration of a long period starting in 1970 and a large literature analysing economic developments since the 1990s. This literature points to the markedly reduced dynamism of the country's labour productivity, the dampening of inflation (and the latter's eventual replacement by deflationary pressures), the emergence of a liquidity trap and associated financial problems.⁴ These features of the Japanese economy are considered here to imply important modelling decisions that have a bearing on our analysis of oil price shocks' effects on consumer prices and economic activity. Second, we introduce a novel characterisation of the way oil prices contribute to Japanese macroeconomic developments by identifying episodes of "high" oil prices periods for which we compute historical decompositions.⁵ In order to achieve these goals, this paper considers a wide array of linear and non-linear models.

The rest of the paper is organised as follows. Section 2 describes the methodology and data used. Section 3 describes tests concerning the time series properties of the data. Section 4 presents the empirical results from our identified VAR models. Section 5 concludes.

⁴ The associated financial problems during the 1990s are the burst of the asset bubble in early 1990s, followed by a situation often characterised as a "credit crunch" and accompanied by a large volume of non-performing loans (see, *e.g.*, Woo, 2003, and Kuttner and Posen, 2005). The role of monetary and fiscal policy in Japan's stagnation and deflation is still a matter of debate. For recent analyses, see Hamada (2004), Kuttner and Posen (2004), and Ito and Mishkin (2006). Other studies have focused on the real side of the economy, *e.g.* pointing to a slowdown in labour productivity and the reduction in the workweek as factors contributing to lower real output growth (see Hayashi and Prescott, 2002, and Morana, 2004).

⁵ Budd and Dicks (1982) compare the effects of oil prices during OPEC I (1973-74) and OPEC II (1979-80), although their informal comparisons were not based on econometric results.

2. Methodology and Data

The main results of this study are obtained by estimating and identifying a vector autoregression model of order p , or simply, $VAR(p)$. The reduced form of $VAR(p)$ can be written as $Y_t = AX_t + \varepsilon_t$, where Y_t is an $(n \times 1)$ vector of endogenous variables, X_t is an $(np \times 1)$ vector grouping all lagged terms of Y_t up to order p , A is an $(n \times np)$ rectangular matrix of coefficients, and ε_t is the $(n \times 1)$ generalisation of a white noise process with variance-covariance matrix Ω . The suitable lag length for the VAR is chosen on the basis of the Bayesian Information Criterion (BIC).

The vector of endogenous variables used here includes the following set of variables: industrial production, real oil price, consumer prices, real wage, real effective exchange rate (REER), and real short- and long-term interest rates. Some variables (industrial production, real oil price, consumer prices, real wage and REER) are expressed in logs, while the remaining ones are simply defined in levels. Oil prices, industrial production and consumer prices are the main variables of interest of this paper. The remaining variables in the model are added in order to capture the most important transmission channels through which oil prices may affect economic activity indirectly, in part by inducing changes in economic policies. Those channels include a variety of demand- and supply-side effects of oil prices operating via exchange rates, financial variables, and the labour market.

We use the following data in the present study. Our measure of real oil prices is defined as the ratio of the price of the internationally traded variety of crude most relevant to Japan (Dubai) converted into yen and then deflated by the country's CPI (all underlying data from IMF's International Financial Statistics – henceforth IFS –

except for CPI, from OECD's Main Economic Indicators – henceforth MEI).⁶ Industrial production data come from IFS; interest rates from MEI; wages from IFS; and REER (based on CPI) from MEI.

Prior to setting up our VAR model, we look at the time series properties of the data employed, as described in Section 3. Then, in Section 4 we identify the VAR model by means of a Cholesky decomposition, which amounts to using exclusion restrictions on the contemporaneous impact of the structural shocks.⁷ More specifically, we use the following recursive ordering for the variables in the system: industrial production, real oil price, consumer prices, real short-term interest rate, real long-term interest rate, real wage, and REER. This ordering presupposes that industrial output does not contemporaneously react on impact to the rest of the variables in the system. Oil prices are allowed to be immediately affected by unpredictable industrial activity developments,⁸ while being modelled as in turn having an immediate impact on wages, consumer prices, interest rates and the exchange rate.

VAR models are estimated for both a linear specification of oil prices and the four leading non-linear approaches considered in the literature. The latter are the following: i) Mork's (1989) asymmetric model; ii) Lee *et al.*'s (1995) scaled model; iii) Hamilton's (1996) net model; and iv) Hamilton's (2003) net3 model. The asymmetric specification allocates positive realisations of the rate of change in the oil

⁶ Oil prices are used directly in the linear approach to VAR estimation, and are transformed - in ways discussed below - for their use in non-linear specifications.

⁷ Formally, the structural form of the model can be expressed as $BY_t = CX_t + v_t$, where $v_t \equiv B\varepsilon_t$ is the vector of so-called structural form errors, B is an $(n \times n)$ matrix and $C \equiv BA$ is an $(n \times np)$ matrix. Under this scheme, B is assumed to be a lower triangular matrix with unit coefficients along the principal diagonal. The triangular factorisation of the positive definite symmetric matrix Ω is given by $\Omega = B^{-1}(B^{-1})'$. Without loss of generality, we assume that the variance-covariance matrix of the structural shocks is equal to the identity matrix.

⁸ Permitting Japanese industrial activity to influence oil prices is in line with the country's status as a major global player.

price to variable o_t^+ , and the corresponding negative realisations to o_t^- . In the scaled approach, the relevant oil variable - standing for "scaled oil price increases" - is

$$SOPI_t = \max\left(0, \hat{e}_t / \sqrt{\hat{h}_t}\right),$$

where \hat{e}_t and \hat{h}_t are, respectively, the estimates of the error and the conditional variance of oil prices from a AR(4)-GARCH(1,1) representation. The net specification uses the "net oil price increase" variable, defined as the amount by which oil prices (in logs), p_t , exceed - if at all - the maximum value over the previous 4 quarters; that is, $NOPI_t = \max(0, p_t - \max\{p_{t-1}, p_{t-2}, p_{t-3}, p_{t-4}\})$.

Finally, the net3 model defines its "oil price increase" variable – which we label $NOPI3$ – in terms of the rise in oil prices over the last 12 quarters, *i.e.* three years. We similarly construct scaled, net and net3 oil price decrease variables, that is, $SOPD_t$, $NOPD_t$, and $NOPD3_t$, respectively. As we mention in the next section, none of these three are found to have a statistically significant macroeconomic impact.⁹

3. Time Series Properties of the Data

One difficulty tackled in this paper is that, in light of the long time period considered (spanning almost 4 decades, from 1970:I to 2008:II), we need to look not only at the order of integration of the time series but also the possibility of structural change. The information derived from structural break tests is taken into consideration when setting the unit root tests and, consequently, to define the relevant sample period and model specification for estimation purposes.

⁹ We make all specifications comparable to each other by scaling down the impulse responses to oil disturbances in the cases of $SOPI$ (dividing by the sample mean of the standard deviation $\sqrt{\hat{h}_t}$), $NOPI$ and $NOPI3$ (in both cases dividing by the average effective number of quarters over which each transformation of oil prices is computed).

3.1. Detecting Multiple Structural Breaks

The literature provides several techniques for testing and locating structural breaks in the intercept and trend (including the well-known Bai–Perron-type tests), but only a few are able to consider breaks in the variance (see Inclán and Tiao, 1994, McConnell and Pérez-Quirós, 2000, Wang and Zivot, 2000, and Herrera and Pesavento, 2005). The possible existence of several breaks in the time series considered leads us not to follow the McConnell and Pérez-Quirós’ (2000) methodology, which has been developed to detect only the existence of one break in volatility. Furthermore, the possible occurrence of breaks in intercept/trend and variance at the same time leads us not to use the Inclán and Tiao (1994) or Herrera and Pesavento (2005) methodologies. We instead opt for Wang and Zivot’s (2000) methodology, which allows us to detect multiple structural breaks in the intercept, trend and variance at the same time. In so doing, we are able to detect the existence of breaks, the number of breaks and to identify break dates. Once we establish whether there are breaks and their dating, we analyse the order of integration of the time series considering the structural breaks (if any) that are detected.

Wang and Zivot (2000) postulate that a time series y_t evolves according to the segmented deterministically trending and heteroskedastic autoregressive model

$$y_t = a_t + b_t t + \sum_{i=1}^p \varphi_i y_{t-i} + s_t u_t, \quad (1)$$

for $t = 1, 2, \dots, T$ where $u_t | \Lambda_t \sim iidN(0, 1)$ and Λ_t denotes the information set at time t .

They assume that parameters a_t , b_t and s_t are subject to $m < T$ structural changes, with m initially known and break dates k_1, k_2, \dots, k_m ($1 < k_1 < k_2 < \dots < k_m \leq T$), so that the observations can be separated into $m+1$ regimes. Let $k = (k_1, k_2, \dots, k_m)$

denote the vector of break dates. For each regime i ($i = 1, 2, \dots, m+1$), the parameters a_t , b_t and s_t are given by: $a_t = \alpha_i$, $b_t = \beta_i$, $s_t = \sigma_i \geq 0$ for $k_{i-1} \leq t < k_i$ with $k_0 = 1$ and $k_{m+1} = T + 1$.

Let I_A denote an indicator variable such that I_A is equal to one if the event A is true and zero otherwise. Then (1) can be rewritten as

$$y_t = \sum_{i=1}^{m+1} I_{\{k_{i-1} \leq t < k_i\}} (\alpha_i + \beta_i t) + \sum_{i=1}^p \phi_i y_{t-i} + s_t u_t \quad (2)$$

Given the assumption of normality of the errors u_t , Wang and Zivot (2000) obtain the likelihood function of (2). The estimation of the model is possible by using the Gibbs sampler. The number of breaks is determined on the basis of the BIC.

We employ the Wang and Zivot test to look at the time series behaviour of the macroeconomic variables entering our VAR models. Using 8 lags (*i.e.*, the specification minimising the BIC), we find that, in the case of short- and long-term interest rates in real terms, three breaks are estimated to occur, namely in 1975:IV, 1997:II and 1997:III. As with Égert *et al.* (2006), we use the notion that, if two breaks are obtained in consecutive quarters, they are considered as one single break.¹⁰ Therefore, the breaks in the end considered short- and long-term real interest rates are for 1975:IV and 1997:II. Other variables subjected to structural breaks according to the Wang and Zivot test are consumer prices in 1973:I (with the optimal univariate specification being one of 8 lags), real wages in 1975:III and 1987:IV (with the optimal univariate specification corresponding to 8 lags), and REER in 1977:IV and 2004:I (with the optimal univariate specification being one of 4 lags). The results for

¹⁰ Égert *et al.* (2006) consider monthly data and use the idea that, if two or more breaks are obtained within an interval of 2 consecutive quarters, the set of such breaks is considered as one single break with the interim period being a period of adjustment.

industrial production indicate the no existence of breaks with the optimal univariate specification corresponding to 4 lags.

Having thus established the presence of structural breaks (or lack thereof) in macroeconomic variables, we turn to the analysis of the order of integration of the series in subsection 3.2, followed by the more substantive model results in section 4. The detection of structural breaks is used in three different ways. First, in light of the presence of a number of breaks at around the mid-1970s, we opt for defining a final sample period of 1976:I-2008:II for model estimation purposes. This has the advantage of avoiding the first of two breaks in three macroeconomic variables, namely, short- and long-term real interest rates, and real wages. In the case of REER, the two breaks detected occur after 1976:I. Second, we identify four possible periods of regime change on the basis of the break dates for (at least) one of the four variables presenting structural breaks since 1976:I. Associated with these four periods, we set up four dummies related to breaks in real interest rates (1997:II), REER (1977:IV and 2004:I) and real wages (1987:IV). As we shall see in section 4, these four dummies are allowed to enter the VAR model alternatively as step or impulse dummies – the reason being that the Wang and Zivot (2000) test does not specify whether the regime shift impinges solely on the intercept or the trend of the variables involved. Third, depending on the number of breaks (0,1,2) detected since 1970 and since 1976, we shall use different types of diagnostic tests for the (non)stationarity of our model's different variables – something we turn to in the next subsection.

3.2. Unit Root Tests

As we have just said, the tests used to evaluate the (non)stationarity of our model's different variables take into account the number of breaks exhibited by them.

The two main types of tests used to assess the order of integration of time series are “conventional” unit root tests (for variables displaying no structural breaks) and unit root tests that factor in possible structural breaks. Concerning the latter type of unit root tests, the literature in turn provides several tests of which we employ two here. First, we resort to the unit root test developed by Busetti and Taylor (2003), which has the advantage of allowing for a shift in both intercept or/and slope and variance, and is suitable for those variables undergoing only one break. For those variables with two structural breaks over the full sample (since 1970), we use the test developed by Lee and Strazicich (2003), which allows for two breaks in the intercept and trend.¹¹

Let us turn to evaluate the order of integration of the variables that display structural breaks both since 1970 and since 1976, namely, short- and long-term real interest rates, real wages and REER, leaving for later the application of “conventional” unit root tests to the remaining variables. Starting with short- and long-term real interest rates, the Wang and Zivot (2000) test showed that these variables display two breaks for the full sample period and one break during the final sample period starting in 1976:I. We thus apply the Lee and Strazicich (2003) test to the full sample period and the Busetti and Taylor (2003) for the final sample period. The Lee and Strazicich (2003) unit root test implemented for the full sample indicates that we cannot reject the presence of a unit root in the levels of real interest rates at the 5% significance level (see Table 1, Panel A). The Busetti and Taylor (2003) test¹² implemented for the sample starting in 1976:I reveals the existence of a unit root in the levels of real interest rates at the 5% significance level (see Table 1, Panel B).

[Insert Table 1 around here]

¹¹ Other tests allowing for structural breaks in the intercept or/and slope are Perron (1989, 1997), Zivot and Andrews (1992) and Lumsdaine and Papell (1997).

¹² We thank the authors for graciously sharing their code.

The unit root test results in Table 1 also point to the nonstationarity at the 5% level for the other two variables displaying structural breaks, namely, real wages and REER.¹³ The evidence from the Buseti and Taylor (2003) test for real wages is mixed, rejecting stationarity for 4 lags but not rejecting stationarity for 8 lags (see Table 1, Panel B). Given that the Lee and Strazicich (2003) test unambiguously fails to reject that real wages are nonstationary (see Table 1, Panel A) we conclude that this variable appears to be I(1).

Having established the (non)stationarity of the variables subjected to structural breaks, we finally turn to assess the order of integration for the remaining variables included in our model. These are real oil prices and the two macroeconomic variables (namely, consumer prices and industrial production), which appear not to display breaks from 1976:I onwards. Table 2 summarises the results obtained from a series of conventional unit root tests for these three variables. More concretely, there we present results from the *DFGLS* and P_T tests of Elliott *et al.* (1996), and the *DFGLS_u* and Q_T tests of Elliott (1999), as well as the Augmented Dickey-Fuller (ADF) test. According to these tests, the growth rates of the real oil price, industrial production and consumer prices appear to be all stationary.

[Insert Table 2 around here]

On the basis of all unit root test results (with and without structural breaks), we define the VAR models to be estimated to have in the vector of endogenous variables the first log-differences of four variables (industrial production, REER, real oil price, and real wage), the first differences of short- and long-term real interest rates, along with inflation rates in levels.

¹³ It is worth noting that, in light of the break in REER detected for 1977:IV, we report the Buseti and Taylor (2003) test only for the period since 1978:I (that is, two years after the final sample period starts), given that this test only allows one break,

4. Vector autoregression results

This section reports all the VAR results of this paper. Subsection 4.1 reports a number of preliminary tests for significance and model selection. Subsection 4.2 describes our results for accumulated responses of industrial output growth and inflation to oil price shocks. Subsection 4.3 reports historical decompositions which are analysed with a focus on the impact of oil shocks across dated “high” oil price periods. In all cases, we describe results concentrating on the specific type of (linear or non-linear) model that is found to perform best.

It is worth clarifying that, for the best-performing (linear or non-linear) econometric model, we present impulse response and historical decomposition results for two versions, namely: a) a baseline specification, in which the four structural change dummies referred to above¹⁴ enter as step dummies (i.e. with each adopting values equal to 1 from the reference quarter on, and being equal to 0 otherwise); and b) an alternative specification, in which the four structural change dummies in question enter as impulse dummies (i.e. with each adopting values equal to 1 only on the reference quarter, and being equal to 0 otherwise).

4.1 Testing for significance and model selection

In this subsection we start by investigating the significance of the relationship between oil prices and the other variables of the model, considering linear and non-linear specifications. We test for the significance of the oil price variables under consideration for the VAR system as a whole, using the null hypothesis that all of the oil price coefficients are jointly zero in all equations of the system but its own equation (see Table 3). The likelihood ratio test is informative in that oil prices, in

¹⁴ As we have seen at the end of subsection 3.1, the four dummies in question relate to reference quarters 1977:IV, 1987:IV, 1997:II and 2004:I.

addition to their direct effect on industrial output and inflation, could well impact the latter two variables through the rest of the system. The likelihood ratio tests show that oil price increase variables are significant at the 5% significance level in all models, while oil price decrease variables are insignificant in all cases and, consequently, deleted from our model.

[Insert Table 3 about here]

For model selection purposes, we look at two selection criteria as given by the Akaike Information Criterion (AIC) and BIC. On the basis of these two criteria, we conclude that the best-performing model is in all cases the scaled (see Table 4).¹⁵

[Insert Table 4 about here]

4.2 Accumulated responses

This subsection contains a discussion of the impact of oil price shocks on industrial production and inflation. Table 5 reports accumulated responses of industrial output growth and inflation to a permanent 100% oil price shock (focusing on best-performing specification given by the so-called scaled model). We distinguish between baseline results (attempting to control for structural shifts using step dummies) and those obtained from the alternative specification (employing impulse regime change dummies).

[Insert Table 5 about here]

Table 5 reports the accumulated impulse responses, with the baseline model results being given in the first line and the second line reporting those obtained from the alternative specification. The results indicate that, throughout the three-year horizon considered, an increase in oil prices drives inflation up and industrial

¹⁵ Our sample is quarterly and runs from 1976:I to 2008:II; this implies an effective sample that starts in 1976:3 after excluding (an optimal number of) 2 lags in all cases. It is worth reminding that the optimal number of lags is determined using the BIC.

production down, as expected. For both specifications considered, the macroeconomic reactions to oil price shocks are found to build over the time horizon. The adverse effect on industrial production growth is broadly consistent between the baseline and alternative models, with the former pointing to a somewhat larger contraction from oil price hikes. According to the baseline specification, an exogenous 100% oil price hike leads to an industrial production loss of some 1.3% after the first year, 2.7% at the end of the second year and 3.3% at the end of the third year. The inflationary effect also induced by the oil shock is found to be considerably larger in the baseline than in the alternative specification, especially during the first two years before results show a smaller discrepancy over the third year. Focusing on the baseline model, an exogenous 100% rise in oil prices has an inflationary impact of 1.0% after the first year, 1.6% at the end of the second year and 1.7% at the end of the third year.

The result that the scaled model turns out to be our preferred approach suggests that it is important to consider not just whether oil prices increase or decline (and by how much), but also the environment in which the movements take place. An oil shock in a stable price environment is likely to have larger economic consequences than one in a volatile price environment. In this regard, the scaled model more specifically highlights the importance of controlling for the time-varying conditional variability of oil price shocks.

4.3 Historical decomposition analysis

Here we consider the economic impact of oil prices for periods of "high" oil prices. In this regard, we perform historical decompositions showing the contribution of oil prices to industrial output and inflation over time. To this aim we use a simple definition of "high" oil price periods that is meant to have heuristic value, and follows

the one given by Jiménez-Rodríguez and Sánchez (2009). According to this criterion, a period of high oil prices is one over which the real oil price lies above the average of the series since 1970. In using this definition, we allow any such period to include individual quarters exhibiting a temporary fall below the sample mean (i.e. both the quarter right before and after them in contrast lie above the sample mean).¹⁶ For the estimation period of our VAR models, we find four periods of high oil prices: HIGH1 in 1979:II-1981:IV, HIGH2 in 1990:III-1990:IV, HIGH3 in period 1999:IV-2001:III, and HIGH4 in 2002:II-2008:II. One possible objection to this approach concerns the use of the mean as a benchmark in our definition, despite our evidence that real oil prices are a nonstationary series – a feature that our VAR models tackle by employing various transformations of oil prices themselves. Having said that, our periodisation presents the advantage that it appears to conform to major events marking world oil markets.¹⁷

Table 6 assesses the role of oil prices in the determination of industrial output and inflation in terms of historical decompositions for our high oil price periods, considering both the baseline and alternative specifications (in Panels A and B, respectively). For each of the two variables we report the actual series and the contribution of oil price shocks to the forecast, both of them in percentage deviations from the base projection.¹⁸ We find that oil price shocks appear to play a visible, if moderate, role both in the industrial sector's downturn and higher consumer inflation during HIGH1, that is, in the late 1970s and early 1980s. This was a period characterised by OPEC's strong influence over world oil markets. In more recent

¹⁶ We found only one such exceptional quarter (1981:III in period HIGH1).

¹⁷ One useful chronology can be found on the US Energy Information Agency's website at <http://www.eia.doe.gov/emeu/cabs/chron.html>

¹⁸ Specifically, both series are computed as deviations from a 4-quarter-ahead projection. We also computed decompositions using different conditional forecasts, which tended to broadly corroborate our substantive results.

episodes of oil price increases, inflationary effects are little visible and there is very limited evidence of oil-induced industrial slowdowns. The only exceptions to the latter result are a small negative contribution of oil price rises to industrial production growth in HIGH3 and – for the alternative specification – in HIGH4. For HIGH2 (1990:III-1990:IV), the signs of the oil shock effects are the opposite of those expected, which could be rationalised in terms of the short-lived nature of such high oil price period.

[Insert Table 6 about here]

V. Conclusions

The present study analyses the role of oil price shocks in Japanese macroeconomic developments. The theory predicts that, in an oil importing economy like Japan, unexpected hikes in oil prices should lead to lower economic activity and higher inflation. The empirical findings concerning the effects of oil shocks on industrial output growth and inflation confirm the expected pattern. In order to derive these results, this paper examines the possibility of structural breaks affecting the Japanese economy. Considering data from 1970, we detect structural breaks in short- and long-term real interest rates, real wages and the real exchange rate, including in the mid-1970s and the second half of the 1990s. The former break makes sense in light of the long period considered here which spans almost four decades. The break in the second half of the 1990s can be rationalised in terms of the discussions in the literature about Japan's widespread economic transformation, involving a dramatic deceleration in the pace of labour productivity growth, the emergence of deflation and the presence of a liquidity trap and credit crunch-type features in the financial sector. The detection of structural breaks in variables that are key to macroeconomic

transmission has prompted us to adopt a suitable modelling strategy. In this regard, our econometric approach has incorporated information about structural breaks into the study of (non)stationarity of time series, the choice of the sample period and the introduction of dummies to control for regime changes.

Our main empirical results provide evidence of non-linear macroeconomic impacts stemming from oil prices. More specifically, the scaled model – one of the leading non-linear approaches – is here found to dominate all of its alternatives. The scaled model, by controlling for the time-varying conditional variability of oil prices, highlights the importance of considering not only the magnitude and direction of actual oil price changes, but also the context in which the latter occur. The same oil price movement will normally entail a larger macroeconomic impact in an environment of stable as opposed to volatile prices for the commodity.

Historical decompositions are used to compute the contribution of oil price shocks across high oil price periods. Our analysis shows that the impact of oil prices on industrial activity and consumer prices is detectable only in the second half of the 1970s and early 1980s – an era in which OPEC had considerable influence on world oil markets. Except for some limited evidence of an oil-induced industrial slowdown in the two high oil price episodes experienced since the late 1990s, industrial production growth and inflation have been largely unaffected by the largest oil price hikes of the last quarter of a century.

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Table 1: *Unit Root Tests with Structural Breaks*

Panel A: Two-break minimum LM unit root test, sample period 1970:II-2008:II				
	$k=4$	$k=8$	$k=4$	$k=8$
	Real short-term interest rate		Real short-term interest rate	
LM_t	-3.294	-2.805	-4.287	-4.284
	Real wage		REER	
LM_t	-2.656	-2.897	-2.589	-2.126
Panel B: Buseti-Taylor stationarity tests, sample period 1976:I-2008:II				
	m_4	m_8	m_4	m_8
	Real short-term interest rate		Real long-term interest rate	
S^{***}	0.471**	0.352**	0.404**	0.320*
	Real wage		REER [†]	
S^{***}	0.308**	0.213	0.350**	0.297**

Note: Panel A) The entries are the outcomes of the LM_t statistics (developed by Lee and Strazicich, 2003) with the two shifts in level and slope in 1975:IV and 1997:II for real interest rates, in 1975:III and 1987:IV for real wages, and in 1977:IV and 2004:I for REER. The null hypothesis of this test indicates the existence of a unit root test with two breaks $[I(I)]$ versus the alternative hypothesis of a stationary time series with two breaks $[I(0)]$. We consider two values for the lags used to estimate the LM statistics, namely, $k=4$ and $k=8$. One/two asterisks mean a p -value less than 5%/1%. The critical values are taken from Lee and Strazicich (2003).

Panel B) The entries are the outcomes of the S^{***} stochastic stationarity statistics (developed by Buseti and Taylor, 2003) with the breakpoint in 1997:II for real interest rates, in 1987:IV for real wages, and in 2004:I for REER. The null hypothesis of these tests is that the time series is stochastically stationary $[I(0)]$ versus the alternative hypothesis of the existence of a unit root $[I(I)]$, controlling for shifts in slope, level and variance. The statistics are computed using two values for lag-truncation parameter $m_x(n) = \text{integer}(x(n/100)^{1/4})$, where x denotes the lag used to estimate the statistics and n the sample size used to calculate the statistic (for further details, see Buseti and Taylor, 2003). More concretely, we employ two values for x , namely, $x=4$ and $x=8$, which yield – for given $n - m_4$ and m_8 , respectively. One/two asterisks mean a p -value less than 5%/1%. The critical values are taken from Nyblom and Harvey (2000).

[†] The sample starts in 1978:I for this case, since it is detected a break in 1977:IV and Buseti and Taylor (2003) only allows one break. Thus, the breakpoint considered to implement the S^{***} stochastic stationarity statistics is 2004:I.

Table 2: *Conventional unit root tests*

	IP		Real Oil Price (Dubai)		CPI	
	Levels	First log-diff.	Levels	First log-diff.	Levels	First log-diff.
Panel A: Model with constant and trend						
<i>ADF</i>	-2.17	-5.74**	-1.40	-5.71**	-1.92	-4.08**
<i>DFGLS</i>	-1.00	-4.53**	-1.72	-5.57**	-0.96	-1.18
<i>P_T</i>	35.45	2.25**	9.57	2.15**	52.78	37.48
<i>DFGLSu</i>	-1.63	-5.45**	-1.74	-5.44**	-1.03	-2.63
<i>Q_T</i>	13.29	0.78**	5.79	1.27**	23.32	9.69
Panel B: Model with constant						
<i>ADF</i>	-2.35	-5.53**	-0.44	-5.66**	-2.51	-3.66**
<i>DFGLS</i>	0.77	-2.53*	0.78	-5.46**	0.63	-0.20
<i>P_T</i>	115.4	2.56*	22.43	0.74**	147.1	50.27
<i>DFGLSu</i>	-1.77	-5.47**	-0.75	-5.62**	-0.65	-2.78*
<i>Q_T</i>	58.05	0.96**	14.61	1.23**	68.64	16.69
Panel C: Model without constant						
<i>ADF</i>	2.04	-5.02**	1.37	-5.45**	1.20	-3.37**
<i>DFGLS</i>	2.04	-5.02**	1.37	-5.45**	1.20	-3.37**

Note: IP stands for industrial production and CPI for consumer price index. The sample is 1976:I-2008:II for the variables in levels, and starts one quarter later for the variables in first differences. We use data-driven lag selection procedures for the Augmented Dickey-Fuller tests, taking 1.645 as the critical value used for significance of lagged terms and 4 as the maximum number of lags allowed in these procedures into account. The same number of lags is used in the other tests considered. We denote with one/two asterisks the rejection of the null hypothesis at a 5%/1% significance level.

Table 3: *Likelihood ratio test*

	Linear	Asymmetric		Scaled		Net		Net3	
	o_t	o_t^+	o_t^-	$SOPI_t$	$SOPD_t$	$NOPI_t$	$NOPD_t$	$NOPI3_t$	$NOPD3_t$
Baseline specification (step dummies)	0.3170	0.0006***	0.3822	0.0020***	0.2247	0.0229**	0.6561	0.0120**	0.3082
Alternative specification (impulse dummies)	0.2613	0.0008***	0.3797	0.0015***	0.1413	0.0100**	0.5100	0.0006***	0.1215

Note: The entries are test statistics for the null hypothesis that all oil price coefficients are jointly zero in all equations of the system but its own equation (*i.e.*, $D_2 = 0$ below). One/two/three asterisks denote a p -value from the asymptotic distribution below 10%/5%/1%.

The test statistic is constructed as follows. Let the p -th order VAR model be rewritten as follows:

$$y_{1t} = k_1 + D_1' x_{1t} + D_2' x_{2t} + D_3' x_{3t} + \varepsilon_{1t}$$

$$o_t = k_2 + C_1' x_{1t} + C_2' x_{2t} + C_3' x_{3t} + \varepsilon_{2t}$$

where y_{1t} is the vector of variables other than o_t , x_{1t} contains lags of y_{1t} , o_t represents the real oil price change, x_{2t} contains lags of o_t , and x_{3t} contains the dummies.

The statistic is as follows:

$$2 \times [\mathcal{L}(\theta_1) - \mathcal{L}(\theta_2)] \sim^a \chi^2(\text{rows}(y_{1t}) \times p)$$

where $\mathcal{L}(\theta_1)$ and $\mathcal{L}(\theta_2)$ denote the value of the log likelihood function of the unrestricted and restricted models, respectively.

Table 4: *Relative performance of the models*

		Linear	Asymmetric	Scaled	Net	Net3
Baseline specification (step dummies)	<i>AIC</i>	18.472	17.591	16.172	17.037	16.531
	<i>BIC</i>	21.435	20.555	19.136	20.000	19.495
Alternative specification (impulse dummies)	<i>AIC</i>	18.668	17.780	16.375	17.242	16.790
	<i>BIC</i>	21.632	20.744	19.338	20.205	19.754

Note: The results are based on a seven-variable VAR with dummies that excludes oil price decrease variables from all four non-linear models, namely, the asymmetric, scaled, net and net3 specifications.

Table 5: *Accumulated responses*

	<i>Industrial production growth</i>			<i>Inflation</i>		
	after 1 year	after 2 years	after 3 years	after 1 year	after 2 years	after 3 years
Baseline specification (step dummies)	-1.30	-2.72	-3.25	1.03	1.63	1.66
Alternative specification (impulse dummies)	-1.42	-2.15	-2.46	0.03	0.65	0.82

Note: The entries refer to the accumulated impulse responses (in percentages) attributed to a 100% oil price shock.

Table 6: *Macroeconomic developments and oil shock impacts in high oil price periods*

Panel A: Baseline specification (step dummies)				
	high oil price periods			
	HIGH1	HIGH2	HIGH3	HIGH4
	1979:2-1981:4	1990:3-1990:4	1999:4-2001:3	2002:2-2008:2
Actual industrial production growth	0.17	1.45	-0.63	0.49
Oil-induced	-0.07	0.09	-0.04	0.01
Actual consumer price inflation	0.58	0.44	-0.10	0.03
Oil-induced	0.08	-0.03	0.02	-0.01
Panel B: Alternative specification (impulse dummies)				
	high oil price periods			
	HIGH1	HIGH2	HIGH3	HIGH4
	1979:2-1981:4	1990:3-1990:4	1999:4-2001:3	2002:2-2008:2
Actual industrial production growth	0.54	0.98	-1.13	0.24
Oil-induced	-0.05	0.09	-0.02	-0.03
Actual consumer price inflation	0.72	0.36	-0.44	-0.21
Oil-induced	0.04	-0.02	0.00	-0.01

Note: For each variable, both actual values and the historical decomposition components attributed to oil shocks are computed subtracting a 4-quarter-ahead projection of the relevant series. All entries are reported as average annualised rates for each high oil price period.