WORKING PAPERS

Castellón (Spain)

UNIVERSITAT JAUME•I

The effect of economic growth and oil price variations on CO_2 emissions: Evidence from Spain (1874-2011)

Economics Department

Jacint Balaguer
Manuel Cantavella-Jordá

2014 / 22

The effect of economic growth and oil price variations on CO_2 emissions: Evidence from Spain (1874-2011)

Jacint Balaguer
Universitat Jaume I
Department of Economics & IEI
jacint.balaguer@eco.uji.es

Manuel Cantavella-Jordá
Universitat Jaume I
Department of Economics & IEI
mcantave@eco.uji.es

2014 / 22

Abstract

A structural analysis on an environmental Kuznets curve (EKC) with oil prices is carried out for Spain from 1874 to 2011. The dynamics of the long and short-term relationships among carbon dioxide (CO_2), economic growth and oil prices is captured through an autoregressive distributed lag (ARDL) model. The EKC hypothesis is supported in a context where real oil price increases play a significant role on the reduction of CO_2 emissions. Hence, the results suggest that oil taxes should be taken into account as an effective tool of environmental policy.

Keywords: Environmental Kuznets curve, CO_2, oil prices, autoregressive distributed lag model, time series

JEL classification: Q53, Q56

The Effect of Economic Growth and Oil Price Variations on CO_2 Emissions: Evidence from Spain (1874-2011)

Jacint Balaguer* and Manuel Cantavella-Jorda*
**Department of Economics and International Economics Institute*
Universitat Jaume I, Campus Riu Sec, 12071 Castelló, Spain.

Abstract

A structural analysis on an environmental Kuznets curve (EKC) with oil prices is carried out for Spain from 1874 to 2011. The dynamics of the long and short-term relationships among carbon dioxide (CO_2), economic growth and oil prices is captured through an autoregressive distributed lag (ARDL) model. The EKC hypothesis is supported in a context where real oil price increases play a significant role on the reduction of CO_2 emissions. Hence, the results suggest that oil taxes should be taken into account as an effective tool of environmental policy.

Keywords: Environmental Kuznets curve, CO_2, oil prices, autoregressive distributed lag model, time series.

1. **INTRODUCTION**

Over time, humankind has contributed to atmosphere contamination through different processes of production and consumption. The massive consumption of fossil fuels turned atmosphere pollution into a severe trouble. This started with the scientific and technical revolution back on the mid-eighteenth century. As a result, global warming and depletion of ozone layer has become two of the most important environmental problems today for humans and ecosystems. Their origin is therefore mostly human-induced (anthropogenic). Global warming is mainly caused by rising concentrations of carbon dioxide (CO_2) aside from other greenhouse gasses in the atmosphere.

From the seventies, when a major concern on the environment's deterioration took place at international level, some economists argued that enhancement in per capita income could eventually reduce the level of environment degradation. We can distinguish at least two arguments widely used to explain this possible phenomenon. That is, on the one hand, the variation towards a new productive structure which uses a lower level of polluting energies. On the other hand, the social willingness to pay for the extra cost associated to the cleaner energies. These underlying ideas were enthusiastically accepted when an early set of papers (e.g., Grossman and Krueger, 1991; Shafik and Bandyopadhyay, 1992; Panayotou, 1993) provided the first formal evidence about an inverted U-shape relationship between per capita income and environmental degradation following, therefore, a Kuznets type curve.[1] Nowadays, we can benefit from many studies that show the existence of an environmental Kuznets

[1]Kuznets (1955) had originally suggested that a changing relationship between per capita income and income inequality could be represented by an inverted-U-shaped curve.

curve (EKC) for several countries by using different sample periods, econometric specifications and empirical methods.

With the aim of contributing to the knowledge on the EKC for CO_2 emissions, this paper provides new evidence for Spain by exploiting time series from a large period (1874-2011). Three recent studies, using a similar span of time period to the one employed in this paper, explored the relationship between economic growth and pollution for the Spanish case (i.e., Esteve and Tamarit, 2012a, 2012b; Septhon and Mann, 2013). They all considered per capita CO_2 and per capita income. The different approach among them basically consisted of the functional forms assumed for bivariate models. Unlike those papers, we analyze the possible effect on Spanish CO_2 emissions generated not only from the economic growth but from the changes in real oil price as well.

The level of oil prices may presumably affects CO_2 emissions through two ways regardless of their indirect effect via GDP.[2] On the one hand, it is well known that an increase in oil prices could imply a reduction on energy consumption. This might be compensated, in order to sustain GDP levels, by using more units of either labor or capital. On the other hand, we must bear in mind that fuel combustion is, after coal, the most pollutant of all energy alternatives. Thus, higher oil prices may drive towards substitution of fuel combustion by other cleaner energy resources. Taking into account the potential effect of oil prices to reduce CO_2 emissions, its introduction in the model specification will allow us to know the degree to which taxation on oil products may be considered as a useful environmental policy. Obtaining more accurate estimates of the per capita income effect is another compelling reason to extent the simplest EKC

[2] The effect of oil prices on GDP is widely recognized, and has been received notable attention in the empirical literature (e.g., Lardic and Mignon, 2006; Bachmeier, et al. 2008; Kilian and Vigfusson, 2013).

specification. That is, if oil price is a relevant variable and it is correlated with GDP, introduction of oil prices will avoid the estimation bias on the per capita income effect.

The existing empirical literature gives us light on the consideration of oil prices in the EKC framework (i.e., Agras and Chapman, 1999; Heil and Selden, 2001; and Richmond and Kaufmann, 2006). These authors claim the importance of oil prices and indicate that measures oriented to increase domestic prices on the most polluting energies constitute a valuable tool to reduce the level of CO_2 emissions. Moreover, the results obtained by Richmond and Kaufmann (2006) from the US data suggested that including energy prices in the model could have a considerable impact on the estimated income coefficients. In fact, in the analyzed context, the inclusion of energy prices removed statistical support for typical turning points.

The estimation process that we use is the autoregressive distributed lag (ARDL) bounds testing procedure of Pesaran and Shin (1999) and Pesaran et al. (2001). A major advantage of this method is that allows us to make valid inferences on both parameters and functional forms regardless of whether the time series are *I(1)* or *I(0)*, or a combination of both. This advantage makes the method particularly suitable to our purpose. The reason is that different historical stages included in our long-time series imply presumable presence of structural breaks, which introduces uncertainty as to the true order of integration of the variables. This means that, it is possible that any of the variables used here are stationary around some probable structural breaks,[3] but can be erroneously classified as *I(1)* from conventional tests. Another noteworthy advantage is that the ARLD bounds testing approach is superior to that of the traditional Johansen's cointegration methodology, which in general, requires a very large sample size. In

[3]Moreover, the ARDL will allow us conveniently test whether or not there is underlying structural breaks that affect the long-run stability of estimated coefficients.

particular, Pesaran and Shin (1999) demonstrate that the ARDL procedure has better properties in a sample size as the used here (i.e., less than 150 observations).

The rest of the paper is organized as follows. Section 2 refers to the empirical background directly related to the approach used in this paper. In section 3, we present the most relevant historical facts that characterize the sample period and the subsequent implications on the evolution of the time series involved in our analysis. In section 4, we will introduce the econometric specification and the implemented methodology. Section 5 will report the empirical results and discuss them. Finally, section 6, will provide conclusions and some policy implications.

2. EMPIRICAL BACKGROUND

Since the initial empirical studies of the aforementioned economists in the early nineties, a large number of papers have tested the existence of an inverted U-shape relationship between pollution level and per capita income. There are several recent surveys on this topic offering a fairly comprehensive overview of the state of the question (e.g., Kijima, et al. 2010; Bo, 2011; Pasten and Figueroa, 2012). In this section we will only focus on those papers that are related in some way to either the context and/or the type of model specification which are object of our research. First, we briefly review those papers that provide evidence for Spain. Second, we concentrate on those papers that not only estimate income effects on CO_2 emissions but also explore the role of energy prices on the environmental quality.

The paper by Roca et al. (2001) is the first research that estimates a long-run relationship between income and CO_2 emissions for the Spanish case by using time series analysis. The empirical results, from a sample period that ranges from 1973 to 1996, do not reveal the existence of an EKC since the estimated elasticity between (per capita) income and (per capita) CO_2 emissions is positive and greater than one (1.24). Because the long-run relationship is assumed to be stable over time, this outcome has been recently questioned in Esteve and Tamarit (2012a). They introduced potential breaks in a bivariate model and considerably extended the data sample used by Roca et al. (2001). From a sample that goes from 1858 to 2007, the authors found that the long-run elasticity between (per capita) income and (per capita) CO_2 is defined by three regimes where estimates decreased over time. This outcome has been interpreted as the existence of a declining growth path pointing to a prospective turning point even though the EKC does not follow an inverted U-shaped curve.

Two additional papers are found in the literature using alternative functional forms. Those reexamined the relationship between (per capita) income and (per capita) CO_2 emissions for Spain through the same data sample utilized by Esteve and Tamarit (2012a). More specifically, Esteve and Tamarit (2012b) employ a bivariate model with two threshold regimes in order to combine the idea of cointegration with nonlinearity (in the adjustment) between income and CO_2 variables.[4]The paper does not provide information about a possible turning point as the standard EKC approach points out. However, their results suggested, once more, that economic development is compatible with pollution reduction. By adopting a more complex functional form, where the cointegration relationship between (per capita) income and (per capita) CO_2 is assumed non-linear,[5] such outcome is also obtained by Septhon and Mann (2013).

[4]The bootstrapping method applied to the bivariate model rejects the hypothesis of linear cointegration relationship in favour of threshold cointegration relationship.

[5]The justification of this complexity is that threshold cointegration per se does not imply non-linearity to the equilibrating relationship. Thus, the attractor itself could still be linear in Esteve and Tamarit (2012b).

All the papers described above pay no attention to the potential contribution of energy consumption with respect to the level of CO_2 emissions. It is obvious that the utilization of energy, especially combustion of fossil fuels, is, at the same time, a large source of pollution. Furthermore, disregarding the role of energy use in models may generate estimation bias if energy use and income are related.[6]Thus, the inclusion of an energy variable, which collects the possible effect of the energy consumption, seems to be a reasonable empirical strategy.

Due to the standard build procedure of CO_2 variable,[7]energy consumption is contemporaneously correlated with CO_2 emissions. Stemming from it, we can see as some authors have instrumented the energy consumption variable through energy prices. The effect of energy prices on CO_2 emissions (via energy use) is explored in the literature. To our knowledge, the paper by Agras and Chapman (1999) is the first to introduce energy prices in an econometric EKC framework. The research is done for a large set of high, middle and low-income countries between 1971 and 1989. The authors used gasoline prices given that combustion of fossil fuels is the main source of CO_2 pollution in the analyzed countries. Results derived from a partial adjustment model indicated that oil price is an essential variable to explain the level of CO_2 emissions. Moreover, it is revealed that their inclusion (together with other no relevant variables)[8] removes the empirical evidence in favor of a turning point for income. Therefore, it is

[6]Nowadays, a large number of studies reveal causality between energy consumption and economic growth for a large set of countries as can be seen from a review of the literature. See, for example, Payne (2010).

[7]The direct inclusion of fuel consumption in models may introduce endogeneity problems. In fact, the Carbon Dioxide Information Analysis Center (CDIAC) measures CO_2 by multiplying different primary energy sources (coal, petroleum, and natural gas) by their respective emission rates.

[8]Trade variables are also included but they are not revealed as significant.

highlighted that economic growth per se will not reduce carbon emissions. Thus, the omission of oil prices can lead to recommend wrong environmental policies.

Similar conclusion is obtained by Heil and Selden (2001), and Richmond and Kaufman (2006). The former use panel data information related to a large set of countries for the period 1951-1992. They forecast the cumulative emissions for the period 1991 to 2001and compare them with the period between 1881 and 1990. The cumulative emissions would be a seven-fold the preceding period without new active environmental policies. This indicates that it may have very adverse effects on climate change. The authors point out that the common oil price used in their analysis should be considered as a proxy of domestic energy prices. It obviously varies across the countries due to taxes, subsidies, and others idiosyncratic distortions of each region. A small magnitude of the coefficient associated to oil prices is found. However, the authors empathize that this should not be interpreted as evidence that energy price increases would have little impact on CO_2 emissions. Richmond and Kaufman (2006) use panel data information for several OECD countries related to 1978-1997 period. Once again, it is shown that there is no turning point when oil prices are included in the econometric model. Oil prices reveal to be quite effective for reducing energy use and CO_2 emissions. The authors explicitly claim that raising real energy prices may be considered an effective policy measure for environmental improvement.

The three studies discussed above employ relatively short time series. Nevertheless, they use a wide range of cross-sectional data. Although the results are restricted to be common to a set of countries, they give an interesting overview about the phenomenon we are concerned to. The economic arguments as well as the relevant omission bias suggested by the empirical results encouraged us to reexamine, through longer time series, the EKC for specific countries. The availability of a reasonable time span data for the Spanish case is an advantage. This case is particularly interesting given the different results and approaches showed by existing papers (e.g. Roca et al., 2001; Esteve and Tamarit, 2012a; Esteve and Tamarit, 2012b; Mann, 2013), some of them very recent. However, the role of energy prices and in particular real oil prices is not taken into account. We expect that our analysis in the same context can enrich the knowledge on the EKC. Furthermore, the use of oil prices and not energy consumption may avoid endogeneity problems as well as some potential biasness in the final results.

3. SAMPLE PERIOD

We focus on a sample period between1874 and 2011 for which annual time series for CO_2, GDP, population, and international oil prices are available. They are collected from different data sources. Thus, CO_2 emissions measured in metric tons were obtained from the Carbon Dioxide Information Analysis Center (CDIAC) of the US Department of Energy (available at the websitehttp://cdiac.ornl.gov/)The GDP, in US dollars (1990 base year), and the population were compiled by the Maddison Historical Statistics (http://www.ggdc.net/maddison/maddison-project/home.htm) and the *Instituto Nacional de Estadística* (http://www.ine.es/). Finally, crude oil prices measured in real dollars (2010 base year) were gathered from the Statistical Review of World Energy 2013 provided by the British Petroleum company (http://www.bp.com/).

Casual inspection of the three different time series in their original form can give us a good approximation to the historical stages as well as to the relevance of the studied phenomenon. As we can see in Figure 1, the overall evolution of CO_2 emissions seems a priori a bit worrying. Indeed, the emissions in 1950 were twenty four times those generated in 1874. Whereas in 2011, they have been about two hundred and fifty times higher than those emitted at the beginning of the shown period. Regardless of the concerns caused by the emissions trend, Spain may have reached a certain level of the per capita income from which some reducing effect on emissions would have been taken place. This is what the EKC hypothesis advocates. In fact, the per capita income in this country has also experienced a rather significant growth. Although in 1950 the per capita income was only about fifty percent greater than in 1874, the increase achieved in 2011 is more than eleven times that on the corresponding starting year.

[Figure 1]

The moderate economic growth at the beginning of the sample and the extraordinary development in the past five decades respond to certain historical facts that it is worth of briefly mentioning. In broadly terms, this moderate growth sub-period in Spain came associated to a predominantly agricultural economy, with low utilization of energy resources, and scarce external relations.[9]This economic growth was transitory interrupted by the Spanish Civil War (1936-1939). Under the following dictatorship government, a reconstruction programme was initiated in order to rebuild and expand the industrial base. The idea of state interventionism and autarky reflected those times. The development was somewhat conditioned by the access to limited domestic energy sources[10]. Therefore, the overall benefits of the development program turned out to be relatively modest. With the aim to boost economic growth, some technocrats advocated implementing a package of deep policy reforms introduced in 1959 and in the immediate following years. Thus, the country abandoned the underlying idea of

[9] The reasonable performance of the economy in Spain during this first stage is described, for example, in Molinas and Prados de la Escosura (1989), and in Fernández Navarrete (2005).

[10] A discussion on energy restrictions can be obtained from Sudrià (2010).

economic self-sufficient through the well-known *Stabilization and Liberalization Plan*. Consequently, this change encouraged trade in general and imports of petroleum products in particular.

As we can also see in Figure 1, international oil prices display some generalized growth and turbulences, which are especially relevant in the last decades. More specifically, oil prices rebounded from 1973 mainly due to the Yom Kippur War. Later on in 1978 they extraordinarily increased once more as a consequence of Iran revolution. The return to social stability in the next decade was accompanied by a falling in oil prices. They rose again but moderately this time when Iraq invaded Kuwait in 1990. After the Asian financial crisis of 1997, and mainly after Iraq's conflict and their invasion in 2003, the oil prices greatly augmented again. Later on, the renowned Arab spring in2010 ended a new stage of falling prices. Itseems quite reasonable to think that these set of changes along with others, have had some influence in both the Spanish oil demand and, therefore, in the evolution of their CO_2 emissions.

4. MODEL SPECIFICATION AND METHODOLOGY

Our analysis is based on an econometric model, in natural logarithms, that specifies per capita real income proxied by per capita GDP ($ln\ y_t$), its quadratic term, and real crude oil prices ($ln\ p_t$) as explanatory variables of the CO_2 emissions ($ln\ CO_2$):

$$\ln CO_{2_t} = \beta_0 + \beta_1 \ln y_t + \beta_2 (\ln y_t)^2 + \beta_3 \ln p_t + \varepsilon_t \qquad (1)$$

where ε_t is the error term which is assumed to be independent and normally distributed. According to the typical EKC form, we would expect that the elasticity of CO_2 with respect to per capita income be positive ($\beta_1>0$) and the income elasticity of its square would become negative ($\beta_2<0$). Oil price elasticity for CO_2 emissions would be expected to be negative ($\beta_3<0$), meaning that higher prices would discourage the use of energy and therefore emissions would be lower.

It is well recognized that working with time series may involve serious problems related to spurious regressions if variables are not stationary. Differencing variables that are non-stationary in levels solves the problem but, then, we lose long-run information. Since Engle and Granger (1987), different approaches have been proposed to also obtain long-run estimates.

The empirical methodology that we use in this paper is the Autoregressive Distributive Lag (ARDL) bounds test proposed by Pesaran et al.(1997, 2001). Then, the error correction model (ECM) can be easily derived from the ARDL framework making also possible to estimate the long-run adjustment process towards equilibrium. In a Montecarlo experiment, Haug (2002) proves good properties of this technique regarding cointegration. One of the advantages is that the time series regression can be carried out regardless of the nature of variables, that is, whether or not they are either I(1) or I(0). Given that most of the macroeconomic variables are proved to be either one of those two orders, then this methodology is convenient with the aim of examining long-run relationships. As Pesaran and Shin (1999) demonstrated, another great advantage is that serial correlation and endogeneity problems are removed when long-run and short-run components are simultaneously taken with appropriate lags.

The relationship among per capita CO_2, per capita income, and oil prices postulated in Eq. (1) follows a time path before a long-term nexus is achieved. Thus, the Eq. (1) would be written as an unrestricted error correction representation:

$$\Delta \ln CO_{2_t} = \alpha_0 + \sum_{i=1}^{p} \alpha_i \Delta \ln CO_{2_{t-i}} + \sum_{i=1}^{p} \varphi_i \Delta \ln y_{t-i} + \sum_{i=1}^{p} \gamma_i \Delta (\ln y_{t-i})^2 + \sum_{i=1}^{p} \delta_i \Delta \ln p_{t-i} + \\ + \lambda_1 \ln CO_{2_{t-1}} + \lambda_2 \ln y_{t-1} + \lambda_3 (\ln y_{t-1})^2 + \lambda_4 \ln p_{t-1} + \varepsilon_t \quad (2)$$

The estimation procedure used here involves two stages. In a first stage we will analyze, through the ARDL bounds test, whether or not there is evidence of a cointegrating relationship. With this purpose, the null hypothesis of no cointegration $(H_0 : \lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = 0)$ should be tested against the alternative hypothesis. Ordinary least squares report F-statistics which are compared to the critical values given in Pesaran and Shin (1996) and Pesaran et al. (2001). If they go beyond the upper bound then the null hypothesis will be rejected and there will be a cointegrating relationship among the variables. On the contrary, if the F-statistics are below the lower bound, the null hypothesis will not be rejected. In the case that the F-statistics are in between the upper and lower critical values, then the test result should be considered inconclusive. The second stage, then, is to estimate the long-run coefficients of the cointegrating relation and make inferences about their values.

Finally, the empirical methodology involves the modeling of a restricted error correction representation, which takes a similar form of Eq. (2), but now including the long-run terms in the error correction variable lagged one period:

$$\Delta \ln CO_{2_t} = \alpha_0 + \sum_{i=1}^{p} \beta_i \Delta \ln CO_{2_{t-i}} + \sum_{i=1}^{p} \varphi_i \Delta \ln y_{t-i} + \sum_{i=1}^{p} \gamma_i \Delta (\ln y_{t-i})^2 + \sum_{i=1}^{p} \delta_i \Delta \ln p_{t-i} + \lambda\, ect_{t-1} + \varepsilon_t \quad (3)$$

where ect_{t-1} is the error correction term represented by the OLS residuals series from the long-run cointegration relationship, and the λ coefficient indicates the speed of adjustment towards this long-run equilibrium. Diagnostic and stability tests will reveal the soundness of the model.

5. RESULTS AND DISCUSSION

Since the ARDL approach does not contemplate an order higher than I(1), unit root tests will still be performed. We are aware that any structural break in the variables would reduce the power of this type of tests. If that were the case, then I(1) could not be rejected but anyhow, the aforementioned method admits both degrees I(1) and I(0) or even if it is an I(0) plus trend (TSP). The Augmented Dickey Fuller (ADF) unit root tests presented in Table 1 indicates that the four variables turned out to be of order one.

[Table 1]

We now estimate $(p+1)^k$ number of regressions, where p is the maximum number of lags and k is the number of variables in the model, to determine the model lag selection. In order to select the optimum lag order for the model we focus on the Akaike Information Criterion (AIC) as well as the Schwarz Bayesian Criterion (BSC). We use a high enough order to ensure that the optimal one is not exceeded. As we can see in Table 2 the optimum lag order is (2, 3, 0, 2) according to the AIC, but it is (2, 2, 0, 1) according to the SBC. Based on the minimum value of the standard error of regression, we finally choose the order selected by the SBC.

[Table 2]

When working with long time series it is advisable to pay special attention to structural changes. Thus, in order to provide stability to the model, we consider the possibility that a break or more than a break exist for intercept and trend. Since we do not know a pre-specified date for possible breaks, we will look for endogenous ones. Thus, using a trimming of 10% of the observations in each of two subsamples we check a break recursively. The one-by-one-break method based on that of Banerjee et al.

(1998) keeps the analysis simple and at the same time provides easily interpretable results based on the critical values tabulated in Andrews and Ploberger (1994).For each time point, the *F*-statistics for testing the null hypothesis of no break is computed. Then, a break point is defined at point for which the *F*-statistics attains its maximum. Empirical results indicate the presence of two structural breaks, which refer to years 1917 and 1973.[11]The first break coincides with World War I, which is not surprising given that Spain was benefitted from its neutrality basically through increased production in textile and metallurgical sectors. The second break point matches with the first oil shock, time when Spain enters into a deep recession given that it heavily depended on oil imports.

Next, we carry out another test to be sure that the selected model with both breaks is dynamically stable. Thus, we check that all the inverse roots of the characteristic equation associated with our ARDL model lie strictly inside the unit circle. Since this is a two-lag model (according to the selection test shown in Table 2), the number of inverse roots will also be two. The chart in form of Argand diagram in Figure 2 indicates that the two roots are real roots (they are in the X-axis) and lie inside the unit circle. Therefore, it can be confirmed that our dynamic model is stable.

[Figure 2]

Once the properties of time series have been analyzed, the optimum lag order was determined, and different stability tests were done, we have to check whether or not a cointegrating relationship (long-run nexus) exists in the ARDL context and estimate the long run coefficients. Thus, after estimating by ordinary least squares (OLS) an Eq. (2) type (with two structural breaks in the intercept and trend), a bounds test is carried

[11] Results of including breaks one by one are available from the authors upon request.

out. As we can see in Table 3, the computed F-statistics is 6.3399 which indicates that there is a cointegrating relationship among $ln\ CO_2$, $ln\ y$, $(ln\ y)^2$, and $ln\ p$ at 1% level.

[Table 3]

In order to check for the existence of an EKC we have to observe the signs of the coefficients associated to $ln\ y$ and $(ln\ y)^2$. Both coefficients have the expected signs, that is, positive for per capita income and negative for its quadratic form. The two of them are statistically significant. Now, we can obtain the estimated turning point regarding per capita income ($\hat{y} = exp\left(-\hat{\lambda}_2 / 2\hat{\lambda}_3\right)$). It is reached at 9,236 US$ (1990 base), which approximately corresponds to per capita income in year 1980. Then,we could interpret this result by saying that from that date on, CO_2 starts decreasing as per capita income grows.

Regarding the oil price long-run parameter, we can see that it is quite significant. Its inelasticity explains a less than proportional effect on emissions each time that there is a variation in real crude oil prices. More specifically, its magnitude is -0.49 and should be interpreted as a 1% increase in oil prices causes a 0.49% reduction of CO_2 emissions. Moreover, in order to know the importance of real oil price variable in our empirical analysis, we have estimated a second model without including this variable. Results suggested some estimation bias if oil price movements are excluded. More specifically, from the estimated coefficients presented in Table 4 we can see that, in this case, the EKC would reach an earlier turning point at 8,103 US$ (1990 base), which corresponds to per capita income in year 1974.

[Table 4]

Lastly, in Table 5 we can see the error correction representation for the selected model. The underlying regression passes all diagnostic tests (such as Lagrange Multiplier test for serial correlation, and heteroscedasticity, Jarque-Bera normality test of residuals, Reset test on functional form, or model specification). Furthermore, the sensitivity analysis makes the model econometrically robust. The estimated coefficients have expected signs and values. Thus, past values of CO_2 (in differences) explain the evolution of CO_2. The remaining short-run elasticities are (in absolute terms) lower than the long-run elasticities, which is something expected. That is, income elasticity and its quadratic form show the same signs as in the long term but now, the absolute magnitude is clearly lower. The oil price elasticity has a positive sign, but the coefficient is practically marginal and not significant. Moreover, its lagged value is significant but again close to zero. The error correction coefficient is statistically highly significant and has the correct sign (negative), which confirms the established long run relationship among the variables. This last coefficient value entails that the rate of adjustment toward the long-run equilibrium is about 3% over each year.

[Table 5]

6. CONCLUSIONS

The aim of this paper has been the estimation of an environmental Kuznets curve dynamic structure for Spain during 1874-2011. We made use of the availability of very long time series for income, population, and CO_2 emissions. Unlike previous papers, the role of real oil prices has been taken into account. The motivations for this extended econometric model are based first, on the fact that Spain has been traditionally highly dependent on energy derived from petroleum and second, on the idea that real oil prices could enhance the results so far obtained.

Our empirical results support the idea that changes in real oil prices are relevant in order to explain the evolution of CO_2 emissions. The estimated coefficients associated to per capita income and to its quadratic form have opposite signs, suggesting the presence of a turning point. This is obtained for some years later than in the case where the oil price variable was omitted. More specifically, we can infer that from 1980 the economic growth in Spain has experienced an environmental improvement through the reduction of CO_2 emissions. Likewise, we show evidence that the rise in real oil prices also helps to decrease these emissions.

The revealed importance of oil prices suggests that the there is margin for implementing specific policy measures in order to improve the environmental quality. The design of a new carefully oil tax structure to reduce consumption of fossil fuels and to promote the energy efficiency should be seriously taken into consideration. Thus, for example, part of taxes that consumers would pay to the government could be levied on fuel for vehicles. This practice could significantly reduce the utilization of cars and encourage toward the use of more efficient vehicles avoiding this way emission of carbon dioxide to the atmosphere. It is obvious that citizens may have high resistance to implementation of new taxes, but an oil tax can be more accepted if policymakers present it as part of a broad revenue-neutral fiscal reform, where the oil tax could replace other specific taxes.

Acknowledgements

Financial support from the European Union and the Spanish Ministry of Economy and Competitiveness (ECO2011-28155) is gratefully acknowledged.

References

Agras J., Chapman D., 1999. A dynamic approach to the environmental Kuznets curve hypothesis. Ecological Economics. 28, 267-277.

Andrews D.W.K., Ploberger W., 1994. Optimal tests when a nuisance parameter is present only under the alternative. Econometrica. 62, 1383-1414.

Bachmeier, L., Li, Q., Liu, D., 2008. Should oil prices receive so much attention? An evaluation of the predictive power of oil prices for the US economy. Economic Inquiry. 46, 528-539.

Baldwin, R., 1995. Does sustainability require growth? in: Goldin, I., Winters, L.A. (Eds.), The Economics of Sustainable Development. Beacon Press, Boston, MA, pp. 19-47.

Banerjee, A., Lazarova, S., Urga, G. 1998. Bootstrapping sequential tests for multiple structural breaks. European University Institute.

Bo, S., 2011. A literature survey on environmental Kuznets curve. Energy Procedia. 5, 1322-1325.

Engle, R., Granger, C. 1987. Co-integration and error correction: representation, estimation, and testing. Econometrica. 55, 251-276.

Esteve, V., Tamarit, C., 2012a. Is there an environmental Kuznets curve for Spain? Fresh evidence from old data. Economic Modelling. 29, 2696-2703.

Esteve, V., Tamarit, C., 2012b. Threshold cointegration and nonlinear adjustment between CO2 and income: the environmental Kuznets curve in Spain, 1857–2007. Energy Economics. 34, 2148-2156.

Fernández Navarrete, D., 2005. La política económica exterior del franquismo: del aislamiento a la apertura. Historia Contemporánea. 30, 49-78.

Grossman, G., Krueger, A. 1991. Environmental impacts of North American free trade agreement. National Bureau of Economic Analysis. Technical report.

Haug, A., 2002. Temporal aggregation and the power of cointegration tests: A Monte Carlo study. Oxford Bulletin of Economics and Statistics. 64, 399-412.

Heil, M.T., Selden T.M., 2001. Carbon emissions and economic development: future trajectories based on historical experience. Environment and Development Economics 6, 63-83.

Kilian, L.,Vigfusson, R.J., 2013. Do oil prices help forecast us real GDP? The role of nonlinearities and asymmetries. Journal of Business and Economic Statistics. 31, 78-93.

Kijima, M., Nishide, K., Ohyama, A., 2010. Economic models for the environmental Kuznets curve: A survey. Journal of Economic Dynamics and Control. 34, 1187-1201.

Kuznets, S., 1955. Economic growth and income inequality. American Economic Review. 45, 1-28.

Lardic, S., Mignon, V., 2006. The impact of oil prices on GDP in European countries: An empirical investigation based on asymmetric cointegration. Energy policy. 34, 3910-3915.

Ministerio de Industria, Energía y Turismo (2011). Libro de la energía en España 2011. (http://minetur.gob.es.)

Molinas, C., Prados de la Escosura, L., 1989. Was Spain different? Spanish historical backwardness revisited. Explorations in Economic History. 26, 385-402.

Panayotou, T., 1993. Empirical tests and policy analysis of environmental degradation at different stages of economic development. Working Paper WP-238, Technology and Employment Programme, ILO, Geneva.

Pasten, R., Figueroa B.E., 2012. The environmental Kuznets curve: a survey of the theoretical literature. International Review of Environmental and Resource Economics. 6, 195-224.

Payne, J.E. 2010. Survey of the international evidence on the causal relationship between energy consumption and growth. Journal of Economic Studies. 37, 53-95.

Pesaran, M.H., Shin, Y., 1996. Cointegration and the speed of convergence to equilibrium. Journal of Econometrics. 71, 117-43.

Pesaran, M.H., Pesaran, B., 1997. Working with Microfit 4.0: Interactive econometric analysis. Oxford University Press, Oxford.

Pesaran, M.H., Shin, Y., 1999. An autoregressive distributed lag modelling approach to cointegration analysis, in: Strom, S. (Eds.), Econometrics and Economic Theory, in

20th Century: The Ragnar Frisch Centennial Symposium. Cambridge University Press, Cambridge.

Pesaran, M.H., Shin, Y., Smith, R. J., 2001. Bounds testing approaches to the analysis of level relationships. Journal of Applied Econometrics. 16, 289-326.

Richmond A.K., Kaufman R. K., 2006. Energy prices and turning points: The relationship between Income and energy use/carbon emissions. Energy Journal.7, 157-180.

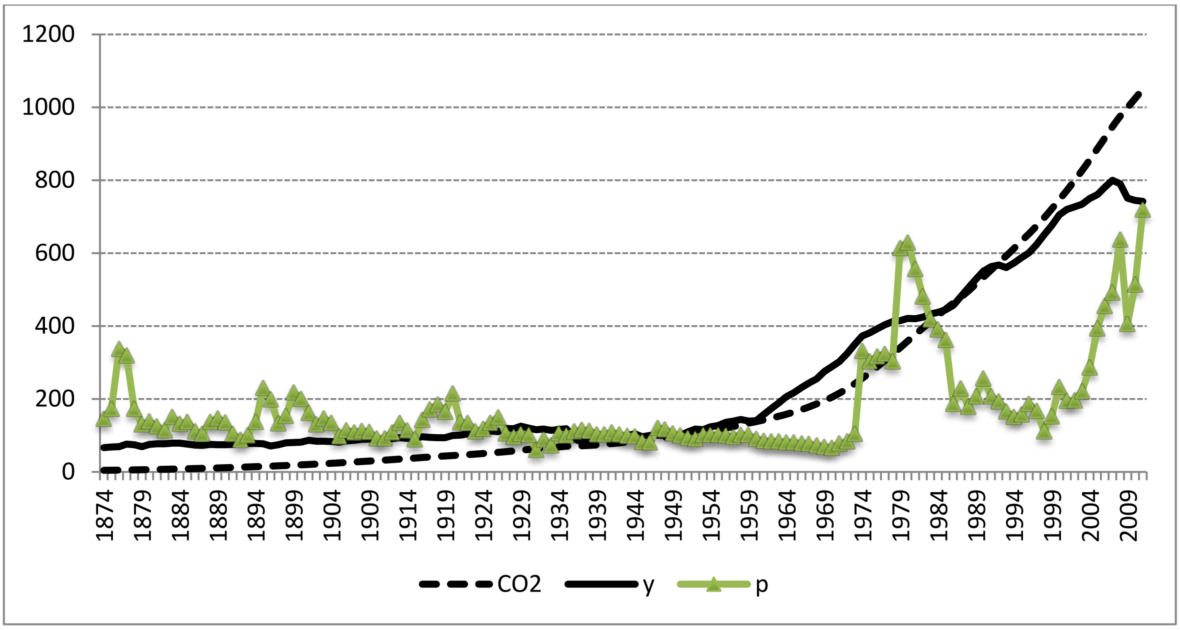
Roca J., Padilla E., Farre M., Galletto V., 2001. Economic growth and atmospheric pollution in Spain: discussing the environmental Kuznets curve hypothesis. Ecological Economics. 39, 85-99.

Shafik, N., Bandyopadhyay, S., 1992. Economic growth and environmental quality: time series and cross-country evidence. Background Paper for World Development Report 1992. World Bank, Washington, DC.

Sephton P., Mann J. 2013. Further evidence of an environmental Kuznets curve in Spain. Energy Economics. 36, 177-181.

Sudrià C., 1995. Energy as a limiting factor to growth, in Martín Aceña, Pablo y Simpson, James, The Economic Development of Spain since 1870, Edward Elgar, Aldershot.

Figure 1. CO_2, per capita GDP, and oil prices (100 for 1950 base year)



Data sources: The CO_2 emissions series were obtained from Carbon Dioxide Information Analysis Center (CDIAC) of the US Department of Energy.The per capita income (noted as *y*) is expressed as the ratio of GDP to population. GDP is in real terms of 1990 US dollars. Both variables are compiled by Maddison "Historical Statistics" and the Spanish National Institute of Statistics (INE). The oil prices (labeled as *p*) were gathered from the Statistical Review of World Energy 2013 provided by the British Petroleum.

Figure 2. Inverted AR Roots from the ARDL Model

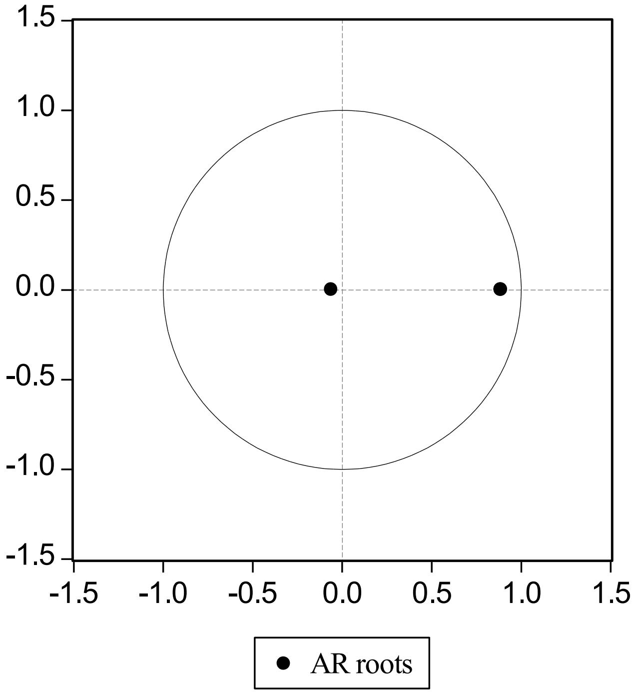


Table 1. Augmented Dickey Fuller Tests

	ADF statisctics I(1) versus I(0)	ADF statistics I(2) versus I(1)	ADF statistics DSP versus TSP
ln CO_2	-1.15(1)	-3.12(2)	-1.91(1)
ln y	0.70(1)	-7.40(1)	-1.21(1)
$(ln\ y)^2$	0.87(3)	-4.95(3)	-1.13(3)
ln p	-2.71(1)	-7.23(3)	-2.93(1)
Critical values	*-2.88*	*-2.88*	*-3.44*

The numbers in brackets are the lags used in the ADF test in order to remove serial correlation in the residuals. DSP stands for difference stationary process and TSP means trend stationary process.

Table 2. Optimum Lag Order for Model Selection

	AIC	SBC
Order for $ln\ CO_2$, $ln\ y$, $(ln\ y)^2$, $ln\ p$	(2, 3, 0, 2)	(2, 2, 0, 1)
Standard error of regression	0.00219	0.00216

The regressions are run based on the autoregressive distributed lag method.

Table 3. Cointegration Results and Long Run Estimates of the ARDL Model

	Lower bounds			Upper bounds			*F*-statistics
	10%	5%	1%	10%	5%	1%	
	2.45	2.86	3.74	3.52	4.01	5.06	F(4,122)= 6.3399

	Coefficients	*T*-statistic	P-value
α	-10.31	-0.77	(0.444)
ln y	7.67	2.53	(0.012)
$(ln\ y)^2$	-0.42	-2.40	(0.018)
ln p	-0.39	-2.24	(0.027)
Breack (1917)	-0.48	-2.27	(0.025)
Breack (1973)	0.65	2.61	(0.010)

Sensitivity analysis and other statistics are provided in the error correction model.

Table 4. Long Run Estimates of the ARDL Model Without Oil Prices

	Coefficients	*T*-statistic	P-value
α	-25.63	-1.81	(0.073)
ln y	11.16	3.34	(0.001)
$(ln\ y)^2$	-0.62	-3.17	(0.002)
Break (1917)	-0.54	-1.89	(0.061)
Break (1973)	0.74	2.26	(0.025)

Table 5. Error Correction Representation for the Selected Model

	Coefficient	*T*-statistics	P-value
α_0	-0.047	-0.95	(0.346)
dln CO_2*(-1)*	0.791	16.78	(0.000)
dln y	0.027	1.87	(0.064)
$d(ln\ y)^2$	-0.003	-0.32	(0.746)
dln p	0.008	1.01	(0.313)
dln p(-1)	0.002	2.20	(0.030)
dBreak (1917)	-0.001	-2.05	(0.042)
dBreak (1973)	0.005	3.16	(0.002)
ect (-1)	-0.030	-3.26	(0.001)
Statistics			
R^2 adjusted	0.97		
S.E. of regression	0.02		
F-Statistics (9, 128)	526.26		(0.000)
Diagnostic tests			
Serial correlation			(0.716)
Functional form			(0.740)
Normality			(0.735)
Heteroscedasticity			(0.135)

Variables starting with a *d* means differenced once; variables lagged one period are expressed as (-1); *ect* is the error correction term.