

Output volatility in the OECD: Are the member states becoming less vulnerable to exogenous shocks?

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ABSTRACT

This paper analyses the vulnerability of OECD member states to external shocks by estimating the degree of asymmetric effects from positive and negative shocks. We use asymmetric conditional heteroscedasticity models with endogenously determined regime changes in a context of progressive moderation in both moments. The results suggest that recessions are associated with higher volatility and significant leverage effects. The estimated impacts of negative and positive shocks amount to 0.961 and 0.028 respectively. The disaggregated analysis over different periods reveals an increasing pattern of these asymmetries, as well as huge differences among the countries. The country-specific analysis suggest an increasing vulnerability to negative exogenous shocks in Australia, Denmark, Finland, Japan, Mexico, the Netherlands, Turkey and the United Kingdom, although with different levels, and decreasing vulnerability in Canada, Greece, Italy and New Zealand. Finally, some economies seem to have developed higher levels of immunity to external shocks by reaching balanced effects from positive and negative shocks. Among these are the largest European economies, together with the northern economies, the United States and the wealthiest economies of Luxembourg and Switzerland.

1. INTRODUCTION

The linkage between economies' growth rates and their volatility has long been a subject of intense debate on both theoretical and empirical grounds. The relevance of this issue rests on the implications of growth volatility on countries' economic performance and the usefulness of getting knowledge on its behaviour for proposes of policy design. This issue poses a particular challenge as real gross domestic product (GDP) growth involves a long run perspective, over which structural changes in volatility are very likely to occur. Their occurrence has, in fact, been widely documented in the literature (see, among other, Kim and Nelson, 1999; McConnell and Perez-

Quiros, 2000; and Blanchard and Simon, 2001, for the United States; Bhar and Hamori, 2003; Mills and Wang, 2003; and Summers, 2005, for Japan and other G7 countries). All of these studies report rather dramatic reductions in the volatility of output growth; and the occurrence of this phenomenon in several countries has been known in the literature as the 'Great Moderation' (Bodman, 2009).

However, while the decline of output volatility has been widely confirmed by empirical evidence, a lack of consensus on the linkage between the growth rate and volatility has emerged on theoretical grounds. On the one hand, a positive relationship is suggested by the perspective that agents choose to invest in riskier and hence more volatile assets only if the associated risk is offset by the expected rates of return. Schumpeterians postulate that the economic instability generated by the process of 'creative destruction' would improve economic efficiency and thereby boost long term growth. The positive relation is found in Grier and Tullock (1989) and Caporale and McKiernan (1996, 1998), for example. On the other hand, suspicions of a negative relationship are raised by the idea that higher volatility is associated with higher uncertainty which, in turn, constrains investments. Studies confirming this evidence include, for example, Ramey and Ramey (1995), Martin and Rogers (2000), and Henry and Olekalns (2002). Finally, several other studies, including Speight (1996) and Grier and Perry (2000), report no significant relation.

On methodological grounds, some form of Generalised Autoregressive Conditional Heteroscedasticity (GARCH) modelling strategy has been adopted in dealing with output growth volatility. However, most studies assume a stable modelling approach or even an asymmetric approach without considering the occurrence of potential structural changes. This issue raises several problems in the analysis, as documented in the literature. Diebold (1986) first argued that structural changes may confound persistence estimation in GARCH models. He claimed that Engle and Bollerslev's (1986) integrated GARCH may result from instability of the constant term of the conditional variance (i.e., nonstationarity of the unconditional variance). Neglecting such changes can generate spuriously measured persistence, with the sum of the estimated autoregressive parameters of the conditional variance heavily biased towards one. Lamoureux and Lastrapes (1990) provide confirming evidence that ignoring regime changes in the unconditional variance can bias upwards GARCH estimates of persistence in variance, while the use of dummy variables to account for such shifts diminishes the degree of GARCH persistence. Alternatively, Hamilton and Susmel (1994) and Kim *et al.* (1998) suggest that the long-run variance dynamics may include regime shifts, but within a given regime it may follow a GARCH process. Empirical evidence on this is also provided by, *inter alia*, Kim and Nelson, 1999; Bhar and Hamori, 2003; Mills and Wang, 2003; Mikosch and Stărică, 2004; Hillebrand, 2005; and Summers, 2005.

Furthermore, there is evidence that the neglect of structural breaks in output growth and/or unconditional or conditional volatility, have led to high persistence in the conditional volatility (see Hamori 2000; Ho and Tsui 2003; and Fountas *et al.* 2004, among others).

Another relevant issue is that most studies postulate that the relation between volatility and growth is symmetric across economies' business cycles. More specifically, most empirical models implicitly assume that the sign (and size) of the volatility-growth relation is the same whether the economy is in contraction or expansion. However, there is no *a priori* reason to believe that is the case and it is conceivable that the sign of the volatility-growth relation depends on business cycle phases.

In this context, most studies have attempted to give a contribution to clarify the nature of the growth-volatility relationship in a limited set of countries OECD countries (see, among others, Fountas and Karanasos, 2006; Fang and Miller 2008, 2009; Fang *et al.*, 2008) or the trade-off between the variability of inflation and output gap (Lee, 1999, 2002, among others).

The evidence of the 'Great Moderation' and structural changes in output growth volatility, combined with high persistence in conditional volatility for several countries, raises the question how different is the reaction of volatility to negative versus positive shocks in growth rates. This constitutes the opportunity that motivates us to revisit the issue of conditional volatility of GDP growth rates in the OECD, addressing the issue of potential asymmetry of the relationship between volatility and business cycles in the presence of structural breaks.

Specifically, our objective is to analyse whether the OECD member-states are becoming more vulnerable to external shocks, by estimating the asymmetry of effects from positive and negative shocks. On methodological grounds, we model the asymmetric growth behaviour over the sample period and estimate the effects of news on volatility over the business cycle, to assess the vulnerability to external shocks, accounting for different country-specific experiences, in a context of a generalised reduction of GDP volatility in all OECD member-states. Using a dataset for the whole sample period and for distinct sub-sample periods, we gather knowledge about the growing/decreasing vulnerability of the OECD and each member-state to asymmetric external exogenous shocks. By developing a country-specific analysis, we gather information on how individual states have contributed to results at the aggregate level.

Another relevant contribution comes from the use of the Lee and Strazicich (2003) structural breaks test. It consists on an endogenous two-break Lagrange Multiplier test that allows for breaks under both the null and alternative hypothesis. In this way, the rejection of the null unambiguously implies trend stationary, and therefore this test overrules the criticisms of other tests used in previous studies (Zivot and Andrews, 1992; and its extension, Lumsdaine and Papell, 1997) and benefits from greater power.

The paper is organised as follows. Section 2 presents some stylised facts on GDP growth and volatility, and the existence of structural breaks.

Section 3 reports the methodological background. Section 4 presents the results on GDP volatility, focusing on the asymmetric effects across the business cycle, its evolution for the OECD, in particular in the periods before and after the adoption of the euro by the euro-zone member-states, and on a country-specific basis. Finally, Section 5 reports the main conclusions.

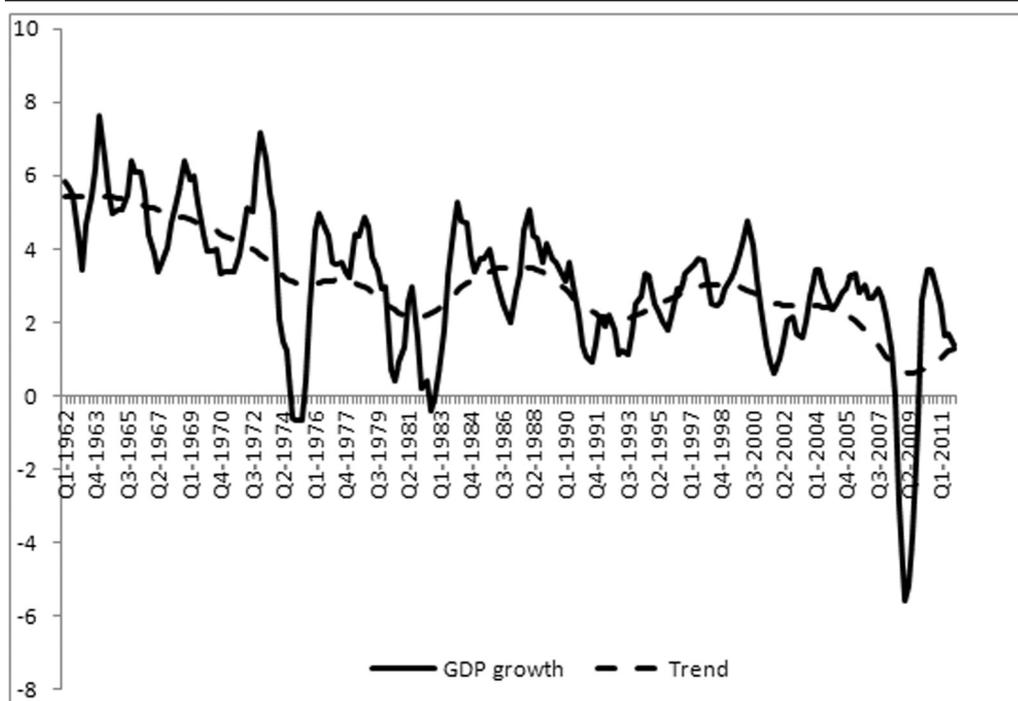
2. ASSESSING ‘THE GREAT MODERATION’ IN THE OECD: DATA, STATISTICS AND REGIME CHANGES

2.1 Data sources and descriptive statistics

This paper uses data on quarterly real GDP over the period from 1961:1 to 2011:4, for the OECD and the individual member-states, in a total of 34 countries.⁴ The data are seasonally adjusted and come from the OECDs online statistical database.

A preliminary analysis of the evolution of GDP annualised growth rates in the OECD is depicted in Figure 1. It shows clearly the decreasing trend observed since the 1970s. Simple quantitative measures of the sample statistical moments are summarised in Table 1. Panel A exhibits the average growth

Figure 1. Preliminary evidence on trend and volatility of GDP growth in OECD (1962:1-2011:4)



Source: OECD and authors' calculations.

rates and volatility measure over the sample period. The annual average growth reached 3.07 per cent, with maximum and minimum values of 7.62 per cent in 1964:1 and -5.56 per cent in 2009:1, while output volatility, measured by the standard deviation, was 1.94.

Table 1. Summary statistics of real growth rates

| <i>Panel A: general statistics for the sample period</i> | | |
|--|-----------------|---------------------------|
| Period: 1961:1-2011:4 | | |
| Mean: 3.07% | | |
| Maximum: 7.62% | Minimum: -5.56% | |
| St. Dev.: 1.94 | | |
| <i>Panel B: moment statistics by decade</i> | | |
| Period | Sample mean (%) | Sample standard deviation |
| 1962:1 – 1969:4 | 5.26 | 0.968 |
| 1970:1 – 1979:4 | 3.60 | 1.78 |
| 1980:1 – 1989:4 | 2.85 | 1.53 |
| 1990:1 – 1999:4 | 2.59 | 0.84 |
| 2000:1 – 2011:4 | 1.83 | 2.16 |
| Source: OECD and authors' calculations. | | |

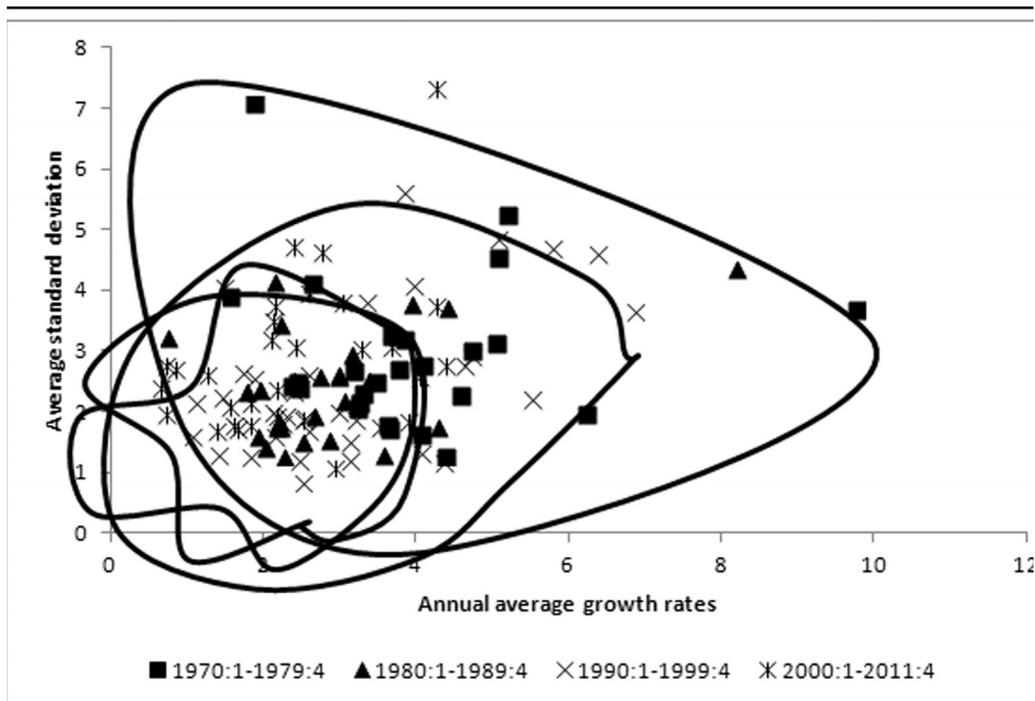
The analysis over shorter time periods is displayed in Panel B. This clearly mirrors the decline in both moments over time. The average growth rates of 5.26 per cent per annum in the 60s and 3.60 per cent in the 70s declined to 2.85 per cent in the 1980s, 2.59 per cent in the 1990s and 1.83 per cent over the last decade. The results also illustrate the significant decline in real GDP volatility since the late 1970s. In fact, after an increase over the 1970s, the standard deviation reduced to 1.53 over the 1980s, and 0.84 over the 1990s. A slightly increase was observed again in the last decade.

Therefore, the analysis clearly illustrates the occurrence of the 'Great Moderation' phenomenon in the OECD, by the end of the 1970s. A more detailed country by country analysis is also illustrative of this phenomenon. Figure 2 reveals that the reduction of average growth rates is common to all countries, decade after decade, along with a notable reduction of volatility.

Several explanations, not exclusive to any member-state in particular, have been proposed for these occurrences. They include a change in the economies' structure due to advances in information technology, increased

resilience of economies to oil shocks, increased access to financial markets, changes in financial market regulation, improvements in the conduct of monetary policy, and a reduction in the size and volatility of domestic and international shocks, among other factors. On the other hand, structural changes are very likely to occur in GDP growth time series for any number of reasons, such as economic crisis, changes in institutional arrangements, policy changes and regime shifts.

Figure 2. Growth rates and volatility among OECD member states



Source: OECD and authors' calculations.

Although many studies find no apparent break in the average growth rate of GDP for the United States (US), other studies report permanent falls in GDP average growth in almost all countries (see McConnell and Perez-Quiros, 2000; and Blanchard and Simon, 2001, among others). Examples include Portugal, with a decline from around 5 per cent in the 70s to just over 3 per cent per annum in the last decade. This picture is fairly similar in timing and magnitude to the fall experienced in other countries, such as Canada (Debs, 2001; Voss, 2004;) and Australia (Bodman, 2009). The US also experienced a similar decline in volatility in the mid-80s. Therefore, this preliminary analysis clearly illustrates the suspicion that GDP growth in OECD has gone

through fluctuations in trend and volatility, in much the same way as in each member-state. Therefore, this issue should not be neglected in the analysis that follows.

2.2 Assessing the occurrence of regime changes

Following the idea that allowing for only one structural break may be too restrictive, and attending to econometric advances on unit root tests with structural changes that emerged in the literature after Perron's (1989) influential article, we use the Lee and Strazicich (2003) test (LS Test, hereafter).⁵

The models proposed by these authors are two of the three structural break models considered in Perron (1989): the 'crash' model A allows for a one-time change in level; and the C model allows for a change in both the level and trend.⁶

Considering the following data-generating process (DGP):

$$y_t = \delta'Z_t + e_t, \quad e_t = \beta e_{t-1} + \varepsilon_t \quad (1)$$

where Z_t is a vector of exogenous variables and ε_t is a gaussian error term. Two structural breaks can be considered as follows: Model A allows for two level-breaks, while Model C allows for two level and two slope-breaks. As such, Model A can be described by $Z_t = [1, t, D_{1t}, D_{2t}]'$, where $D_{j,t} = 1$ for $Z_t = [1, t, D_{1t}, D_{2t}]'$ and 0 otherwise; and Model C can be described by $Z_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]'$ where $DT_{jt} = t - T_{Bj}$ for $t \geq T_{Bj} + 1, j=1,2$ and 0 otherwise. Considering the more general Model C, the DGP allow breaks both under the null ($H_0: \beta = 1$) and the alternative ($H_1: \beta < 1$) hypothesis, which are represented as follows:

$$\begin{aligned} H_0 : y_t &= \alpha_0 + d_1 B_{1t} + d_2 B_{2t} + d_3 DT_{1t} + d_4 DT_{2t} + y_{t-1} + \mu_{1t} \\ H_A : y_t &= \alpha_1 + \gamma t + d_1 B_{1t} + d_2 B_{2t} + d_3 DT_{1t} + d_4 DT_{2t} + (1-\alpha) y_{t-1} + \mu_{2t} \end{aligned} \quad (2)$$

where μ_{1t} and μ_{2t} are stationary error terms, $B_{jt}=1$ for $t=T_{Bj}+1, j=1,2$ and 0 otherwise.⁷ The LM unit-root test statistic is generated from the following regression:

$$\Delta y_t = \delta' \Delta Z_t + \theta S_{t-1} + \sum_{i=1}^{\infty} \pi_i \Delta S_{t-i} + \vartheta_t$$

where the de-trended series \tilde{S}_t is defined as $\tilde{S}_t = y_t - \tilde{\varphi}_x - Z_t \tilde{\delta}, t=2, \dots, T$; $\tilde{\delta}$ are the coefficients of the regression of Δy_t onto ΔZ_t ; $\tilde{\varphi}_x$ is given by $y_1 - Z_1 \tilde{\delta}$ where y_1 and Z_1 correspond to the first observations of y_t and Z_t , respectively. The lagged terms $\Delta \tilde{S}_{t-i}$ are considered to correct for serial correlation. The LM test statistic equals the t -ratio, $\hat{\tau}$, testing the unit-root null hypothesis, $\theta=0$. Lee and Strazicich (2003) show that the asymptotic null distri-

bution of the two break LM unit root test for Model C depends on the location of the breaks

$$\left(\lambda_1 = \frac{TB_1}{T}, \lambda_2 = \frac{TB_2}{T} \right)$$

although not on their magnitude.⁸ To determine endogenously the location of the two breaks, the minimum LM unit root test uses a grid search given by $LM_\rho = \text{Inf}_\lambda \tilde{\rho}(\lambda)$ and $LM_\tau = \text{Inf}_\lambda \tilde{\tau}(\lambda)$ whose asymptotic distributions and critical values are tabulated in Lee and Strazicich (2003). For testing purposes, we consider Model C that allows structural breaks on both the intercept and slope of the series.⁹

The results of the LS Test are reported in Table 2. The null hypothesis of a unit root in GDP growth rate is rejected for the conditional mean, which confirms that the series is trend stationary. Two breaks in the intercept and slope are identified in 1973:4 and 2001:4. To test for the instability of the volatility (the conditional variance), the identified structural breaks are included in the growth rate series and the non-constant mean removed. The null hypothesis of a unit root is again rejected in the case of the conditional variance, in favour of the existence of two structural breaks, in 1974:1 and 2000:1. Once again, the coefficients of the dummy variables report statistically significant impacts of structural changes on both the intercept and slope.

Table 2. Endogenous two break LS unit root test

| TB ₁ TB ₂ | y_{t-1} | B_{1t} | B_{2t} | DT_{1t} | DT_{2t} | Critical Value Break Points | k |
|--|------------|------------|------------|------------|------------|--------------------------------|-----|
| Conditional mean | | | | | | | |
| 1973:4 | -0.3638** | -2.4453*** | 0.9499** | 0.1761** | -0.4764*** | $\lambda = (0.24, 0.80)$ | 9 |
| 2001:4 | (-5.754) | (-4.496) | (1.757) | (1.718) | (-3.181) | | |
| Conditional standard deviation (breaks in the mean) | | | | | | | |
| 1974:1 | -1.4026*** | 1.8653*** | -1.1171*** | -1.2038*** | 1.1057*** | $\lambda = (0.27, 0.80)$ | 10 |
| 2000:1 | (-9.829) | (8.001) | (-4.883) | (-9.488) | (9.510) | | |

Notes: the numbers in parenthesis are the t-statistics for the estimated coefficients. TB_1 and TB_2 are the break dates; k is the lag length; the coefficient on y_{t-1} tests for the unit root; B_{1t} and B_{2t} correspond to the breaks in the intercept and DT_{1t} and DT_{2t} correspond to the breaks in the slope. Critical values for the coefficients on the dummy variables follow the standard normal distribution. The critical values for the unit root test are tabulated in LS (2003, Table 2) and depend upon the location of the breaks. For $\lambda_1=0.24$ and $\lambda_2=0.8$ the critical values, calculated by interpolation, are -6.348, -5.698, and -5.328 for 1%, 5% and 10% levels, respectively. For $\lambda_1=0.27$ and $\lambda_2=0.8$ the critical values, calculated by interpolation, are -6.362, -5.689, and -5.327 for 1%, 5% and 10% levels, respectively.

Source: Authors' calculations.

3. METHODOLOGICAL FRAMEWORK

The ARCH models are design to model and forecast the conditional variance. In each case the variance of the dependent variable is specified to depend upon past values of the dependent variable using some formula. A general ARMA(r,s)-GARCH(p,q)-M process is specified as follows,

$$\Phi(L)y_t = \mu + \Theta(L)u_t + \delta h_t \quad (4)$$

$$B(L)h_t = \varpi + A(L)u_t^2 \quad (5)$$

where

$$\Phi(L) = 1 - \sum_{j=1}^r \varphi_j L^j; \quad \Theta(L) = -\sum_{j=1}^s \theta_j L^j; \quad B(L) = 1 - \sum_{i=1}^p \alpha_i L^i; \quad A(L) = \sum_{i=1}^q \beta_i L^i$$

The potential dependence of the nature of the volatility-growth relation on the business cycle phase requires the use of methods that account for this asymmetry. Our options go to the Threshold GARCH model (TGARCH) and the Exponential GARCH model (EGARCH). The TGARCH model, introduced by the works of Zakoian (1990) and Glosten *et al.* (1993), is defined as:

$$B(L)h_t = \varpi + A(L)u_t^2 + C(L)u_t^2 \quad (6)$$

where $C(L) = \sum_{i=1}^q \beta_{i+1} I_{t-i}$ and I_{t-i} for $u_t < 1$ and zero otherwise.

This specification allows the impacts of lagged squared residuals to have different effects on volatility, depending on their sign. While good news, given by $u_{t-i} > 0$ has an impact of α_i , bad news, expressed by $u_{t-i} < 0$ has an impact of $\alpha_i + \sum_{i=1}^q \beta_{i+1}$. Significant values for the leverage effect coefficients suggest asymmetries, with negative (positive) shocks having a greater impact upon volatility whether $\sum_{i=1}^q \beta_{i+1} > 0$ ($\sum_{i=1}^q \beta_{i+1} < 0$).

The EGARCH model, first developed by Nelson (1991), is defined as:

$$B(L)\ln(h_t) = \varpi + C(L)z_t$$

$$z_{t-i} = \beta_1 \frac{u_{t-i}}{\sqrt{h_{t-i}}} + \beta_2 \left[\frac{|u_{t-i}|}{\sqrt{h_{t-i}}} - E \left[\frac{|u_{t-i}|}{\sqrt{h_{t-i}}} \right] \right] \quad (7)$$

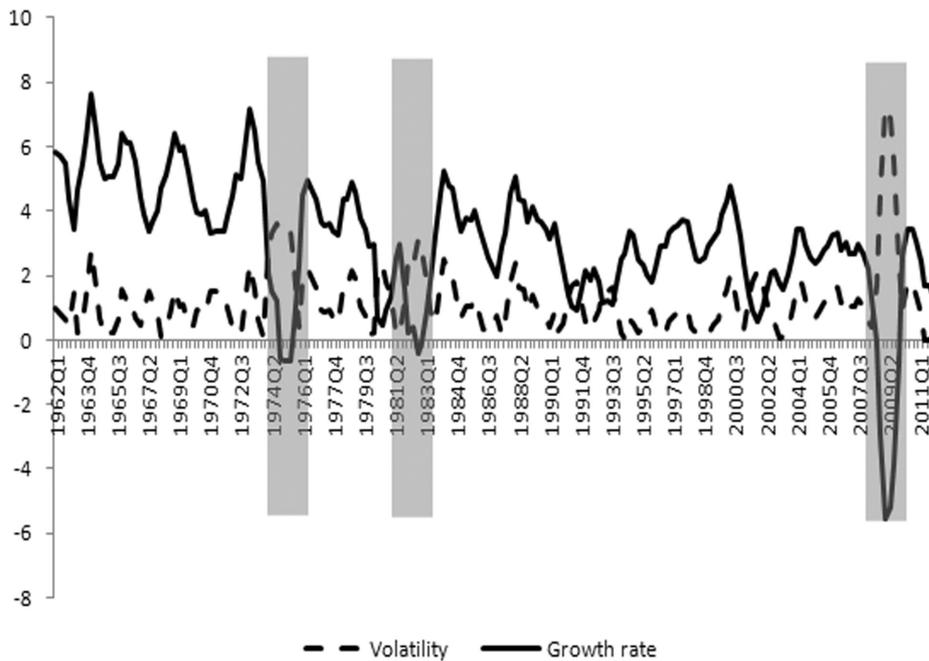
where $C(L) = \sum_{i=1}^q c_i L^i$ and $B(L) = \prod_{i=1}^p (1 - \alpha_i L)$ with $c_i = 1$. We estimate both models to check whether or not they provide similar results and, therefore, to assess their robustness. However, the choice upon which model best fits the data is based on the Akaike Information Criterion and Schwartz Information Criterion (AIC and SIC, respectively).

4. VOLATILITY AND GROWTH CYCLES: ASYMMETRIES AND TIME VARYING PATTERNS

4.1 *Volatility and business cycles: stylized facts and asymmetries*

The de-measured real GDP growth, i.e., real filtered GDP growth obtained by removing the non-constant mean, provides a measure of GDP volatility and is represented in Figure 3, together with GDP annual growth rates. Two remarks are in order. First, periods of positive growth seem to be characterised by a positive relationship between growth rates and volatility. Because expansions last longer than contractions, the volatility average values lie closer to the values they reach during expansions. Consequently, deviations from the output growth average are larger during periods of lower growth. This cyclical pattern seems to suggest the existence of potential asymmetries associated with the business cycle. Second, we observe periods of increased volatility, in particular, around 1975, 1982 and 2009. As these years coincide with recessions in the OECD, it seems that the asymmetry of the business cycle may account for part of the increase in measured volatility during recessions.

Figure 3. GDP volatility and the business cycle



Note: the shaded bars indicate recessions. Volatility is computed using the absolute value of the de-measured annual growth rate.

Source: OECD and authors' calculations.

Following the standard Box-Jenkins Autoregressive Integrated Moving Average (ARIMA) modelling procedure, and considering the model selection criteria, GDP growth is best modelled as an ARMA(1,3). The results are provided in column 1 of Table 3.¹⁰ The diagnostic tests on the residuals indicate significant persistence in the squared residuals and the Lagrange Multiplier tests (LM Test) for ARCH show that the assumption of constant error variance is not appropriate when modelling the GDP growth rate, as there is a significant non-captured structure in the second moment. Further analysis using the Brock, Dechert and Scheinkman (BDS) Test,¹¹ a portmanteau test for time based dependence in a series, indicates the existence of nonlinearities in the residuals.

To address these inadequacies and allow for time varying conditional variances, a GARCH(p,q) modelling procedure of the squared residuals is implemented. The corresponding results are reported in column 2 of Table 3, along with the residuals diagnostic tests. The coefficient estimates in the conditional mean specification are still significant at reasonable levels and the process stability is guaranteed in the conditional variance specification. The coefficient of the lagged square residual is also highly statistically significant, although the coefficient of the lagged conditional variance term is not. The dummies' coefficients are significant and negative, confirming the shift from higher to lower volatility regimes.

4.2 The cyclical features of volatility and business cycle dependence

Given the asymmetry observed in different phases of the business cycle, we estimate a TGARCH model and an EGARCH model, whose results are reported in columns 3 and 4 of Table 3. Both models report statistically significant leverage effects with similar magnitudes. The estimated leverage effect in the TGARCH is positive, suggesting that while the impact of good news on variance is 0.038, the impact of bad news is more than 16 times higher, 0.644. The leverage effect is also confirmed in the context of the EGARCH model. While the impact of positive shocks is 0.028, the impact of negative shocks of the same magnitude amounts to 0.961. Therefore, the statistical significance of the leverage effects, along with their signs, suggests that negative shocks to GDP growth cause higher volatility than positive shocks, thereby increasing the degree of uncertainty during recessions, and causing asymmetries of the corresponding news impact curves. These two models confirm the higher magnitude of the effects of negative shocks on volatility, when compared to positive shocks. This configures the robustness of our results. However, given the AIC and BIC values, the model that best fits the data is the EGARCH.

4.3 The time varying asymmetric nature of volatility

Having detected volatility change in GDP growth rates, analysis is conducted to further investigate whether the asymmetric effects exhibit a persistent pattern over time, or the volatility decline is associated with a change in the business

Table 3: Time varying volatility and asymmetric responses of volatility: 1962:1-2011:4 (regime shifts in the mean in 1973:4 and 2001:4)

| Parameters | ARMA(1,3) | ARMA(1,3)- GARCH(1,1) | ARMA(1,3)- TGARCH(1,1) | ARMA(1,3)- EGARCH(1,1) |
|---------------|----------------------|--------------------------|---------------------------|---------------------------|
| μ | 1.2661*** (0.302) | 1.3099*** (0.292) | 1.1482*** (0.235) | 1.3660*** (0.226) |
| η_1 | -0.0185 (0.175) | -0.8122*** (0.218) | 0.0245 (0.167) | -0.0746 (0.153) |
| η_2 | -0.4833* (0.258) | -0.3127** (0.151) | -0.1280** (0.061) | -0.2719* (0.164) |
| ϕ_1 | 0.6154*** (0.058) | 0.5776*** (0.048) | 0.5688*** (0.051) | 0.5295*** (0.061) |
| θ_1 | 0.8989*** (0.039) | 0.9249*** (0.037) | 0.9478*** (0.020) | 0.9559*** (0.013) |
| θ_2 | 0.9369*** (0.040) | 0.9436*** (0.029) | 0.9642*** (0.017) | 0.9823*** (0.001) |
| θ_3 | 0.9642*** (0.027) | 0.9023*** (0.016) | 0.9805*** (0.015) | 0.9742*** (0.014) |
| ω | | 0.1771* (0.101) | 0.2695*** (0.070) | -0.3625 (0.235) |
| α_1 | | 0.3824 (0.336) | 0.1873** (0.091) | 0.6176*** (0.099) |
| β_1 | | 0.2114*** (0.064) | 0.0377 (0.019) | -0.4666*** (0.118) |
| β_2 | | | 0.6063** (0.301) | 0.4945** (0.252) |
| ψ_1 | | -0.0863*** (0.032) | -0.1339* (0.074) | -0.2942*** (0.106) |
| ψ_2 | | -0.1383* (0.083) | -0.2323*** (0.069) | -0.7119*** (0.156) |
| R^2 | 0.935 | 0.935 | 0.933 | 0.932 |
| $J-B$ | 49.373*** [0.000] | 47.176 [0.000] | 14.829 [0.001] | 10.455 [0.005] |
| $LM ARCH (1)$ | 10.291*** [0.001] | 0.8123 [0.367] | 0.082 [0.775] | 1.4343 [0.231] |
| $LM ARCH (2)$ | 16.554*** [0.000] | 1.2465 [0.536] | 0.1237 [0.940] | 1.6976 [0.428] |
| AIC | 1.4993 | 1.3417 | 4.0464 | 1.3399 |
| BIC | 1.6152 | 1.5568 | 4.0964 | 1.5550 |

Note: Bollerslev-Wooldridge robust standard errors in square brackets; p-values in brackets; * indicates statistical significance at 10% level; ** indicates statistical significance at 5%; *** indicates statistical significance at 1%.

$$\text{Model ARMA}(1,3): [1 - \phi_1 L] y_t = \mu + \eta_1 DT1_t + \eta_2 DT2_t + \sum_{j=1}^3 \theta_j L^j u_t$$

$$\text{Model ARMA}(1,3)\text{-GARCH}(1,1): [1 - \alpha_1 L] h_t = \omega + \beta_1 L u_t^2 + \psi_1 DT3_t + \psi_2 DT4_t$$

$$\text{Model ARMA}(1,3)\text{-TGARCH}(1,1): [1 - \alpha_1 L] h_t = \omega + \beta_1 L u_t^2 + \beta_2 I_{t-1} L u_t^2 + \psi_1 DT3_t + \psi_2 DT4_t$$

$$\text{Model ARMA}(1,3)\text{-EGARCH}(1,1): (1 - \alpha_1 L) \ln(h_t) = \omega + z_{t-1} + \psi_1 DT3_t + \psi_2 DT4_t$$

DT1 and DT2 are dummy variables to reflect the regime changes in the mean: DT1=0 before 1973:4 and after 2001:4; DT2=0 before 2001:4. DT3 and DT4 are dummy variables to reflect the regime changes in the variance: DT3=0 before 1974:1 and after 1999:4; DT4=0 before 2000:1. Source: Authors' calculations.

Table 4: Time varying volatility and asymmetric responses of volatility in different periods (EGARCH specification): 1962:1 to 1999:4 and 2000:1 to 2011:4 (regime shifts in the mean in 1973:4 and 2001:4)

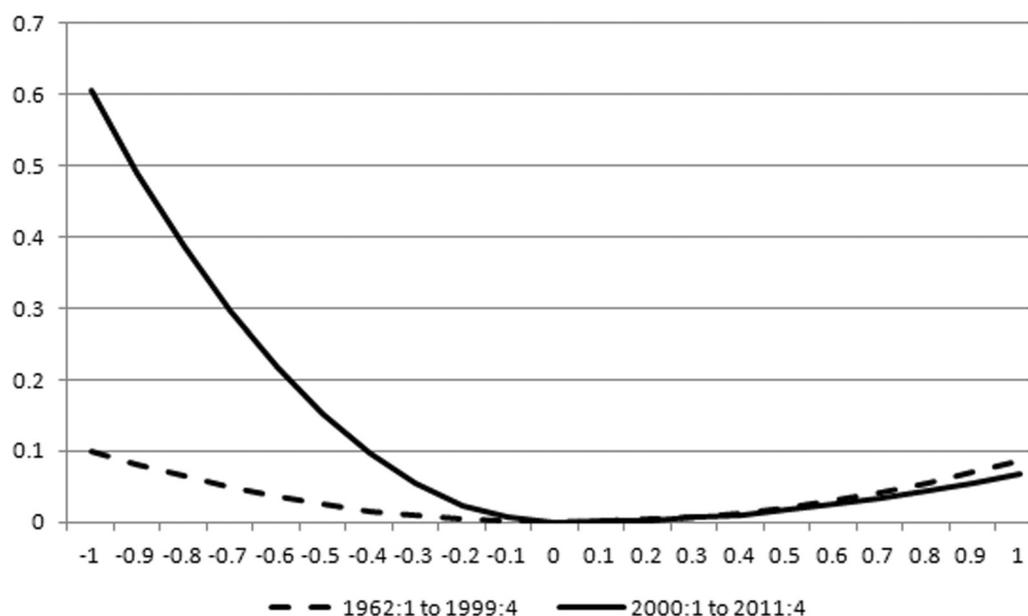
| Parameters | 1962:1 to 1999:4 | | 2000:1 to 2011:4 | |
|-------------|---------------------------|---------------------------|---------------------------|---------------------------|
| | ARMA(1,3)- TGARCH(1,1) | ARMA(1,3)- EGARCH(1,1) | ARMA(2,3)- TGARCH(1,1) | ARMA(1,3)- EGARCH(1,1) |
| μ | 2.4684*** (0.455) | 1.7751** (0.356) | 1.0907** (0.365) | 0.4662** (0.186) |
| η_1 | -0.7923*** (0.247) | -0.1967 (0.208) | | |
| ϕ_1 | 0.4621*** (0.078) | 0.5215** (0.081) | 0.8159** (0.181) | 1.1343*** (0.109) |
| ϕ_2 | | | -0.3832** (0.166) | -0.4417*** (0.076) |
| θ_1 | 0.8223*** (0.058) | 0.9362** (0.029) | 0.8203** (0.158) | 0.5418*** (0.088) |
| θ_2 | 0.7484*** (0.059) | 0.9389** (0.025) | 0.5859** (0.159) | 0.7909*** (0.089) |
| θ_3 | 0.8244*** (0.047) | 0.9392** (0.022) | 0.7252** (0.123) | 0.5834*** (0.077) |
| ω | 0.2508 (0.197) | -0.1536** (0.077) | 0.0904** (0.039) | -0.1778 (0.281) |
| α_1 | 0.0948 (0.805) | -0.9816*** (0.029) | 0.0471 (0.241) | 0.4496*** (0.112) |
| β_1 | -0.0838 (0.094) | 0.1659 (0.188) | -0.2007** (0.048) | -0.3539*** (0.133) |
| β_2 | 0.0158 (0.091) | 0.1766 (0.285) | 0.2716** (0.091) | 0.3930*** (0.099) |
| ψ_1 | -0.0010 (0.039) | 0.0366 (0.043) | | |
| ψ_2 | | | | |
| R^2 | 0.905 | 0.911 | 0.936 | 0.939 |
| J-B | 27.632 [0.000] | 5.338 [0.069] | 0.3372 [0.845] | 0.2071 [0.902] |
| LM ARCH (1) | 0.001 [0.985] | 1.191 [0.275] | 0.1026 [0.749] | 1.0125 [0.314] |
| LM ARCH (2) | 2.529 [0.282] | 1.981 [0.371] | 1.0257 [0.599] | 3.6506 [0.161] |
| AIC | 1.5959 | 1.4591 | 1.1242 | 0.9641 |
| BIC | 1.8148 | 1.6779 | 1.5296 | 1.3696 |

Note: Bollerslev-Wooldridge robust standard errors in square brackets; p-values in brackets; * indicates statistical significance at 10% level; ** indicates statistical significance at 5% level; *** indicates statistical significance at 1% level or less. Source: Authors' calculations.

cycle asymmetry effects on volatility. The structural change in variability in 2000:1 probably reflects the change from national currencies to the euro in most of the member-states and the consequent indirect impacts on other non-European or non-euro-area states. Therefore, in attempting to analyse the time volatility evolution, we establish that date as our benchmark. The estimated results for the periods before and after the regime change in volatility, centred on 2000:1 and considering parsimonious specifications (lowest AIC

and SIC values), are reported in Table 4. Once again, both the TGARCH and EGARCH specifications provide consensual results, but we give special attention to the EGARCH specification as it reports lower values of both information criteria (AIC and BIC). The corresponding news impact curves are reported in Figure 4.

Figure 4. Impacts of positive and negative shocks on GDP growth volatility [EGARCH(1,1)]



Source: Authors' calculations

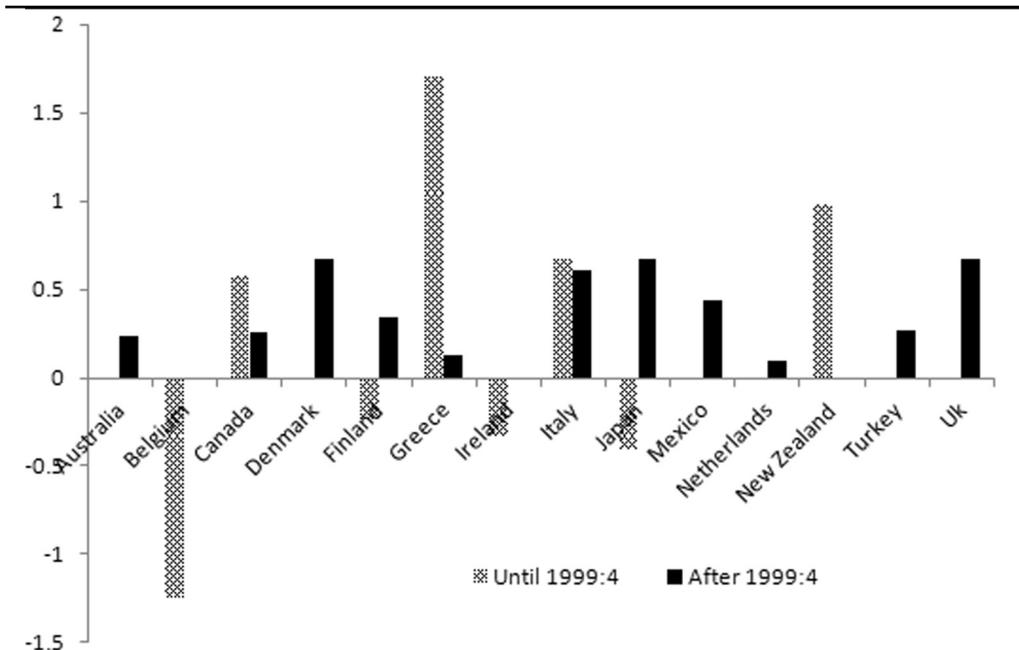
This analysis is informative on some important points. First the leverage effect is not significant over the period from 1962:1 to 1999:4; i.e., negative and positive shocks to GDP do not induce impacts of different magnitude on volatility. However, after 2000, a significant positive leverage effect is observed. Considering the EGARCH model, those impacts are 0.747 and 0.0391, respectively. This means that negative shocks have an impact of about 19 times higher than the impact of good news of the same magnitude. Therefore, there has been a change in the pattern of impacts over the sample period, with negative shocks to GDP growth inducing higher volatility than positive shocks of identical magnitude after 2000.

4.4 Country-specific contributions to volatility-asymmetry behaviour

The country-specific information for the estimation of volatility asymmetric behaviour over the two periods is reported in Table A in the Annex.¹² We iden-

tify two structural changes in both the mean and the variance in all countries. Each regime change is identified directly with the occurrence of external economic shocks, such as the oil shocks of the 1970s, the international financial crisis in the middle 1980s and, more recently, the euro crisis. Other regime changes are explained by the occurrence of internal events, such as democratic revolutions (for example in Portugal and Spain in the 1970s), or internal crisis such as the steel crisis in Luxembourg in the 1970s, or the constitutional crisis in New Zealand in the 1980s.¹³ The estimation of the leverage effects derived from the EGARCH models for each country is shown in Figure 5. In accordance with the absence of asymmetries evidence at the aggregate level, over the period until 1999:4 we found no evidence of asymmetries in fifteen countries, namely, Australia, Austria, Denmark, France, Germany, Luxembourg, Mexico, the Netherlands, Norway, Spain, Sweden, Switzerland, Turkey, the United Kingdom and the United States. However, we found evidence of asymmetries in eight countries, namely Belgium, Canada, Finland, Greece, Ireland, Italy, Japan and New Zealand.

Figure 5. Leverage effects



Source: Authors' calculations

Most of these countries still report significant leverage effects over the second period, although with smaller magnitudes. In fact, countries like Canada, Greece and Italy have shown a tendency toward effects symmetry, while Belgium, Ireland and New Zealand have managed to reach balanced

effects. Therefore, the economic vulnerability to negative shocks seems to have been reduced over time. On the other hand, Finland and Japan continue to report significant asymmetries, while other countries like Australia, Denmark, Mexico, the Netherlands, Turkey and the United Kingdom have become more vulnerable to negative shocks after 2000. The two largest European economies, Germany and France, the northern economies, Norway and Sweden, together with Luxembourg, Spain, Switzerland, Austria and the United States have managed to get balanced effects over the sample period.

5. CONCLUSIONS

This paper investigates the asymmetric effects of external shocks on real GDP growth in the OECD, using quarterly data over the last five decades in the context of 'the Great Moderation' phenomenon, characterised by a declining trend in both growth and volatility across almost all the member-states. The absence of information on this issue for the OECD at the aggregate level, as well as for the individual member-states, together with the lack of consensus in the literature about the behaviour of volatility across the business cycle, attributed mostly by methodological issues, are open points in the research agenda that constitute a motivation for this study.

We adopt a GARCH modelling strategy, which accounts for the occurrence of regime changes in both the trend and volatility of GDP series, to identify signs of "the Great Moderation" in the OECD, the time-varying nature of volatility and its symmetric/asymmetric nature across the business cycle and over the sample period.

The results reveal a progressive moderation, both at the aggregate and disaggregated levels, characterised by a decline in both GDP growth rates and associated volatility. Asymmetries in the behaviour of growth volatility seem to emerge over the business cycle. The results suggest that periods of positive growth are characterised by a positive relationship between growth rates and volatility, while periods of negative growth are characterised by a negative relationship. We estimate that the impacts of negative shocks on volatility exceed by far those of positive shocks over the sample period. However, this asymmetric pattern is not stable over time and the time-disaggregated analysis uncovers a pattern of increasing asymmetry, which may provide a sign of increased economic vulnerability to exogenous negative shocks. The increase in the persistence of this asymmetry is observed particularly when the analysis is performed considering two different periods, before and after 2000:1. This date is particularly relevant since it is associated with the currency change in the euro-zone member-states. The general equilibrium of the effects caused by positive and negative shocks before that date gives place to a disequilibrium in the following period, with the effects of negative shocks to exceed those of positive shocks by more than 19 times.

A detailed analysis on a country-by-country basis reveals decreasing vulnerability to negative shocks in Canada, Greece, Italy and New Zealand,

but increasing vulnerability in Australia, Denmark, Finland, Japan, Mexico, the Netherlands, Turkey and the United Kingdom. The largest European economies, together with the northern economies, the United States and the wealthiest economies of Luxembourg and Switzerland are among those that were able to keep their resilience to the effects of negative shocks.

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ENDNOTES

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3. The authors thank the anonymous reviewers for the insightful comments and helpful suggestions. The authors also thank to financial support from Fundação para a Ciência e Tecnologia and FEDER/COMPETE (grant Pest -C/EGE/UI4007/2011).
4. Australia, Austria, Belgium, Canada, Chile, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Japan, South Korea, Luxembourg, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Slovenia, Spain, Sweden, Switzerland, Turkey, the United Kingdom and the United States of America.
5. We thank an anonymous referee for suggesting this test.
6. Model B is omitted as it is commonly held that most economic time series can be adequately described by models A or C.
7. The corresponding null and alternative hypothesis for Model A are particular cases of those for Model C, considering d_3 and d_4 equal to zero.
8. The authors also show that the asymptotic null distribution of the two break LM unit root test for Model A is invariant to the location and magnitude of the structural breaks.
9. Whenever the null is not rejected, the test should be repeated for Model A.
10. Results of the model selection are not provided here, but are available upon request.
11. The BDS test was developed by Brock, Dechert and Scheinkman (1987) and later published as Brock, Dechert, Scheinkman and LeBaron, (1996).
12. Some countries have been excluded from the analysis because of limited data:

Chile, Czech Republic, Estonia, Hungary, Iceland, Israel, Slovak Republic and Slovenia. Details on model specification for each country, as well as regime changes in both the mean and variance are available from the authors by request.

13. Here we report a short and general explanation of the identified structural changes timings. A more detailed description can be provided upon request.

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