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By

Kyriakos C. Neanidis and Christos S. Savva

Centre for Growth and Business Cycle Research, Economic Studies,
University of Manchester, Manchester, M13 9PL, UK

July 2010
Number 145

Macroeconomic Uncertainty, Inflation and Growth: Regime-Dependent Effects in the G7*

Kyriakos C. Neanidis[a] and Christos S. Savva[b]

[a] *Economics, University of Manchester and Centre for Growth and Business Cycle Research, Manchester, United Kingdom*
[b] *Department of Commerce, Finance and Shipping, Cyprus University of Technology, Cyprus*

First completed draft: June 24, 2010
This version: June 24, 2010

Abstract

We analyze the causal effects of real and nominal macroeconomic uncertainty on inflation and output growth and examine whether these effects vary with the level of inflation and location on the business cycle. Employing a bivariate Smooth Transition VAR GARCH-M model for the G7 countries during the period 1957-2009, we find strong nonlinearities in these effects. First, uncertainty regarding the output growth rate is related with a higher average growth rate mostly in the low-growth regime, supporting the theory of "creative destruction". Second, higher inflation uncertainty induces lower growth rates, increasingly so at the high-inflation regime. Third, real and nominal uncertainties have mixed effects on average inflation. Nevertheless, there is a trend in favour of the Cukierman-Meltzer hypothesis in the high-inflation regime. Our results can be viewed as offering an explanation for the often mixed and ambiguous findings in the literature.

Keywords: Business cycles; inflation; output growth; uncertainty; regimes; GARCH model

JEL Classification: C32; C51; E31; E32; O40

* We are grateful to Denise Osborn for helpful discussions and comments on a preliminary draft.

1. Introduction

Considerable research efforts have been directed through the years among macroeconomists on the link between inflation and output growth.[1] More recently, however, attention has been focused on the effects of macroeconomic uncertainty on both inflation and output growth.[2] On both theoretical and empirical fronts, macroeconomic uncertainty, defined as nominal (volatility of inflation) and/or real (volatility of output growth), has been found to have contradictory effects on the mean values of inflation and output growth providing inconclusive outcomes of its average impact.[3] For this reason, a few studies have attempted to investigate whether these effects are contingent to a particular set of conditions, such as the level of economic and financial development, institutional development, and trade openness (see Hnatkovska and Loayza 2004). The current study, following this line of thought, contributes to the literature by focusing on the regime-dependent effects of real and nominal macroeconomic uncertainty on inflation and output growth. Our innovation lies in that the regimes are determined by the rates of inflation and output growth themselves, consistent with studies that stress the importance of the state of the economy – level of inflation rate and phase of business cycle – for the transmission of such effects (see Baillie et al. 1996; Henry and Olekalns 2002; Chang and He 2010).

The theoretical literature that examines the effects of macroeconomic uncertainty on inflation and growth is rich both in terms of the number of explanations offered and of the sign of the assessed impact.[4] In particular, there are some theories that support an effect of inflation uncertainty on output growth that takes up a negative sign (Friedman 1977; Pindyck 1991) while some others a positive sign (Abel 1983; Dotsey and Sarte 2000; Blackburn and Pelloni 2004). Similarly, the influence of nominal uncertainty on the rates of inflation has been deemed to be positive by Cukierman and

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[1] For early theoretical models, Tobin (1965) shows that inflation raises growth, while Stockman (1981) derives a diminishing effect. Finally, Sidrauski (1967) depicts the super-neutrality of inflation. More recent models in support of a negative effect include Gomme (1993) and Jones and Manuelli (1995). Early empirical papers, all in support of the negative impact, include among others De Long and Summers (1992), De Gregorio (1993), and Barro (1996). Few other studies, however, illustrate a nonlinear impact of inflation on output growth (Sarel 1996; Bruno and Easterly 1998; Khan and Senhadji 2001).

[2] There are also studies that examine the reverse direction of causality, that is, the impact of the rates of inflation and of economic growth on their respective volatilities. See, for example, Grier and Perry (1998), Fountas (2001), Conrad and Karanasos (2005), Fountas and Karanasos (2006).

[3] Even though volatility (fluctuations in a variable) and uncertainty (unpredictability of fluctuations) are two different concepts, it is common practice to use them interchangeably.

[4] A detailed description of these theories is presented in the following section.

Meltzer (1986) and Cukierman (1992) and negative by Holland (1995). In a same fashion, output growth uncertainty has had supporters of a positive (Deveraux 1989) but also of a negative (Taylor effect jointly with the Cukierman-Meltzer hypothesis) influence on inflation. Finally, the impact of real uncertainty on output growth has been theorized to go in all possible directions covering positive (Black 1987), zero (Friedman 1968), and negative (Bernanke 1983; Pindyck 1991) effects. There have also been some studies that relate the sign of the correlation on the size of the coefficient of relative risk aversion (de Hek 1999; Jones et al. 2005), the type of the learning behavior that guides productivity improvements (Blackburn and Galindev 2003), the impulse source of fluctuations (Blackburn and Pelloni 2004), and the behavior of agents toward working versus learning (Blackburn and Varvarigos 2008).

A sizable empirical literature investigating the above effects can be broadly categorized into two groups of studies. On the one hand, there are studies that are based on cross-section or panel data approaches. On the other, the analysis is conducted with the use of time series data within univariate and multivariate Generalized Autoregressive Conditional Heteroskedasticity (GARCH) techniques.[5] The findings of both classes of studies, however, produce mix and often contradictory results in line with the conflicting theoretical argumentation.

The cross-section and panel data approaches have mainly focused on the effects of nominal and real uncertainty (proxied by their variability) on output growth rates, with more weight on the latter relationship. Judson and Orphanides (1999), using annual data from 1959 to 1992 for 119 countries, find that both inflation and inflation uncertainty lower growth. Clark (1997), on the other hand, shows that neither average inflation nor inflation volatility is robustly related to economic growth. The effect of real uncertainty on growth is even more mixed as there are studies that support a positive effect (Kormendi and Meguire 1985; Grier and Tullock 1989; Dejuan and Gurr 2004), a negative effect (Ramey and Ramey 1995; Kneller and Young 2001; Hnavotska and Loayza 2003), and even a zero effect (Dawson and Stephenson 1997;

[5] Cross-section and panel data studies have traditionally used the variance (or standard deviation) of a variable as a measure of its uncertainty. This, however, being a measure of a variables' variability, rather than uncertainty, has been criticized. Following the development of the ARCH approach by Engle (1982), further studies have mostly adopted a time-varying conditional variance as the measure of uncertainty.

Dawson et al. 2001; Chatterjee and Shukayev 2006) with a variety of country and time period sets.

The class of studies that utilizes time series techniques encompasses a greater number of papers that examines a greater number of testable relations. The first technique is the univariate GARCH method where in the first step the conditional variances of inflation and output growth are estimated independently from each other and in the second step Granger causality tests are performed to examine the relationships between macroeconomic uncertainty and performance (see Grier and Perry 1998; Fountas and Karanasos 2006). Two-step estimation methodologies, however, may encounter the problem of a "generated regressor", as suggested by Pagan (1984), which leads to biased estimates of the standard errors and hence to problems in inference. For this reason, a lot of studies adopt a simultaneous approach where a bivariate GARCH-in-mean (GARCH-M) model is estimated to provide estimates of the conditional variances and at the same time test for the impact of uncertainty on macroeconomic performance (Grier and Perry 2000; Grier et al. 2004; Fountas and Karanasos 2007). The countries under consideration in these studies, however, is more limited compared to the cross-country approach as the need for relatively long time series limits the set typically to industrialized nations, with more emphasis in the G7 countries.[6]

Focusing on the findings of the studies that utilize *simultaneous* bivariate GARCH-M models, the sign of the effects once again varies – even with the use of more homogeneous country sets. Table 1 presents the coefficient estimates of the most popular studies for the four kinds of effects under consideration.[7] We use the usual terminology in the literature and coin the coefficients of interest as Ψ_{ij} with $i,j = 1,2$. Ψ_{11} and Ψ_{21} describe respectively the impact of output growth uncertainty on output growth and the inflation rate, while Ψ_{12} and Ψ_{22} reflect the impact of inflation uncertainty on output growth and the inflation rate, respectively. Starting with the United States, which is the most studied country, Table 1 shows that the three most well-known studies unanimously agree only on the negative impact of inflation

[6] A notable exception to this rule is a recent study by Bredin et al. (2009) who use a bivariate GARCH-M model for a set of South-East Asian countries.
[7] Coefficients in bold type represent significance at least at the 5% level.

uncertainty on output growth (Ψ_{12}), even though the magnitude of the effect differs. The significance of the remaining three estimates varies across studies.[8] This situation, however, is not limited to the United States. For Japan, Bredin and Fountas (2005) and Wilson (2006) agree on the sign and significance of Ψ_{21} only. The same is also true for the United Kingdom, but now we also observe differences in the direction of the effects described by Ψ_{11} and Ψ_{12} which appear to be significant in all three cases. For Germany, the agreement is restricted on the insignificance of Ψ_{21}, while for France Bredin and Fountas (2005) show all four effects to be significant in contrast to the findings of Bredin and Fountas (2009) that find them to be statistically equal to zero. Finally, Italy resembles the cases of Japan and the United Kingdom in that only one of the effects shows up with a consistent sign and significance, Ψ_{22}. One, therefore, concludes that the results of the literature are at best mixed and often antithetical.

The above conflicting evidence of the effects of macroeconomic uncertainty on macroeconomic performance, even for the same countries and similar period sample, suggests that the effects may differ because of the presence of nonlinearities or regime shifts. This is corroborated by a few recent papers that examine the conditional or regime-dependent impact of real and nominal uncertainties. Such studies can be found both in the panel data and time-series literature.

In a cross-country environment, a few recent studies support a nonlinear impact of real volatility on long-run economic growth. Hnatkovska and Loayza (2004) illustrate that, for 79 countries over the period 1960-2000, growth volatility reduces economic growth in countries that are economically and institutionally underdeveloped and undergoing intermediate stages of financial development. They also show that the negative effect is mostly due to large recessions. They attribute these findings to the lack of stabilization policies, institutional safeguards, and insurance markets in less advanced nations that would allow them to neutralize the long-run effects of volatility. A different view is offered by Herrero and Vilarrubia (2005) who offer support to an inverted U-shaped relationship between real volatility and growth. Using a set of more than 100 countries during the period 1978-2002, they find

[8] It is interesting to note that all three studies use the same variables to proxy output and prices (industrial production index and producer price index respectively) at the same frequency (monthly). Furthermore, the period coverage of the three studies does not vary considerably.

evidence that a moderate degree of volatility is growth-improving while high-volatility is detrimental for growth.

The conditional effect of real uncertainty on output growth has also been examined with reference to business cycle movements. Kroft and Lloyd-Ellis (2002) use the original Ramey and Ramey (1995) dataset but decompose aggregate volatility into fluctuations that may be interpreted as year-to-year uncertainty and fluctuations that reflect structural business cycle shifts between recessions and expansions. They find a significant negative correlation between growth and structural business cycle fluctuations, and a significant positive correlation between growth and year-to-year fluctuations. Henry and Olekalns (2002) focus explicitly on the effect of recessions on the relationship between output variability and growth for the US in a GARCH-M framework. They provide evidence that recessions result in increased output uncertainty, which dampens subsequent growth. As the economy expands, the impact of real uncertainty on growth vanishes.

Using the GARCH approach, the analysis also considers the impact of nominal uncertainty. Baillie et al. (1996) show that higher uncertainty about the rate of inflation leads to higher inflation rates only in high-inflation countries. Chen et al. (2008), on the other hand, in an examination of the four little dragons (Hong-Kong, Singapore, South Korea, and Taiwan), show that inflation uncertainty induces a “Laffer curve” effect on inflation in the case of Taiwan. Finally, Chang and He (2010) examine the effect of inflation uncertainty on output growth as a function of the rate of inflation. They employ a bivariate Markov regime switching model for the US economy during 1960Q1-2003Q3 and demonstrate that nominal uncertainty inhibits growth in both low-and high-inflation regimes. The size of the effect, however, is greater in the high-inflation regime by threefold. Importantly, when there is no distinction between regimes, the use of a single-regime GARCH model shows the effect of inflation uncertainty to be negligible, pointing to a misspecification bias.[9]

[9] A related study by Bredin and Fountas (2009) investigates for regime switching in the conditional variance of inflation and output growth in 14 European Union countries via the estimation of Switching ARCH in-mean (SWARCH-M) models. In spite of the presence of significant regime switching in both the volatilities of inflation and growth, they do not find any significant effects of macroeconomic uncertainty on performance. They state, however, that these findings could be misleading as they test for regime switching in a univariate setting rather than (the most appropriate) multivariate.

Following the terminology of the Ψ_{ij} terms above, these GARCH-related studies imply non-monotonic effects of real uncertainty on output growth (Ψ_{11}) and of nominal uncertainty on both output growth (Ψ_{12}) and inflation (Ψ_{22}). The findings of these set of papers is the starting point of the current paper, where we consider the impact of nominal and real uncertainty as being dependent on the rates of inflation and output growth respectively. The use of the rate of inflation as the measure by which inflation uncertainty influences inflation and growth is adopted by Baillie et al. (1996) and Chang and He (2010). The use of output growth rates as the transition variable in estimating the impact of growth rate uncertainty on inflation and output growth draws from a long literature that uses this variable to distinguish between periods of positive and negative growth, that is, between expansions and contractions (see, among others, Teräsvirta and Anderson 1992; Teräsvirta et al. 1994; Van Dijk et al. 2002).[10]

Keeping with the above reasoning, our methodology employs a bivariate Smooth Transition VAR GARCH-M model that is used to generate the conditional variances of inflation and output growth as proxies of inflation and output growth uncertainty, respectively. These measures are then used to test for the effect of real and nominal uncertainty on inflation and output growth. The innovation in our technique is that we generalize the conditional mean and conditional variance to allow for the possibility that they are affected by the state of the economy as captured by the rate of inflation and the phase on the business cycle. An advantage of this specification is that the regime is recognized at each time point by using the smoothed transition process without exogenously assuming structural change points and threshold values. In this way, our technique allows us to obtain more reliable and accurate estimates of macroeconomic uncertainty. This offers an important advantage compared to traditional GARCH models where macroeconomic uncertainty is by construction invariant to the direction of change in inflation and output growth. Our dataset covers the G7 countries with monthly data for the period 1957-2009.

Our findings offer strong support to the presence of threshold, or regime-dependent, effects for most of the hypotheses tested. First, there is evidence that output growth

[10] More studies that follow this procedure, in models that examine the effectiveness of monetary policy over the business cycle, include Garcia and Schaller (1995), Weise (1999), and Sensier et al. (2002).

uncertainty enhances mean output growth (Ψ_{11}). The effect, however, is mainly restricted in the low-growth regime during economic contractions. Second, with the exception of Italy, higher inflation uncertainty leads to lower mean output growth (Ψ_{12}). This effect becomes especially pronounced during periods of high inflation. Third, output growth uncertainty does not appear to have a significant effect on mean inflation (Ψ_{21}) across the G7. There is weak evidence though of a diminishing negative, or increasing positive, effect as moving to the high-growth regime. Fourth, there are mixed findings for the effect of inflation uncertainty on mean inflation (Ψ_{22}) being mostly positive. At the same time, there is a trend for an amplifying impact as economies move to the high-inflation regime. These results can be viewed as offering an explanation for the often mixed and ambiguous findings of the literature on the effects of real and nominal uncertainty on the rates of inflation and output growth.

The paper is organized as follows. Section 2 offers the analytical background to the theories that relate macroeconomic uncertainty and performance. Section 3 presents the dataset while Section 4 outlines our econometric model. In Section 5 we report the results of the single-regime model and of the regime-dependent model. These results are then discussed in the context of some recent studies. Finally, Section 6 summarizes our conclusions.

2. Analytical background

Concerning the impact of inflation uncertainty on output growth, Friedman (1977) argues informally that increased inflation uncertainty reduces the effectiveness of the price mechanism in allocating resources efficiently and hinders long-term contracting, thus reducing output growth. In a more formal framework, Pindyck (1991) and Huizinga (1993) show that inflation uncertainty reduces output growth by increasing the option value of delaying an irreversible investment, thus causing investment projects to be postponed. Blackburn and Pelloni (2004), using a model with nominal rigidities, also argue that nominal variability exerts a negative effect on growth, channelled through its adverse impact on aggregate employment. In contrast, Abel (1983) in a model that allows for symmetric adjustment costs of investment show that inflation uncertainty raises investment and growth. Similarly, Dotsey and Sarte (2000), using a cash-in-advance model with risk-averse agents, predict that more inflation uncertainty leads to higher output growth. This outcome is a result of higher precautionary savings in response to nominal uncertainty, which in turn leads to higher investment. More recently, Varvarigos (2008) supports the finding of Dotsey and Sarte (2000) through a human capital accumulation channel.

The impact of inflation uncertainty has also been investigated with respect to the rate of inflation. Cukierman and Meltzer (1986) and Cukierman (1992) develop a game-theoretic model of central bank behavior according to which higher inflation uncertainty raises the average inflation rate. Using the Barro-Gordon model, where agents face uncertainty about both the rate of money supply growth (and hence inflation) and the policy-maker's objective function, they show that an increase in uncertainty about money growth and inflation raises the optimal average inflation rate because it provides an incentive to the policymaker to create an inflation surprise in order to stimulate output growth. Thus the lack of a commitment mechanism to control the inflation rate produces an inflationary bias in equilibrium. By contrast, Holland (1995) argues that inflation uncertainty has a negative impact on the inflation rate owing to the central bank's stabilizing policy. With an independent central bank and a clear commitment to long-run price stability, monetary authorities when faced with more inflation uncertainty apply tight monetary policy, and hence reduce average inflation, in order to minimize the real costs of inflation uncertainty.

Moving to the macroeconomic effects of real uncertainty, uncertainty about output growth has been argued to influence both the rates of inflation and output growth. Deveraux (1989) based on the Barro-Gordon model shows that real uncertainty increases the average rate of inflation. In a model with endogenous wage indexation and a stochastic element to money growth, Deveraux (1989) demonstrates that an exogenous increase in the variability of real shocks causes workers to lower the optimal amount of wage indexation. From the central bank's perspective, lower indexation invites surprise inflation as a more effective instrument to raise output. In equilibrium, therefore, the increased incentive to inflate translates into a higher average inflation rate. Higher output uncertainty, however, may also lead to a lower average inflation rate. This channel combines the Taylor effect with the Cukierman-Meltzer hypothesis outlined above. As the Taylor effect suggests a negative association between output variability and inflation variability and the Cukierman-Meltzer hypothesis illustrates a positive effect of inflation variability on average inflation, their combination yields a negative impact of output uncertainty on the rate of inflation.

Finally, real uncertainty has been described to have diverse effects on output growth. Prior to the 1980s, the study of growth and business cycles was two separate bodies of literature. Since then they have been brought together in a variety of setups. These recent macroeconomic theories offer explanations for a positive, a zero, or a negative effect of output uncertainty on output growth rates. The positive effect is put forth by Black (1987) who argues for a positive trade-off between aggregate risk and return in the choice of economies for productive technologies. As such, investments in riskier technologies take place only if the expected rates of return (average growth rates) are large enough to compensate for the greater risk. Sandmo (1970) and Mirman (1971) also support a positive link based on the theory of saving under uncertainty. Higher real uncertainty causes higher precautionary savings and subsequent rates of investment that positively impact upon output growth. Another class of models, business cycle models based on the natural rate hypothesis, suggest no relationship between output variability and growth. Implicit in Friedman's (1968) business cycles model, output movements away from the natural rate are the result of price level misperceptions by workers and firms triggered by monetary shocks. Long-run output growth, being determined by real factors (human and physical capital accumulation), is independent of these information asymmetries.

The negative effect of output uncertainty on output growth has been theoretically derived by a few studies. Keynes (1936) argues that in the presence of large fluctuations in economic activity, entrepreneurs perceive investment projects as riskier. This, in turn, lowers the demand for investment and output growth. Bernanke (1983) and Pindyck (1991) relate instead the decline in investment and growth that arises from increased real uncertainty to investment irreversibilities at the firm level. Firm's investment in plant and equipment falls as higher volatility in output growth leads to unexpected changes in output growth and makes future demand for a firms' product more uncertain (increased riskiness). A final explanation is given by Ramey and Ramey (1995) who relate the negative causal link of output uncertainty on growth to the firm's commitment on technology in advance. In such an environment, higher real variability leads firms to produce at ex-post suboptimal levels due to uncertainty-induced planning errors, and hence lower mean output and growth.

There is also a more recent line of literature that uses endogenous growth models to identify the sign of the correlation between output growth variability and output growth. These papers take the view that the correlation could be either positive or negative, the result being a function of the fundamentals governing the behavior of agents and structural characteristics of the economy. These include the agents' attitudes toward risk, their preference for learning, the type of technology shocks, and the characteristics of fluctuations. Smith (1996), de Hek (1999), and Jones et al. (2005) argue that the effect of volatility on growth depends on the magnitude of the elasticity of relative risk aversion. For sufficiently high (low) degrees of risk aversion, an increase in volatility causes an increase (decrease) in precautionary investments in physical or human capital, implying an increase (decrease) in long-run growth. In another class of models it is the mechanism of technological change that is important for the results (e.g., Aghion and Saint-Paul 1998a, b; Martin and Rogers 2000; Blackburn and Galindev 2003). If this mechanism is based on internal (external) learning, then an increase in volatility leads to an increase (decrease) in the amount of learning that takes place, generating an increase (decrease) in trend growth. Blackburn and Pelloni (2004) in a stochastic monetary growth model show that the correlation between output growth and its variability is a function of the type of shocks buffeting the economy. The study concludes that the correlation will be positive (negative) depending on whether the real (nominal) shocks dominate. Finally, Blackburn and Varvarigos (2008) offer another determinant that guides the effect of output growth variability on output growth vis-à-vis the presence of preference and technology shocks. They show that when the disutility of total effort devoted to non-leisure activities is relatively low (high), the effect in the case of preference shocks is more likely to be positive (negative). This outcome is contrasted by the case of technology shocks for which the effect is always positive.

3. Data Description

We use monthly data from 1957.01 through 2009.08 on industrial production index (IPI) and producer price index (PI) as proxies for output and the price level respectively.[11] The data refer to the G7 and the source is the International Financial Statistics (International Monetary Fund). Inflation is defined as the annualized monthly difference in the log producer price index [$\pi_t=log(PI_t/PI_{t-1})*1200$] and similarly output growth is calculated as the annualized monthly difference in the log industrial production index [$y_t=log(IPI_t/IPI_{t-1})*1200$].

Summary statistics and various tests on inflation and output growth have been performed. We first test for the stationarity properties of our data using ADF and Philips and Perron tests. The results of these tests indicate that the null hypothesis of a unit root is rejected at the 1% level indicating that all series are stationary. Descriptive statistics on both series include results on skewness, kurtosis and Jarque-Bera normality tests which provide evidence against normality in inflation and output growth in all G7 countries.

In addition to the above, the Ljung-Box test for up to twelve lags serial correlation indicates the strong presence of serial dependence in the data. Similarly, a Ljung-Box test for serial correlation in the squared data provides evidence of conditional heteroskedasticity. These tests suggest that second moment (nonlinear) dependencies are significant, supporting the use of a specification that captures the instability of the variances of inflation and output growth.[12]

4. Econometric Methodology

In this section we introduce the bivariate Smooth Transition VAR GARCH-M model with constant conditional correlations (CCC) which allows us to model and test the effects of inflation uncertainty and growth uncertainty on the levels of inflation and output growth along different regimes. [13]

[11] Both the IPI and PI data are seasonally adjusted. Price level data for France and Italy refer to the consumer price index due to the lack of early data on PI.

[12] All descriptive statistics and tests are available from the authors upon request.

[13] In the CCC model (Bollerslev 1990) the conditional covariance matrix is time-varying but the conditional correlation across equations is assumed to be constant. We opt to use the CCC specification since the related literature revealed low and stable conditional correlations between growth and inflation. In addition, the assumption of a constant matrix reduces the estimated parameters and computational complexity (see, for instance, Grier and Perry 2000; Wilson 2006; Fountas et al. 2002 and 2006, among others).

Consider a bivariate time series of output growth (y_t) and inflation (π_t), $t=1,.....,n$, the stochastic properties of which are assumed to be described by the model:

$$\begin{aligned} y_t = &\left(\phi_{10,l} + \sum_{\kappa=1}^{p} \phi_{11,l}^{k} y_{t-k} + \psi_{11,l}\sqrt{h_{y_t}}\right)\left[1 - G\left(y_{t-1}^{*};\gamma_1,c_1\right)\right] + \\ &\left(\phi_{10,h} + \sum_{\kappa=1}^{p} \phi_{11,h}^{k} y_{t-k} + \psi_{11,h}\sqrt{h_{y_t}}\right) G\left(y_{t-1}^{*};\gamma_1,c_1\right) + \\ &\left(\sum_{\kappa=1}^{p} \phi_{12,l}^{k} \pi_{t-k} + \psi_{12,l}\sqrt{h_{\pi_t}}\right)\left[1 - G\left(\pi_{t-1}^{*};\gamma_2,c_2\right)\right] + \\ &\left(\sum_{\kappa=1}^{p} \phi_{12,h}^{k} \pi_{t-k} + \psi_{12,h}\sqrt{h_{\pi_t}}\right) G\left(\pi_{t-1}^{*};\gamma_2,c_2\right) + u_t \end{aligned} \tag{1}$$

$$\begin{aligned} \pi_t = &\left(\sum_{\kappa=1}^{p} \phi_{21,l}^{k} y_{t-k} + \psi_{21,l}\sqrt{h_{y_t}}\right)\left[1 - G\left(y_{t-1}^{*};\gamma_1,c_1\right)\right] + \\ &\left(\sum_{\kappa=1}^{p} \phi_{21,h}^{k} y_{t-k} + \psi_{21,h}\sqrt{h_{y_t}}\right) G\left(y_{t-1}^{*};\gamma_1,c_1\right) + \\ &\left(\phi_{20,l} + \sum_{\kappa=1}^{p} \phi_{22,l}^{k} \pi_{t-k} + \psi_{22,l}\sqrt{h_{\pi_t}}\right)\left[1 - G\left(\pi_{t-1}^{*};\gamma_2,c_2\right)\right] + \\ &\left(\phi_{20,h} + \sum_{\kappa=1}^{p} \phi_{22,h}^{k} \pi_{t-k} + \psi_{22,h}\sqrt{h_{\pi_t}}\right) G\left(\pi_{t-1}^{*};\gamma_2,c_2\right) + v_t \end{aligned} \tag{2}$$

The VAR structure of equations (1) and (2) allows for the possibility of bivariate Granger causality between growth and inflation along with their uncertainties, where subscripts l and h denote the low and high regime respectively. $k(=1, \ldots, p)$ determines the lag order and captures any possible mean effects that arise from past growth and/or inflation.[14] To assess whether the effects on growth and inflation vary between low or high regimes, we employ a continuous function $G(s_{t_i};\gamma_i,c_i)$, which changes smoothly from 0 to 1 as s_{t_i} (the transition variable) increases.[15] A popular choice for the transition function is the logistic function[16]

[14] The lag order is determined by the Schwartz Information Criterion (SIC).
[15] In practice we could have used a threshold model rather than a Smooth Transition model. Nevertheless we opt to use a Smooth Transition model as, in principle, it is more general and includes the threshold model as a limiting special case.
[16] The starting values of γ_i and c_i (with $\gamma_i > 0$) are determined by a grid search and are estimated in one step by maximizing the likelihood function.

$$G(s_{t_i};\gamma_i,c_i) = \frac{1}{1+\exp\left(-\gamma_i\left(s_{t_i}-c_i\right)\right)} \qquad (3)$$

The way the model is written in equations (1) and (2) highlights the basic characteristic of the logistic specification in the model, which is, that at any given point in time, the evolution of y_t and π_t (and their explanatory variables) are determined by a weighted average of two different autoregressive models. The weights assigned to the two models depend on the value taken by the transition variable s_{t_i}. For small (large) values of s_{t_i}, $G(s_{t_i};\gamma_i,c_i)$ is approximately equal to zero (one) and, hence, almost all weight is put on the first (second) part of the model.

The parameter c_i, where i=1 (growth), 2 (inflation) can be interpreted as the threshold between the two regimes corresponding to $G(s_{t_i};\gamma_i,c_i)=0$ and $G(s_{t_i};\gamma_i,c_i)=1$, in the sense that the logistic function changes monotonically from 0 to 1 as s_{t_i} increases, while $G(c_i;\gamma_i,c_i)=0.5$. A nice feature of our specification is that we let the model decide the threshold points between low and high regimes of growth and inflation.

The parameter γ_i where i=1 (growth), 2 (inflation) determines the speed at which the weights between the two parts of the specification change as s_{t_i} increases; the higher γ_i, the faster is this change. If $\gamma_i \rightarrow 0$, the weights become constant (and equal to 0.5) and the model becomes linear, whereas, if $\gamma_i \rightarrow \infty$, the logistic function approaches a Heaviside function, taking the value of 0 for $s_{t_i} < c_i$ and 1 for $s_{t_i} > c_i$.

As far as the transition variable (s_{t_i}) is concerned, in practice, the appropriate transition variable is unknown, however a good choice is to use lagged endogenous variables.[17] In our case we opt to use $s_{t_i} = y^*_{t-1}$ or π^*_{t-1} where y^*_{t-1} and π^*_{t-1} denote the lag smoothed growth and inflation rates (averaged over the last twelve months) respectively.[18] In the determinants of equations (1) and (2), all variables related to

[17] This logistic form has been widely used for smooth transition models. For further details we refer to Terasvirta and Anderson (1992), Terasvirta (1994), and van Dijk and Franses (1999).
[18] The transformation effectively smoothes the changes and has been found to be more appropriate for capturing "regimes" than the more noisy monthly changes (see also van Dijk et al. 2002).

growth – constant, past levels, and uncertainty – depend on the function that uses smoothed growth as a transition variable while all variables related to inflation depend on the function that uses smoothed inflation as a transition variable. An interesting feature of this model is that the two variables do not need to be in the same state (low/high regime) at the same time. This further reinforces the independence of the regimes.

Typically the literature allows one transition variable to potentially affect all variables in equations (1) and (2), and then apply restrictions supported by the data. In our model, this would translate into using either the rate of growth or of inflation as the single transition variable in both equations. This would be justifiable as long as the regimes implied by inflation and growth are similar. In practice, however, these two variables may give rise to different regimes. It is not difficult to think of situations where growth and inflation move in opposite directions making the regime determined by each of the transition variables vary. But even if both variables move in the same direction, differences in their size of change may yield different regimes. For these reasons a more natural hypothesis is to assume that the growth transition applies to lags of growth and the inflation transition applies to lags of inflation. In the next section we offer evidence for the presence of the form of nonlinearity assumed in the model based on a Lagrange Multiplier (LM) statistical test.

In equations (1) and (2), the coefficients $\phi_{ij,l}^{k}$ (where *i,j=1,2* and *k=1, ..., p)* capture the effects of lagged growth and inflation on mean growth and inflation respectively when smoothed growth and inflation are below the threshold c_i, while $\phi_{ij,h}^{k}$ capture the effects of lagged growth and inflation when smoothed growth and inflation are above the threshold c_i. Coefficients $\psi_{ij,l}$ and $\psi_{ij,h}$ do the same for the case of uncertainties, while $\phi_{j0,l}$ and $\phi_{j0,h}$ follow the same pattern for the constants. We test for the statistical difference of the "pairs" of estimated coefficients in the two regimes with a Likelihood Ratio (LR) test.[19] This test allows us to observe whether the change of a

[19] The LR test is used to compare the unrestricted model with the restricted one ($LR=-2[\ln L_R - \ln L_{UR}]$). The restricted model imposes the restriction of equal regimes (i.e. $\psi_{ij,l} = \psi_{ij,h}$). Asymptotically, the test statistic is chi-squared distributed, with degrees of freedom equal to the difference in the number of parameters between the two models.

regime is accompanied by a statistically dissimilar effect on mean inflation and output growth.

In order to capture any temporal effects in the error volatilities, the error process of equations (1) and (2) is assumed to follow the process

$$\varepsilon_t \mid \Psi_{t-1} \sim N(0, H_t) \tag{4}$$

where $\varepsilon_t = \{u_t, v_t\}$, Ψ_{t-1} is the information set consisting of all relevant information up to and including time $t-1$, and N denotes the bivariate normal distribution. The conditional covariance matrix of ε_t, H_t, is assumed to follow a time-varying structure given by

$$H_t = E[\varepsilon_t \varepsilon_t' \mid \Psi_{t-1}] \tag{5}$$

$$h_{1,t} = \omega_{1,l}\left[1 - G\left(y^*_{t-1}; \gamma_1, c_1\right)\right] + \omega_{1,h} G\left(y^*_{t-1}; \gamma_1, c_1\right) + \alpha_1 u^2_{1,t-1} + \beta_1 h_{1,t-1} \tag{6}$$

$$h_{2,t} = \omega_{2,l}\left[1 - G\left(\pi^*_{t-1}; \gamma_2, c_2\right)\right] + \omega_{2,h} G\left(\pi^*_{t-1}; \gamma_2, c_2\right) + \alpha_2 v^2_{2,t-1} + \beta_2 h_{2,t-1} \tag{7}$$

$$h_{12,t} = \rho(h_{1,t} h_{2,t})^{1/2} \tag{8}$$

where we assume that the conditional variances $h_{1,t}$ and $h_{2,t}$ both follow a GARCH(1,1) specification. Our choice is motivated by the heavy autocorrelation of the second moments, as indicated by the diagnostic tests, and by the empirical literature that has found this specification to adequately capture the persistence in second moments of industrial production and producer/consumer price indices (e.g., Fornari and Mele 1997; Grier and Perry 2000; Grier et al. 2004). Similar to the parameters in equations (1) and (2), the constant parameters, ω, of equations (6) and (7) are allowed to vary when the transition variable takes values below or above the threshold value. We incorporate this nonlinearity in our GARCH model to avoid a misspecification problem and at the same time to examine for its validity, even though the possibility of regime-dependent variances is almost always ignored in a Smooth-Transition framework.[20] It is worth mentioning that our regime-dependent GARCH(1,1) model allows the intercepts, but not other coefficients, to vary. This is due to the feasibility of the estimation.

[20] Note that the Markov-switching literature generally supports the presence of regime-dependent variances (see, for instance, Kim 1993; Hamilton and Susmel 1994; Kim and Kim 1996; Susmel 2000).

The sizes of α_i and β_i, determine the short- and long-run dynamics of the resulting volatility series, respectively. Large β_i coefficients indicate that shocks to conditional variance take a long time to die out, implying persistent volatility. On the other hand, large α_i coefficients indicate that volatility reacts quite intensively to new information. Consequently, if α_i is large (and significant) and β_i is small, this means the volatility process is characterized by spikes. Finally, parameter ρ denotes the level of the conditional correlation between output growth and inflation.

In addition to the above specification, for comparison purposes, we estimate the restricted version of the model where the effects on mean inflation and output growth are independent of regimes. That is, we estimate a single-regime specification. In this case the model reduces to a VAR GARCH-M specification. This model has been widely used in the related literature (see, for instance, Grier and Perry 2000; Wilson 2006) and it represents a good benchmark to compare our results with previous studies.

The likelihood function at time t (ignoring the constant term and assuming normality) is given by

$$l_t(\theta) = -\frac{1}{2}\ln|H_t| - \frac{1}{2}\varepsilon_t^{'} H_t \varepsilon_t, \tag{9}$$

where θ is the vector of all the parameters to be estimated. The log-likelihood for the whole sample from time 1 to n, $L(\theta)$, is given by

$$L(\theta) = \sum_{t=1}^{n} l_t(\theta) \tag{10}$$

This log-likelihood function is maximized with respect to all parameters simultaneously, employing numerical derivatives of the log-likelihood. To allow for non-normality of $\varepsilon_t \left| \Psi_{t-1} \right.$, robust "sandwich" standard errors (Bollerslev and Wooldridge 1992) are used for the estimated coefficients.

5. Results

Before we proceed to the estimation of our model described by equations (1) to (10), we first test our assumption of nonlinearities based on own-transition effects in equations (1) and (2). For this purpose, we use the Lagrange Multiplier (LM) test statistic developed by Luukkonen et al. (1988) and extended by van Dijk and Franses (1999) who generalize it for the case of multiple regimes.[21] This LM test examines whether an autoregressive model or a multiple regime alternative (under different transition variables) should be used. The results of the test, illustrated in Table 2, offer strong support for the presence of nonlinearities in our model as our specification (i.e., multiple regimes) is preferred to the autoregressive model. This offers evidence of regime switching behavior in both the rates of inflation and output growth for every G7 country in line with our assumption of own-transition effects. In other words, the test supports the use of our restricted form of a four-regime smooth transition model as it identifies the existence of four distinct regimes.

Turning to the model, we begin by estimating the restricted version of our model that does not account for regime switches. This single-regime model is used as a benchmark for comparing findings both with earlier studies and with the extended model that allows for regime-switching. Given that our goal is to test for the four economic hypotheses presented in Section 2 concerning the impact of macroeconomic uncertainty on macroeconomic performance, we focus our attention on the statistical significance and signs of the elements of matrix Ψ. The coefficient estimates are reported in Table 3.[22]

The economic significance of the results can be summarized as follows. In the conditional mean equation for output growth, the sign and values of Ψ_{11} offer weak support to Black (1987) and the precautionary investment hypothesis as real uncertainty enhances output growth for the USA, UK, and Canada – albeit at the 10% level for the latter two. For the remaining countries, the effect is indistinguishable from zero suggesting independence between output uncertainty and growth in line with business cycles models. Worthy of note is that none of the countries exhibit

[21] This test has been conducted in two steps by taking each equation individually.

[22] The low and mainly insignificant estimates for conditional correlations that appear in the bottom row of the table offers support to the use of the CCC specification of our VAR GARCH-M model consisted with the findings of the related literature.

negative effects, largely dismissing the analysis of Bernanke (1983) and Pindyck (1991). The impact of nominal uncertainty on mean output growth, illustrated by Ψ_{12}, is negative for five of the countries with the exceptions of Italy and Canada. In Italy the effect is positive, while in Canada there is no statistically significant effect. These imply the dominance of the negative link put forward by Friedman (1977) and the models of Pindyck (1991), Huizinga (1993), and Blackburn and Pelloni (2004). Only Italy supports the theoretical arguments of Dotsey and Sarte (2000) and Varvarigos (2008). Comparing these results to the literature as presented in Table 1, one can find both similarities and differences. Focusing on the similarities, the findings for the USA agree with those of Grier et al. (2004) and Bredin and Fountas (2005), while Wilson (2006) supports the results for Japan. The findings for the UK and Italy find backing by Bredin and Fountas (2005) and Bredin and Fountas (2009) respectively.

Moving to the conditional mean equation for inflation, output growth uncertainty bears no effect on the rate of inflation as illustrated by the insignificance of Ψ_{21}. The sole exception is France, for which the effect is positive, lending support to Deveraux (1989). Finally, the estimates of Ψ_{22} point to mixed effects of inflation uncertainty on inflation. In Germany, Italy and Canada the effect is positive offering support to the Cukierman-Meltzer hypothesis, while the findings for the UK and France endorse Holland (1995). Moreover, the effect does not seem to take any shape in the USA and Japan. Empirical confirmation for this pair of results is presented by Grier et al. (2004) and Bredin and Fountas (2005) for the USA, Bredin and Fountas (2005, 2009) for Germany, and Bredin and Fountas (2009) for Italy.

In general, as it regards the results of the single-regime specification, about half of the estimated coefficients in Table 3 are not statistically significant. In addition, of those effects that are significant, more than a third are so only at the 10% level. These findings may be an accurate reflection of the presence of weak effects or even of their overall absence. On the other hand, however, the consideration of a single regime may mask potentially different realizations that would be materialized if one would allow for the particular set of conditions that determine them to take shape. This is exactly the issue we tackle next with the consideration of two distinct regimes.

As discussed in the introduction, a number of recent studies have explored the regime-dependent, or conditional, effects of real and nominal uncertainty (Baillie et al. 1996; Henry and Olekalns 2002; Hnatkovska and Loayza 2004; Chang and He 2010). Following their lead, we allow the conditional mean and conditional variance to be affected by the state of the economy, in the sense that we use as transition variables in our regime-determined model the rate of inflation and the rate of output growth; the latter as a proxy for business cycle phases. The use of these transition variables is further reinforced by the nature of our sample, which covers the period 1957-2009 for the G7 countries. During this time, there have been periods of high-inflation (1960s and 1970s) as well as low-inflation (1990s). Similarly, G7 members experienced high growth rates in the 1960s and 1990s and low growth in the 1980s. This variety of experience in the rates of inflation and growth through the years could alter the related uncertainty of the macroeconomic environment, and its subsequent impact on macroeconomic outcomes.[23]

Table 4 reports the results of the model that takes into account these considerations as described in equations (1) to (10). The most important finding is the presence of regime-switching illustrated by the estimated threshold rates of output growth (c_1) and inflation (c_2), which are all statistically significant at the 1% level. This means that the effect of uncertainty related to output growth on mean growth and inflation varies between low-growth and high-growth regimes. Analogously, the impact of inflation uncertainty on mean growth and inflation is different in sign and/or magnitude as inflation moves from the low to the high regime. The location of the endogenously determined threshold points is quite uniform across countries. The switching point for output growth is in the range of 2-4% with Canada being an outlier (-0.13%), while for inflation the threshold rate is in the range of 1.5-4.5% except for Japan (0.29%).[24] As for the speed of the transition process from one regime to the other, the large estimates of γ_1 and γ_2 indicate abrupt switches between regimes making our models

[23] Recall that our specification allows for the two regimes in both the mean equations (1) and (2) and volatility equations (6) and (7). In the mean equations the variables that are regime-dependent include the constant, the lagged values of the dependent variables and the uncertainty related variables. In the volatility equations these include only the constants.

[24] We have plotted both the growth and inflation rate transition functions to see how many observations fall into the upper and lower regimes. Inspection of these plots shows that there is a substantial number of observations in each regime. This allows us to infer that each regime is detecting genuine economic regimes and excludes the possibility of detecting outliers. The plots are available from the authors upon request.

effectively threshold models, rather than smooth transition ones.[25] Once again, the last row of the table shows that conditional correlations are very low. Moreover they are not statistically significant (at 5%) for any of the countries. Finally, the LM test for remaining nonlinearity produces *p*-values in excess of 0.1 implying that there is no further nonlinearity within our model – for further details on this test see van Dijk and Franses (1999).[26]

Turning to the coefficient estimates of interest, the values of Ψ_{11} show that between low- and high-growth regimes there are five countries for which there is a statistically *different* effect of real uncertainty on output growth. With the exceptions of Japan and France, the LR test of coefficient equivalence between regimes suggests this to be the case. The effect of output growth uncertainty appears to be mainly positive in the low-growth regime and switches to zero or remains positive in the high-growth regime. It is only for the UK that the effect becomes negative in the latter. Therefore, in contrast to the findings of the single-regime model presented in Table 3, there is significant variance of the impact once one accounts for asymmetric effects of business-cycle fluctuations. The interpretation of the double-regime findings, as to the positive influence of growth uncertainty on growth in the low-growth regime, is in line with the theory of "creative destruction" put forth by Schumpeter (1939).[27] According to this argument, the volatility of output growth is associated with recessions. During economic downturn, higher research and development spending and/or the destruction of least productive firms take place, thereby boosting growth. As a result, higher long-run growth can occur alongside higher uncertainty. This argument offers an intuitive explanation for our Ψ_{11} estimates as it requires the presence of deep financial markets, active firm turnover, and the ability to conduct counter-cyclical innovation

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[25] In a number of cases the parameter γ_i becomes large and imprecisely estimated, signifying an abrupt change in the regime. In such cases we report the value of γ_i to be 500 as indicative. However, even a γ_i value of 25 (the lowest obtained) also implies a transition with very few observations with transition function values other than zero or one; plots of the transition functions verify this. Furthermore, as noted by van Dijk and Franses (1999) an insignificant estimate of γ_i should not be interpreted as insignificance of the regime switching. What matters is the magnitude of γ_i rather than its significance.

[26] Standardized residual diagnostic tests, available upon request, suggest that our model is well specified. The mean and variance of the standardized residuals are found to have values of zero and one respectively for all the economies. The values of the Ljung-Box and Ljung-Box squared statistics illustrate the absence of serial correlation and conditional heteroskedasticity up to 4th and 12th order in the standardized and squared standardized residuals in both the output growth and inflation equations.

[27] For a modern treatment of this view, see Shleifer (1986), Hall (1991), Caballero and Hammour (1994), and Aghion and Saint-Paul (1998).

expenditures, characteristics associated with developed economies as our country sample.

The influence of inflation uncertainty on output growth (Ψ_{12}) has been found to be mainly negative in the single-regime specification offering support to Friedman (1977). The consideration of two inflation regimes now corroborates this finding by further showing that this negative effect is re-enforced in the high-inflation regime. In particular, the negative effect materializes for six of the G7 countries in the high-inflation regime and for five of them (except France) the size increases in absolute magnitude. But even for Italy, where the effect is positive in the low-inflation regime, it becomes statistically equal to zero in the high-inflation regime highlighting the diminishing impact of inflation uncertainty on growth between regimes.[28] These findings offer support to Change and He (2010) who find a nonlinear effect of inflation uncertainty on growth across inflation regimes for the US economy. We, however, offer further evidence that this also happens in the rest of the G7 nations. In addition, we complement the empirical studies (Bruno and Easterly 1998; Khan and Senhadji 2001) that find a nonlinear harmful effect of inflation on output growth by showing that the growth-inhibiting effect is even larger if one also takes into account the nonlinearity related with the effect of inflation uncertainty.

Looking at the estimated coefficients of Ψ_{21}, we observe that the insignificant relationship between real uncertainty and inflation obtained in Table 3 does not accurately reflect reality for all countries when we control for business cycles effects. We now find that for four countries, as opposed to just France, there are significant effects. From those, the USA and the UK exhibit negative inflation effects in the low-growth regime (the UK also in the high-growth regime), whereas Germany and France display positive effects in the high-growth regime. Therefore, it seems that Deveraux's (1989) prediction as to a positive causal relationship is confirmed for Germany and France only during periods of output expansion. During economic contractions, the direction of the effect goes the other way for the USA and UK supporting the Taylor effect in conjunction with the Cukierman-Meltzer hypothesis.

[28] The fact that the estimates of Ψ_{12} differ between the two regimes for all countries is confirmed by the LR test statistic.

Finally the mixed findings unveiled in Table 3 as to the effect of nominal uncertainty on inflation (Ψ_{22}) are also present in Table 4. Japan is the only country for which the link between nominal uncertainty and inflation is statistically non-existent regardless of the use of one or two regimes. For the rest of the G7 the effects point toward nonlinearities between low- and high-inflation regimes – see the LR test statistic. Specifically, for the UK and France the effect is U-shaped. As these economies move from the low- to the high-inflation regime, inflation uncertainty first increases and then decreases inflation. For the rest of the countries for which significant effects take place (USA, Germany, Italy and Canada), the effect in the low-inflation regime varies between positive, negative, and zero. Most importantly, however, at the high-inflation regime, the effect becomes positive and sizeable. This outcome accords well with Bailie et al. (1996) who find enhancing effects of nominal uncertainty on inflation in high inflation countries.

The above results carry a number of interesting implications for macroeconomic modeling, estimation, and analysis. Most importantly, they offer an explanation for the ambiguous and conflicting findings reported in the literature as to the relationship between macroeconomic uncertainty and macroeconomic performance. For example, the findings of the literature for the USA described in Table 1 show a variety of outcomes for each of the effects. The value of Ψ_{11} reported in the literature as either positive or zero accords well with our finding which suggests a zero effect in the low-growth regime and a positive effect in the high-growth regime. Similarly, the range of values for the negative sign of Ψ_{12} found in studies can be explained by our finding of a negative effect which is increasing in magnitude as moving to the high-inflation regime. The result we obtain as to the negative effect of Ψ_{21} in the low-growth regime and the statistically zero effect in the high-growth regime may also explain these two *average* effects obtained in early studies. Finally, the opposite signs we unveil for the effect described by Ψ_{22} in the low and the high inflation regimes, can explain the findings in the literature as they either offset each other giving rise to a zero effect, or one of the effects dominates. A similar kind of comparison can be carried out along these lines for the rest of the G7 countries. The important thing to note is that our findings reveal significant nonlinearities in the relationships of macroeconomic uncertainty and outcomes that advocate (i) the construction of theoretical macroeconomic models that take this element into account and (ii) their incorporation into empirical analysis to avoid misspecification bias.

6. Conclusion

In this paper, we develop a bivariate Smooth Transition VAR GARCH-M system of inflation and industrial production growth in the G7 countries. Our goal is to investigate the presence of nonlinearities in the effects of real and nominal uncertainty on average inflation and output growth. By making the effects conditional on the rate of inflation and the economy’s position along the business cycle, we find significant regime-switching effects.

We can summarize our findings as follows. First, there is evidence that output growth uncertainty enhances mean output growth with the effect being mainly restricted in the low-growth regime during economic downturns. Second, higher inflation uncertainty leads to lower mean output growth especially during high inflation periods. Third, output growth uncertainty does not appear to have a significant effect on mean inflation with weak evidence of a diminishing negative, or increasing positive, effect as economies expand. Fourth, there are mixed findings for the effect of inflation uncertainty on mean inflation being mostly positive and increasing in size in inflationary periods.

Our results are important as they demonstrate the existence of significant regime-dependent effects of real and nominal uncertainty on the rates of inflation and output growth. Our findings also help to explain the often-mixed results in the literature because we demonstrate that the empirical relationship between macroeconomic uncertainty and macroeconomic performance is quite complex. This implies that studies examining the above links will tend to produce misleading results unless the threshold, or nonlinear, effects are taken into account.

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Table 1
The Effects of Macroeconomic Uncertainty on Inflation and Growth in the Literature

Country	***Study***	***Model***	Ψ_{11}	Ψ_{12}	Ψ_{21}	Ψ_{22}	***Output***	***Inflation***	***Frequency***	***Period***	
USA	Grier and Perry (2000)	GARCH-CCC-M	3.45	**-1.03**	-0.17	-0.03	IPI	PPI	Monthly	Jul-48	Dec-96
	Grier et al. (2004)	VARMA-Asymmetric BEKK-M	**0.08**	**-0.24**	**-0.004**	**-0.02**	IPI	PPI	Monthly	Apr-47	Oct-00
	Bredin and Fountas (2005)	VARMA-Asymmetric BEKK-M	**0.11**	**-0.09**	0.02	-0.01	IPI	PPI	Monthly	Jan-57	May-03
Japan	Bredin and Fountas (2005)	VARMA-Asymmetric BEKK-M	0.19	0.54	**0.32**	1.48	IPI	PPI	Monthly	Jan-57	Apr-03
	Wilson (2006)	EGARCH-CCC-M	-0.02	**-0.42**	**0.02**	**1.61**	GDP	CPI	Quarterly	Q4-57	Q3-02
United Kingdom	Bredin and Fountas (2005)	VARMA-Asymmetric BEKK-M	**0.58**	**-0.53**	**0.40**	0.05	IPI	PPI	Monthly	Jan-57	Apr-03
	Bredin and Fountas (2009)	VARMA-Asymmetric BEKK-M	**-0.46**	**0.75**	**2.68**	**-4.93**	IPI	PPI	Monthly	Jan-62	Dec-03
		Structural VAR-BEKK-M	**-0.75**	**0.30**	**1.55**	-0.60	IPI	PPI	Monthly	Jan-62	Dec-03
Germany	Bredin and Fountas (2005)	VARMA-Asymmetric BEKK-M	**0.54**	0.14	-0.01	**0.30**	IPI	PPI	Monthly	Jan-57	May-03
	Bredin and Fountas (2009)	VARMA-Asymmetric BEKK-M	**-0.97**	**10.54**	-0.04	**0.57**	IPI	PPI	Monthly	Jan-62	Dec-03
		Structural VAR-BEKK-M	**-0.50**	**1.05**	-0.22	0.06	IPI	PPI	Monthly	Jan-62	Dec-03
France	Bredin and Fountas (2005)	VARMA-Asymmetric BEKK-M	**-0.05**	**0.74**	**0.02**	**0.82**	IPI	PPI	Monthly	Jan-57	May-01
	Bredin and Fountas (2009)	VARMA-Asymmetric BEKK-M	-0.01	2.99	0.01	0.10	IPI	CPI	Monthly	Jan-62	Dec-03
		Structural VAR-BEKK-M	-0.36	0.09	0.12	0.09	IPI	CPI	Monthly	Jan-62	Dec-03
Italy	Bredin and Fountas (2005)	VARMA-Asymmetric BEKK-M	**-0.18**	**3.93**	**-0.01**	**0.06**	IPI	PPI	Monthly	Jan-57	Apr-03
	Bredin and Fountas (2009)	VARMA-Asymmetric BEKK-M	0.30	**4.02**	0.03	**0.18**	IPI	CPI	Monthly	Jan-62	Dec-03
		Structural VAR-BEKK-M	**0.08**	0.60	**1.74**	**0.10**	IPI	CPI	Monthly	Jan-62	Dec-03
Canada	Bredin and Fountas (2005)	VARMA-Asymmetric BEKK-M	**0.02**	**0.03**	**-0.03**	0.00	IPI	PPI	Monthly	Jan-57	Apr-03

Notes: Studies in the table reflect those that use simultaneous bivariate GARCH-M models in inflation and output growth. The focus is in studies that use the G7 set of countries, or a subset, for comparison purposes (for example, Bredin and Fountas (2009) examine 14 European Union countries). Coefficients in bold type represent significance at least at the 5% level. IPI: Industrial Production Index; PPI: Producer Price Index; GDP: Gross Domestic Product; CPI: Consumer Price Index.

Table 2
LM-type test for STAR nonlinearity

Transition Variables	*USA*		*Japan*		*United Kingdom*		*Germany*		*France*		*Italy*		*Canada*	
	Growth	Inflation	Growth	Inflation	Growth	Inflation	Growth	Inflation	Growth	Inflation	Growth	Inflation	Growth	Inflation
$\boldsymbol{y^{*}_{t-1}}$, $\boldsymbol{\pi^{*}_{t-1}}$	**47.87**	**109.12**	**182.77**	**27.14**	**49.61**	**22.17**	**68.58**	**154.56**	**86.01**	**87.18**	**49.27**	**81.40**	**26.19**	**33.27**
	(0.000)	**(0.000)**	**(0.000)**	**(0.000)**	**(0.000)**	**(0.038)**	**(0.000)**	**(0.000)**	**(0.000)**	**(0.000)**	**(0.000)**	**(0.000)**	**(0.014)**	**(0.000)**

Notes: p-values in parenthesis. Coefficients in bold type represent significance at least at the 5% level

Table 3
The Values of the Ψ Matrix for the G7 (Single Regime)

	USA	*Japan*	*United Kingdom*	*Germany*	*France*	*Italy*	*Canada*
Ψ_{11}	**0.349**	0.414	**0.116**	-0.083	-0.015	0.078	**0.306**
	(0.044)	(0.439)	**(0.080)**	(0.360)	(0.455)	(0.354)	**(0.080)**
Ψ_{12}	**-0.364**	**-1.34**	**-0.080**	**-0.283**	**-0.921**	**1.772**	0.021
	(0.018)	**(0.020)**	**(0.092)**	**(0.083)**	**(0.000)**	**(0.019)**	(0.473)
Ψ_{21}	-0.043	0.001	0.120	-0.004	**0.154**	0.033	0.020
	(0.213)	(0.498)	(0.158)	(0.449)	**(0.000)**	(0.139)	(0.232)
Ψ_{22}	0.051	-0.076	**-0.652**	**0.503**	**-0.366**	**0.508**	**0.435**
	(0.327)	(0.443)	**(0.076)**	**(0.005)**	**(0.003)**	**(0.000)**	**(0.001)**
Correlation	-0.009	0.020	**0.110**	**0.061**	-0.011	0.014	0.015
	(0.400)	(0.371)	**(0.033)**	**(0.050)**	(0.418)	(0.390)	(0.363)

Notes: p-values in parenthesis. Coefficients in bold type represent significance at least at the 10% level.

Table 4. The Values of the Ψ Matrix for the G7 (Double Regime)

	USA	*Japan*	*United Kingdom*	*Germany*	*France*	*Italy*	*Canada*
$\boldsymbol{\Psi_{11,l}}$	0.211	0.173	**0.334**	**0.394**	**0.324**	**0.248**	**0.923**
	(0.203)	(0.287)	**(0.025)**	**(0.091)**	**(0.097)**	**(0.061)**	**(0.000)**
$\boldsymbol{\Psi_{11,h}}$	**0.991**	0.023	**-0.612**	-0.151	0.282	0.106	**0.195**
	(0.000)	(0.481)	**(0.001)**	(0.371)	(0.112)	(0.307)	**(0.034)**
$\boldsymbol{\Psi_{12,l}}$	**-0.215**	**-0.658**	-0.227	-0.017	**-0.829**	**1.958**	**-0.275**
	(0.013)	**(0.094)**	(0.170)	(0.469)	**(0.000)**	**(0.000)**	**(0.040)**
$\boldsymbol{\Psi_{12,h}}$	**-0.301**	**-0.960**	**-1.182**	**-1.062**	**-0.472**	0.514	**-0.404**
	(0.000)	**(0.006)**	**(0.000)**	**(0.013)**	**(0.000)**	(0.309)	**(0.004)**
$\boldsymbol{\Psi_{21,l}}$	**-0.073**	-0.001	**-0.195**	0.008	0.023	0.027	-0.024
	(0.088)	(0.497)	**(0.000)**	(0.421)	(0.309)	(0.135)	(0.277)
$\boldsymbol{\Psi_{21,h}}$	-0.011	-0.022	**-0.179**	**0.073**	**0.095**	0.013	0.024
	(0.404)	(0.377)	**(0.000)**	**(0.044)**	**(0.038)**	(0.304)	(0.286)
$\boldsymbol{\Psi_{22,l}}$	**-0.227**	0.456	**0.357**	**0.408**	**1.345**	**0.193**	0.149
	(0.086)	(0.228)	**(0.003)**	**(0.000)**	**(0.000)**	**(0.088)**	(0.225)
$\boldsymbol{\Psi_{22,h}}$	**0.156**	-0.266	**-1.207**	**1.136**	**-0.458**	**0.707**	**0.450**
	(0.067)	(0.163)	**(0.000)**	**(0.000)**	**(0.014)**	**(0.005)**	**(0.012)**
Log Likelihood	-4219.03	-4023.25	-4407.85	-4427.47	-4833.47	-4385.13	-4340.38
LR test of $\boldsymbol{\Psi_{11,l}} = \boldsymbol{\Psi_{11,h}}$	**13.20**	0.004	**23.10**	**14.96**	2.240	**52.68**	**8.160**
	(0.000)	(0.841)	**(0.000)**	**(0.000)**	(0.134)	**(0.000)**	**(0.000)**
LR test of $\boldsymbol{\Psi_{12,l}} = \boldsymbol{\Psi_{12,h}}$	**25.40**	**22.90**	**17.08**	**10.32**	**5.680**	**46.24**	**126.82**
	(0.000)	**(0.000)**	**(0.000)**	**(0.001)**	**(0.018)**	**(0.000)**	**(0.000)**
LR test of $\boldsymbol{\Psi_{21,l}} = \boldsymbol{\Psi_{21,h}}$	**23.24**	2.280	**2.800**	**5.640**	**2.790**	2.24	0.160
	(0.000)	(0.131)	**(0.094)**	**(0.018)**	**(0.095)**	(0.134)	(0.689)
LR test of $\boldsymbol{\Psi_{22,l}} = \boldsymbol{\Psi_{22,h}}$	**29.76**	1.890	**37.21**	**51.60**	**13.29**	**53.24**	**121.76**
	(0.000)	(0.169)	**(0.000)**	**(0.000)**	**(0.000)**	**(0.000)**	**(0.000)**
p-values of LM test for remaining nonlinearity	0.459	0.317	0.332	0.159	0.223	0.371	0.628
$\boldsymbol{c_1}$ *– growth rate threshold*	***2.821***	***3.855***	***3.519***	***3.104***	***2.193***	***2.695***	***-0.130***
	(0.000)	***(0.000)***	***(0.000)***	***(0.000)***	***(0.000)***	***(0.000)***	***(0.000)***
$\boldsymbol{c_2}$ *– inflation rate threshold*	***2.961***	***0.290***	***4.365***	***1.539***	***2.149***	***4.569***	***2.476***
	(0.000)	***(0.000)***	***(0.000)***	***(0.000)***	***(0.000)***	***(0.000)***	***(0.000)***
$\boldsymbol{\gamma_1}$ *– growth transition*	497	453.82	201.71	500	493.46	500	148.31
$\boldsymbol{\gamma_2}$ *– inflation transition*	74.85	500	57.71	36.62	25.64	500	500
Correlation	0.003	0.004	**0.067**	0.049	-0.001	0.024	-0.015
	(0.468)	(0.470)	**(0.063)**	(0.115)	(0.488)	(0.268)	(0.374)

Notes: p-values in parenthesis. Coefficients in bold type represent significance at least at the 10% level.