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Business Cycle Moderation - Good Policies or Good Luck: Evidence and Explanations for the Euro Area*

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Abstract

Economic fluctuations in most of the industrialised world have for over the past 30 years been characterised by declining volatility. This decline has also been a trait witnessed for output fluctuations in the Euro Area. This paper has two objectives. The first is to provide a comprehensive characterisation of the decline in volatility using a large number of Euro area economic time series and a variety of methods designed to describe the time-varying time series processes. The second objective is to provide new evidence on the quantitative importance of various explanations for this 'great moderation'. This paper focuses on the central elements in the literature contending why real output growth has stabilised. Such factors include shifts in the structure of the economy, improved policies, and a 'good luck' factor. Further, this paper goes on to investigate whether cross-country linkages in growth have shifted, perhaps in a way that can help rationalise the stabilisation in output. Taken together, the moderation in volatility is attributable to a combination of improved policy (around 5 - 30 percent) and identifiable forms of good luck that manifest themselves as smaller reduced-form forecast errors (40 percent).

JEL *Classification*: E32, E60
Keywords: Output Volatility, Monetary Policy, International shocks

1 Introduction

The history of business cycles can be conveniently summarised by measuring the volatility of economic growth. Using this measure, the past 30 years has witnessed a considerable decline in the volatility of economic activity in most industrialised economies. The reduction in volatility has been widespread across sectors within the G7. It was Kim and Nelson (1999) who coined the phrase the 'great moderation' to describe the increasing stability seen in business cycle fluctuations over the past three decades. Much has been written about the possible causes of this great moderation.

Although the fact that declining business cycle volatility is common wisdom, there is much less agreement about the causes of improved macroeconomic stability, especially with regards to improved output stability and secondly, whether it will endure. Much of the early literature has focused upon the US experience – see

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Kim and Nelson (1999), McConnell and Perez Quiros (2000) and Stock and Watson (2002a, 2003). These studies have contend that economies have become more self-stabilising as a result of the shift in economic activity from the secondary to the tertiary sector, better inventories management by firms and integration of financial markets. Other economists, such as Cogley and Sargent (2005) and Taylor (1998), have put forward the claim that institutional change, such as central bank independence along with more transparent monetary policy and inflation targeting has led to improved economic stability. Consequently, Blanchard and Simon (2001) emphasise the role of inflation volatility in the decline of output growth volatility. In distinction to such theories, Stock and Watson (2002a, 2003) put the stabilisation down to unadorned 'good luck', which allows them to draw the conclusion that the quiescence of the past fifteen years could well be a hiatus before a return to more turbulent economic times. In support, Martin and Rowthorn (2005) contend that the record of recent years is an exception and unlikely to continue[1]. Whilst improved credit markets have allowed households to smooth their spending, automatic stabilisers have also meant that incomes have varied less than production.

This paper provides a comprehensive characterisation of the decline in volatility using a large number of Euro area economic time series and a variety of methods designed to describe the time-varying time series processes. Apart from the US economy, there has been little work undertaken on other industrialised economies examining why output has stabilised over the past two decades. Hence, the primary objective of this paper is to provide new evidence on the quantitative importance of various explanations for the moderation witnessed in the Euro area cycle. Such an analysis for the Euro area has added significance, given the current efforts to understand the workings of the Euro are economy as a whole, in terms of the impulses and propagation that drive the cycle – real, monetary and international – in the design of Euro-wide policies[2]. Understanding the causes of the moderation of business cycles remains a crucial issue (Diebold and Rudebusch, 2001). Increasingly instability in output increases risk and premia associated with risk in the economy. Increases in risk are likely to reduce the level of equilibrium output, possibly leading to both higher saving and a lower capital stock, which may in turn lead to greater capital outflows in an open economy. Policies that reduce anticipated and unanticipated volatility will therefore raise output and welfare

[1] In contrast, Bernanke (2004) paints a more optimistic future.

[2] The introduction of a common currency has increased the interest and need for business cycle analysis at the Euro area level. Such analysis acts as a reference for economic agents due its influence on monetary policy decisions.

in the longer run.

As in Stock and Watson (2002a, 2003), the investigation here falls into five main categories. The first category will examine the evidence for structural change, helping to provide an answer to the question that underlies the bulk of the literature in this topic area; has there been a structural break in post-war real output growth towards stabilisation?[3] In the US case, Kim and Nelson (1999) and McConnell and Quiros (2000) documented a structural break in the volatility of output growth, finding a dramatic reduction in the output volatility in the most recent two decades relative to the previous three[4]. This is investigated using a stochastic volatility model, which allows for the conditional mean and the conditional variance to break (or not) at potentially different dates. The second category, first contended by Moore and Zarnowtiz (1986) and later by McConnell and Quiros (2000), will focus upon the changes in the structure of the economy, which include the shift in output from goods to services. The third category examines the impulse and propagation mechanism for the Euro area to investigate signs of structural shifts in either the impulses or propagation (which will act as a proxy to changes in the structure of the economy). The fourth category will examine whether improved monetary policy has led to a decline in output volatility, as first suggested by Taylor (1998) for the US economy (also see Cogley and Sargent, 2005). This category also extends to a shock based analysis of a variety of different variables and whether such disturbances have become more benign i.e. a 'good luck' category. The fifth, and final, category spotlights external business cycle comovements, as in Doyle and Faust (2005). The comovements analysis will examine the role played by shocks in the business cycles of the four main trading partners of the Euro area, which include Japan, the US and the US. This allows one to gauge their role and contribution to greater stability with regards to the Euro area cycle.

2 Economy-wide Reductions in Volatility

This section documents the widespread reduction in volatility and provides nonparametric estimates of this reduction for major economic time series.

[3] See Sensier and van Dijk (2004) and Stock and Watson (2002).

[4] McConnell and Quiros (2000) suggest that the decline in US output volatiltiy can be traced to a break in the volatility of durable goods production, whose timing corresponds to a reduction in the proportion of durables accounted for by inventories.

2.1 Data Series

The lack of a long time series data set for the Euro area, which decomposes GDP into various major economic time series, as with the NIPA[5] dataset for the US economy, the data in this paper represents a wide range of macroeconomic activity arrived at from a variety of different sources to help ensure a data set long enough for meaningful economic analysis. Most data series used in this paper are available from Datastream. The exceptions are the crude oil price and the raw materials index, both of which are gathered from the 2005 International Financial Statistics (IFS) series from the World Bank. The second is the average hours worked, short-term interest rate, long-term interest rate, total consumption and investment, all of which come from Fagan *et al.* (2001) Euro-wide dataset. Finally, the composite leading indicator is an OECD measure (2005). Seasonally adjusted series were used when available. All of the analysis uses quarterly observations, which are transformed to eliminate trends and nonstationarity.

2.2 Volatility Measures

Table 2 reports the sample standard deviation of 27 leading macroeconomic time series. Each time period standard deviation is presented relative to the full-same standard deviation, so a value less than one indicates a period of relatively low volatility. The key demand and production variables illustrate a decline in volatility, with standard deviation percentages all less than one. All measures of inflation also reflect a decline in volatility. The external sectors also show a decline. With the Euro area's main trading partners being Japan, UK and the US, the exports estimate in Table 1 support the results found in other studies of a decline in consumption and production volatility in the G7 economies (see Mills and Wang 2002, 2003 and Stock and Watson 2002a, 2003). The results in Table 1 for the Euro area economy as a whole, differ from the results by Blanchard and Simon (2001), who found the relative standard deviation of industrial production to be lower in the eighties than was the case for the nineties.

Examining the monetary sector, one finds that the interest rate result in Table 1 is in partial similarity with the result found for the US by Stock and Watson (2002a, 2003). The Euro area experienced a decrease in the variance in the interest rates both at the long and short end, however this decrease in volatility is

[5] National Income Public Accounts

slightly more marked for the long-term interest rate, a statistical observation that differs from that found by Stock and Watson (2002a) for the US economy. The decline in volatility is also reflected in other series. Table 2 shows the relative standard deviation of different sectors in the total labour market. Employment volatility has fallen in the highly volatile industrial and construction sectors. This result coincides with the results shown in Table 1.

Table 1: Standard Deviations of Annual Growth Rates Macroeconomic Time Series

Series	Standard Deviation 1980-2005	Standard Deviations, relative to 1980-2005		
		1980-1987	*1988-1994*	*1995-2005*
GDP	0.039	1.24	1.11	0.69
Consumption	0.039	1.24	1.04	0.62
Private Consumption	0.724	1.62	0.44	0.61
Gov't Consumption	0.506	1.30	0.94	0.77
Capital Consumption	1.918	.	0.21	1.15
Investment	0.026	0.81	1.31	0.83
GFCF Investment	1.502	1.25	0.98	0.80
Residential	1.384	.	0.94	1.02
Non-Residential	1.593	.	1.36	0.84
Exports	1.483	1.15	0.94	0.92
Imports	1.460	1.05	1.17	0.83
Production				
Goods (total)	7.742	1.16	1.08	0.81
Non-Durables	0.854	0.75	1.16	0.94
Capital Goods	1.532	0.76	1.28	0.84
Construction	3.185	1.18	1.08	0.79
Producer Price Index	0.793	1.47	0.43	0.90
Inflation (CPI)	0.779	1.51	0.37	0.79
GDP Deflator	0.628	1.47	0.41	0.85
Employment	0.284	0.92	1.35	0.76
Unit Labour Cost	0.749	0.97	1.28	0.80
Average Hours Worked	0.108	0.61	1.32	0.93
Composite Leading Indicator (OECD)	1.405	1.23	1.01	0.79
Money M1	0.964	0.52	0.92	1.27
Money M3	0.687	0.79	1.13	1.05
Money Stock	1.601	0.43	0.58	1.41
Money Demand ($m_t - p_t$)	2.386	1.03	0.98	0.99
Short Interest Rate	3.488	1.09	0.59	1.17
Long Interest Rate	3.009	1.19	0.38	1.16
Great Ratios				
Consumption : GDP ratio	0.008	1.07	0.83	1.07
Investment : GDP ratio	0.036	1.23	0.97	0.73

B. Average Phase Durations of the Euro Area Cycle[+]

	1980:1990				*1991:2005*			
	P-P	*T-T*	$\overline{C}$	$\overline{E}$	*P-P*	*T-T*	$\overline{C}$	$\overline{E}$
GDP	21.0	27.0	12	13.5	30.0	33.7	17.5	14.7

Notes: *The series are annual growth rates. Inflation is the four-quarter change in the annual inflation rate. The capital consumption series starts from 1991, production non-durables begin at 1985 and finally, non-residential and residential data series starts from 1991.*

\+ *$\overline{C}$ and $\overline{E}$ respectively denote average lengths of expansions and contractions (in months), P-P denotes peak-to-peak cycle and T-T denotes a trough-to-trough cycle. All results were calculated using the Bry-Boschan (1971) algorithm.*

Investigating the average phase durations of the Euro area cycle if one were to split the sample in accord with the GDP break date in Table 2, the sample period after the break data, 1991-2005, is characterised by longer cycles according to the Bry-Boschan (1971) algorithm i.e. less cyclical behaviour.

Table 2: Employment Volatility

	Agriculture	Industry	Construction	Self-employed	Employees	Total market
1981-1987	0.85	1.15	1.9	0.94	0.83	0.91
1988-1994	1.05	2.59	1.83	0.99	1.63	1.44
1995-2005	1.03	0.97	1.4	0.78	0.86	0.7

Notes: *Results represent percentages.*

All measures of money also show a slight rise in the level of volatility. However, as touched upon by Kim, Nelson and Piger (2001), Stock and Watson (2002a) and Basistha and Startz (2004) the situation regarding different monetary indicators is somewhat complex.

2.2.1 Estimates of time-varying standard deviations

Figure 1 attempts to provide a graphical proof of the decline in volatility of real output for the Euro area.

Figure 1 – Euro Area Real GDP Time Series

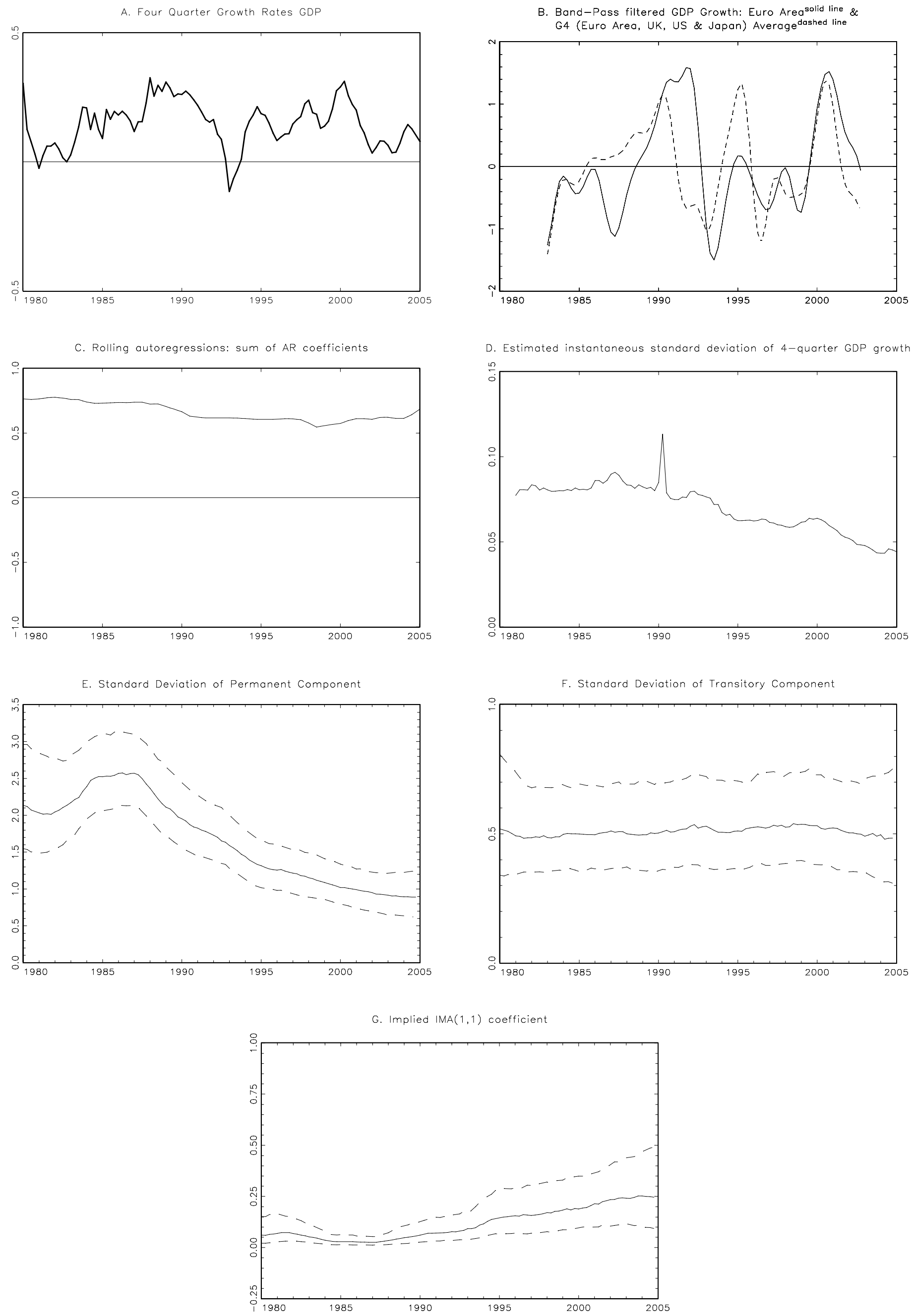


Figure 1.D illustrates the time-varying estimates, where the light line is a raw estimate of the volatility of the series. The smoothened line shows the instantaneous time-varying standard deviation of the series, based on an AR(4) with time varying parameters and stochastic volatility. This model and associated non-Gaussian smoother are conceptually similar to the multivariate approach in Cogley and Sargent (2002) (Stock and Watson, 2002a). Specifically, y_t follows the time-varying AR process, $y_t = \Sigma_{j=1}^{p} \alpha_{jt} y_{t-j} + \sigma_t \epsilon_t$, where $\alpha_{jt} = \alpha_{jt-1} + c_j \eta_{jt}$ and $\ln \sigma_t^2 = \ln \sigma_{t-1}^2 + \varsigma_t$. The standard assumptions apply $\epsilon_t \sim i.i.d.(0,1)$ and $\eta_{1t}, ..., \eta_{pt} \sim i.i.d.(0,1)$. The model allows for large jumps in σ^2, thereby capturing a possible break in the variance, by using a mixture of normal models for the error term ς_t, where the error term ς_t is distributed $N(0, \tau_1^2)$ with probability q and $N(0, \tau_2^2)$ with probability $1 - q$. The model is estimated with $p = 4$. For these calculations the standard calibration parameters are used with $\tau_1 = 0.1$, $\tau_2 = 0.2$ and $q = 0.95$[6]. The estimated instantaneous autocovariance function of y_t are computed using $\sigma_{t|T}^2$ and $\alpha_{jt|T}$. The conditional means of σ^2 and α_{jt} are given by $y_1, ..., y_T$.

Figure 1.D illustrates output following a more stable output path subsequent to the late eighties / early nineties recession in the Euro area. Graphical evidence on the decline in volatility for the principal economic series for the Euro area are also provided in Appendix A[7]. There are however a few notable exceptions to the declining volatility witnessed in the main indicators of the economy. The short and long-term interest rate has seen a slight rise in the level of volatility. Its a point worthy of note, that volatility in short-term interest rates began to rise from 1985 onwards. This period was characterised by stronger commitments from central banks across the Euro area in keeping their currencies within the Exchange Rate Mechanism (ERM), than was the case at the launch of the ERM in 1979. Finally, Figure 1.C graphically measures the change in the persistence of a shock to GDP growth, specifically, the sum of the AR coefficients. Figure 1.C illustrates a decline in the level of persistence to a shock.

The analysis is taken one-step further by decomposing output into its permanent and transitory components, as first suggested in the seminal article by Beveridge and Nelson (1981). Output is decomposed as $y_t = \tau_t + \eta_t$, where $\tau_t = \tau_{t-1} + \epsilon_t$ is a stochastic trend component[8]. The logarithms of the variances

[6] Trying different calilbration parameters has little overall bearing on Figure 1.D. The calibration parameters are based on variance estimates.

[7] No weights, by their average nominal GDP share, are given to any of the figures in Appendix A due to data limitations, which transpires into the case that not all the data has come from the same source.

[8] The approach follows that of Stock and Watson (2003) in using the unobserved components - stochastic volatility model.

are $\eta_t = \sigma_{n,t}\gamma_{n,t}$ and $\epsilon_t = \sigma_{s,t}\gamma_{s,t}$, where $\gamma_t = (\gamma_{n,t}, \gamma_{\epsilon,t}) \sim i.i.d.(0, I_2)$. The logarithms of the variances evolve as independent random walks, $\ln \sigma^2_{n,t} = \ln \sigma^2_{n,t-1} + \psi_{n,t}$ and $\ln \sigma^2_{\epsilon,t} = \ln \sigma^2_{\epsilon,t-1} + \psi_{\epsilon,t}$, where $\psi_t = (\psi_{n,t}, \psi_{\epsilon,t}) \sim i.i.d.(0, \gamma I_2)$ and γ_t and ψ_t are $i.i.d.$ with γ as a scaler which controls the smoothness of the stochastic volatility process. The results are estimated with a vague prior of $\gamma = 0.2$[9]. Figure 1.E shows a substantial decline in volatility from the early eighties, which were characterised by high variations in the permanent component, τ_t, of output. The most pronounced decline occurs from the late eighties, early nineties onwards. This stands in contrast to the transitory component, which shows little, if no, decline in volatility. The results have changed little over time, with the estimates hovering around half a percent. As a result, Figure 1.G, which represents the coefficient in the implied IMA(1,1), tracks the inverse of the smoothed estimates of $\sigma_{\epsilon,t}$. The coefficients rise from close to zero to just under 0.25. The results from the stochastic volatility model would seem to imply a trend break, τ, which has led growth towards stabilisation.

3 Dating the Great Moderation

The evidence presented previously has strongly indicated widespread volatility decline in the major economic time series. In this section, the analysis goes on to investigate whether this decline is associated with a single distinct break in the volatility of these series and if so, when this might have occurred.

In contrast to using the traditional markov-switching model to check for structural breaks, as in Mills and Wang (2003) and Kim and Nelson (1999), Table 3 examines the univariate evidence on whether the change in variance is associated with changes in the conditional mean of the univariate time series process or changes in the conditional variance. Variance changes could arise from changes in the AR coefficients, θ_t, which would represent changes in the conditional mean (given its previous values) or changes in the variance, ϵ_t. The change in the variance of a series can be associated with changes in its spectral shape, changes in the level of its spectrum[10], or both (Stock and Watson, 2002a). The results in Table 2 are estimated from the following AR model

$$y_t = \alpha_t + \theta_t(L)y_{t-1} + \epsilon_t$$

[9] Changing the value of the prior, γ, has little overall effect on the shape of Figures 1.E, 1.F and 1.G

[10] See Cogley and Sargent (2005), Seniser and Dijk (2001), Blanchard and Simon (2001) and Kim and Nelson (1999).

where

$$\alpha_t + \theta_t(L) = \begin{matrix} \alpha_1 + \theta_1(L),\ t \le \kappa \\ \alpha_2 + \theta_2(L),\ t > \kappa \end{matrix} \qquad var(\epsilon_t) = \begin{matrix} \sigma_1^2, t \le \tau \\ \sigma_2^2, t > \tau \end{matrix}$$

where $\theta_t(L)$ is a lag polynomial, where κ and τ are break dates in the conditional mean and variance. The heteroskedasticity-robust Quandt (1960) likelihood ratio (QLR) statistic is used to test for a break in the conditional mean. As mentioned by Stock and Watson (1998, 2002a), the QLR test statistic has power over other forms of time variation such as drifting parameters[11]. The conditional variance break is calculated by the QLR statistic, which looks for a break in the mean of the absolute value of the residuals from the estimated AR model above, where the AR allows for a break in the AR parameters at the estimated break date $\widehat{\kappa}$. The test for the break in the conditional variance is computed with the errors recovered from the above AR equation, which are denoted $\epsilon_t(\kappa)$. The AR coefficients break at date κ, with $\widehat{\epsilon_t}(\kappa)$ denoted as the OLS residuals estimated with a break in the AR coefficients at date κ. Under the null hypothesis of no break in the variance, $E|\epsilon_t(\kappa)|$ is constant. By contrast, under the alternative hypothesis that there is a break date τ, $E\,|\epsilon_t(\kappa)| = \sigma_1 + \lambda 1(t \ge \tau)$, where σ_1 is the first-period standard deviation and λ is the difference between the standard deviations before and after the break. Therefore, the break test is undertaken by computing the QLR statistic in the regression of $|\widehat{\epsilon_t}(\widehat{\kappa})|$ against a binary variable $1(t \ge \tau)$ using homoskedastic standard errors, where $\widehat{\kappa}$ parameter is estimated using OLS. Table 3 also illustrates a trend-augmented version, in which $|\widehat{\epsilon_t}(\widehat{\kappa})|$ is regressed against a constant, $1(t \ge \tau)$ and a time trend t, as well as the $p-$value for the test that the coefficient on t is zero in the regression in which $\tau = \widehat{\tau}$.

The confidence intervals for the conditional variance break data are also computed with OLS from the regression $|\widehat{\epsilon_t}(\widehat{\kappa})|$ against a constant and $1(t \ge \tau)$. Consequently, as noted by Stock and Watson (2002a), if there is a break in the variance of the error term in this regression, it will differ before and after the break. The confidence interval for the break data is then obtained by inverting the test of the break data, which is based upon scaling the distribution differently on either side of the break by the appropriate estimated variance. For that reason, the asymmetric confidence intervals estimated, express greater uncertainty about the break data in the low than the high volatility period.

The estimates from the stochastic volatility model have added significance for the Euro area. An often

[11] For a discussion on the estimated break dates and confidence intervals, see Bai *et al.* (1998)

heard criticism of the empirical research on the Euro area is that the final conclusions and policy implications are based on results obtained using historical pre-Euro area data (Mihov, 2001). The finding of a break date around the time of the Euro's introduction would validate this concern[12].

The model is estimated as an AR(4) to ensure sufficient dynamics. However, the results change little with the model estimated as an AR(2).

Table 3: Estimates and Tests for Changes in the Autoregressive Parameters

$$y_t = \alpha + \theta(L) y_{t-1} + \varepsilon_t$$

	Conditional Mean			**Condiational Variance: Break only**			**Conditional Variance: Trend and break**		
	P-value	Break date	67% confidence interval	*P*-value	Break date	67% confidence interval	*P*-value trend	*P*-value break	Break date
GDP	0.00	1990:1	1989:3 - 1990:3	0.97	.	. - .	0.36	0.39	.
Consumption	0.00	1990:1	1989:3 - 1990:3	0.18	.	. - .	0.95	0.81	.
Private Consumption	0.01	1985:4	1985:2 - 1986:2	0.80	.	. - .	0.23	0.18	.
Gov't Consumption	0.00	1984:1	1983:3 - 1984:3	0.01	1988:4	1983:1 - 1989:3	0.20	0.00	1988:4
Capital Consumption		1995:1	1994:3 - 1995:3	0.00	1994:4	1994:3 - 1995:2	0.50	0.00	1994:4
Investment	0.00	1985:1	1984:3 - 1985:3	0.11	.	. - .	0.08	0.00	1993:1
GFCF Investment	0.00	1985:1	1984:3 - 1985:3	0.12	.	. - .	0.57	0.13	.
Residential	0.00	1998:4	1998:2 - 1999:2	0.09	.	. - .	0.08	0.76	.
Non-Residential	0.00	1996:1	1995:3 - 1996:3	1.00	.	. - .	0.86	0.98	.
Exports	0.00	1998:3	1998:1 - 1999:1	0.02	1991:1	1981:2 - 1991:3	0.99	0.85	.
Imports	0.00	1992:2	1991:4 - 1992:4	0.03	1987:2	1981:1 - 1988:1	0.99	0.70	.
Production									
Goods (total)	0.00	1984:2	1983:4 - 1984:4	0.00	1992:1	1986:4 - 1992:4	0.30	0.00	1992:1
Non-Durables	0.00	1999:2	1998:4 - 1999:4	0.61	.	. - .	0.92	0.91	.
Capital Goods	0.00	1999:2	1998:4 - 1999:4	0.77	.	. - .	0.20	0.25	.
Construction	0.00	1985:1	1984:3 - 1985:3	0.04	1986:1	1981:2 - 1986:2	0.01	0.01	1997:1
Producer Price Index	0.00	1998:4	1998:2 - 1999:2	0.04	1987:1	1985:4 - 1988:4	0.04	0.31	.
Inflation (CPI)	0.00	1988:1	1987:3 - 1988:3	0.01	1991:3	1984:1 - 1992:1	0.02	0.94	.
GDP Deflator	0.00	1998:2	1997:4 - 1998:4	0.54	.	. - .	0.26	0.14	.
Employment	0.00	1993:2	1992:4 - 1993:4	0.47	.	. - .	0.15	0.06	.
Unit Labour Cost	0.00	1993:2	1992:4 - 1993:4	0.07	.	. - .	0.06	0.25	.
Average Hours Worked	0.00	1998:4	1998:2 - 1999:2	0.19	.	. - .	0.00	0.00	1994:2
Composite Leading Indicator (OECD)	0.00	1992:1	1991:3 - 1992:3	0.16	.	. - .	0.86	0.30	.
Money Stock	0.00	1993:4	1993:2 - 1994:2	0.01	1990:1	1981:1 - 1990:2	0.98	0.53	.
Short Interest Rate	0.00	1992:3	1992:1 - 1993:1	0.11	.	. - .	0.07	0.83	.
Long Interest Rate	0.00	1995:1	1994:3 - 1995:3	0.42	.	. - .	0.28	0.60	

Notes: *The p-test results are based on the QLR test for changes in the coefficients of an AR(4). The second column is the OLS estimate of the break date. The final column shows the 67% confidence interval for the break date. The 'Conditional Mean Coefficients' are represented by the parameters α and θ. The 'Conditional Variance' corresponds to* ε_t*, either with or without a time trend in the QLR regression*

Table 3 presents the results of the QLR statistic, which tests the null of no-break. Rejection of the null implies time variation, which may possibly not be of the single break form. The break date for real GDP is estimated to be 1990:1. The 67% confidence interval for the break date is accurate, 1989:3 - 1990:3. This break date coincides relatively closely with the reductions seen in the permanent component of output in Figure 1.E. The finding of a break only in the conditional mean is perhaps not surprising, given the results in Figures 1.E and 1.F. The break date of 1990:1 also matches with the start of a progressive fall in the time-varying standard deviations in Figure 1.D. A break date appears to inflict all variables in Table 3. Further, it is perhaps not surprising to learn that the break period for GDP, 1989:3-1990:3, is characterised

[12] Mihov (2001), who looked at the Lucas critique in a constructive manner, concludes that the overall closeness of his out-of-sample forecasts to the actual data speaks in favour of taking Euro area research based on historical data seriously. He finds no abrupt change from the introduction of the Euro in many Euro area economies.

by structural shifts in the Euro area, the main suspects being German reunification and the collapse of the ERM regime. This result differs from Artis *et al.* (2004), who found a break point in the mid-eighties[13].

The measures of consumption and investment components seem to break in the mid-eighties. Total consumption breaks at exactly the same point as GDP. This is perhaps not so surprising when one looks at Figure 1.D and the consumption figure in Appendix A. The magnitude of troughs and peaks coincide closely to one another. The results for both output and consumption suggest that the 'break model' is appropriate i.e. the decline in volatility has not happened through a discrete reduction in the variance. The results for the other series show widespread instability, especially in the conditional mean. A third of the series reject the null hypothesis of a constant variance. The broad spectrum of results suggest break points in the conditional means are heavily concentrated around the late eighties and early nineties. An observation which is also made for the conditional variance breaks.

The results reported in the final columns of Table 3 provide a further glimpse into the 'trend vs. break' discussion[14]. The last three columns of Table 2 are calculated using the QLR test based regression, $|\epsilon_t| = \phi_0 + \phi_1 t + \phi_2 d_t(\tau) + \eta_t$, where $d_t(\tau)$ is a binary variable that equals one if $t \succcurlyeq \tau$ and equals zero otherwise with η_t is an error term. The results assert that the hypothesis of no break (i.e. the possibility of a time trend in the standard deviation) cannot be rejected at the 95% level for real output. The coefficient on the time trend is not statistically significant different from zero. Hence, the decline in the volatility of GDP growth is perhaps better characterised as a 'break model'. This result stands opposed to the view put forward by Blanchard and Simon (2001), who argued that volatility reduction was better viewed as part of a longer term trend decline, in which the high volatility in the eighties was a temporary aberration. This characterisation can be made of consumption for the Euro area. However, in contrast to consumption, the decline in the volatility of total investment can be characterised by a discrete reduction in the variance, which distinguishes investment from it's sub-components.

[13] Using a Markov switching vector autoregression, Artis *et al.* (2004) identify a common cycle between Germany, UK, France, Italy, Belgium, Netherlands, Austria, Spain and Portugal, for Europe.

[14] For a full discussion see Blanchard and Simon(2001) and Kim and Nelson (1999).

3.1 Multivariate Estimates of Break Dates

As put forward by Hansen (2001), a more precise estimate of the break date can be achieved when multivariate methods are utilised. Bai *et al.* (1998) show that there can be substantial gains from using multivariate inference about the break dates. To estimate common trend breaks in VAR's, the procedure follows that of Bai *et al.* (1998), which builds upon the work in Banerjee *et al.* (1992). The procedure is similar to that of the AR model above. The null of no break is tested against the alternative of a common break in the system of equations, using the QLR statistic which is computed using the VAR residuals. The empirical motivation concerns breaks in the mean growth rate, for which the parameters describing the stationary dependence in the stochastic part of the process - the AR parameters - are treated as nuisance parameters. In summary the Bai *et al.* (1998) test considers the null of a constant mean growth rate. All variables are transformed into $I(0)$ variables before any estimation is undertaken.

Table 4: Estimates of Common Break Dates of Variances of VAR Residuals

Variables	# vbles	QLR p-value	Break date	67% confidence interval
Total Consumption & investment	2	0.11	1993.1	1991:1 - 1995.1
Employment, Unit Labour Costs & Average Hours Worked	3	0.14	1994.2	1992.3 - 1996.1
Money Stock, Short-term Interest Rate & Long-term Interest Rate	3	0.00	1990.1	1989.1 - 1991.1
Exports & Imports	2	0.00	1991.1	1989.1 - 1992.2

Notes: The estimated breaks dates are computed using a method due to Bai *et al.* (1998)

Table 3 reports the OLS break date in the mean absolute residuals and the 67% confidence interval. The first VAR gathers the two main components of GDP in the Euro area. The second VAR captures labour market changes, the third VAR focuses on monetary factors and the fourth VAR captures the external sector. In the first VAR, the hypothesis of a constant variance is narrowly rejected at the 10% level. The third VAR reject the hypothesis of a constant variance at the one percent significance level. The estimated break ranges are all in the early nineties. However, the third VAR has a higher level of accuracy than the first two VAR's. The break date for the third VAR coincides exactly with the break date for output. From purely an objective viewpoint, it resorts one to ask the question of whether monetary factors were a key stabilising force for output.

The results from Table 3 and 4 suggest a break point that lies somewhere in the late eighties to early nineties, which coincides with an observed shift in the volatility of the permanent component of output. The results from Table 3 suggest a break model would be a suitable characterisation of consumption and investment, the two main components of real GDP. Hence, Table 3 suggests a break model would be best suited to further modelling of volatility in GDP. This stand in contrast to the production side, where total good production is best described by a discrete reduction in the variance.

In general, the weight of evidence suggests that the reduction in volatility are associated with changes in conditional means, θ_t, rather than conditional variances, ϵ_t. Accordingly, it can be concluded that real output has been experiencing lower levels of output growth rather than a sizeable reduction in volatility or a trend decline in volatility[15]. From this, it allows one to decipher that the stabilisation of output growth has been achieved at the expense of a slowdown in growth[16]. This result is also supported by Bai *et al.* (1998), who investigated the slowdown in the growth of output in the individual European economies.

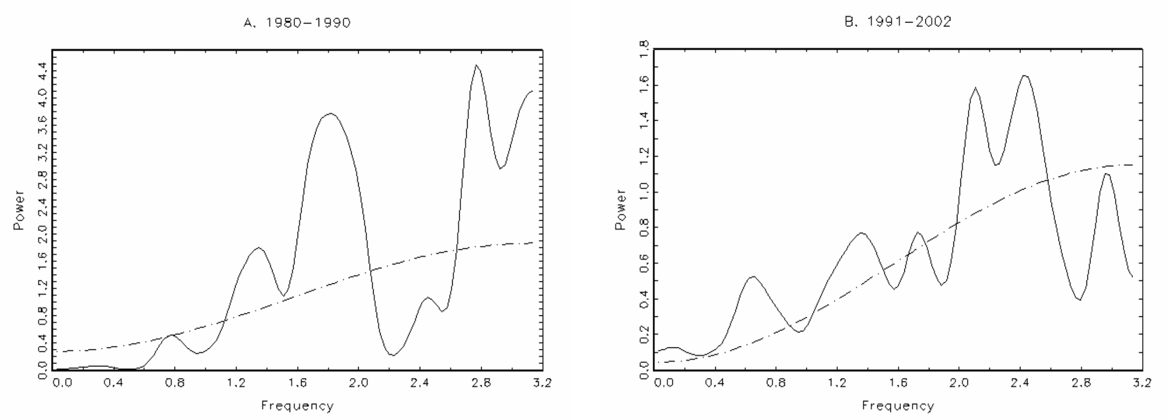
To finish, Figure 2 shows the spectral analysis for real output if one were to split the sample according to the break date for real GDP given in Table 3[17]. Two estimates are reported: a nonparametric estimator (smoothed periodogram) and a parametric IMA(1,1) estimator. The parametric estimate looks like a smoothed version of the nonparametric estimate, suggesting that the IMA(1,1) model fits the data reasonably well. Relative to the first period, the spectrum in the second period is lower in magnitude – this reflects the reduction in volatility between the two periods. Closer inspection reveals that the shape of the spectrum has changed, as well as its level, with the second period having relatively more power at lower frequencies than in the first. An 'eyeball econometric viewpoint' of Figure 2, would suggest the sample period either side of the break date are characterised by significantly differing levels of output growth volatility. This supports the result in Table 4 which shows the two main components of GDP, consumption and investment, reject the null of constant variance.

[15] This results differs from that of the largest Euro area, Germany. Mills and Wang (2003) found no structural break in the mean for Germany, but rather a shift break in volatility. However, they found that stabilisation of Italian business cycles has been achieved at the expense of a lower growth rate with similar evidence for France. The results for France and Italy are more closely aligned to that of the Euro area as a whole.

[16] Lower levels of growth has inflicted all major time series.

[17] Figure 2 allows by illustration, to decipher whether the the sample period after the break date, 1990, is associated with lower levels of output growth volatility relative to the first sample period. A significant change in the volatility between the two sample periods would infer an accurate break date, in which the decline in output growth volatility is associated with a single distinct break.

Figure 2:
Parametric and Nonparametric Estimates of the Spectrum for GDP



4 Impulse or Propagation

The univariate analysis infers that the moderation is perhaps due to breaks in the conditional mean. Hence, this section uses multiple sources of information to compute the conditional mean of output growth. This is achieved in a way similar to that of Ahmed *et al.* (2002), Boivin and Giannoni (2002) and Stock and Watson (2002a, 2003) by using VAR models. This section asks, is the observed reduction in volatility associated with a change in the magnitude of the VAR forecast errors - the impulses - or in the lag dynamics modelled by the VAR - the propagation - or both.

As of the results in Table 3, a break date of 1990 is imposed. Hence, the reduced form VAR is estimated in two separate time periods, 1980 - 1990 and 1991 - 2005. This will allow one to deduce how much of the reduction in mean output growth is due to changes in the VAR coefficients and the corresponding covariance matrix. The reduced form VAR takes the traditional form,

$$X_t = \Phi_i(L)X_{t-1} + u_t, \quad Var(u_t) = \Sigma_i \tag{1}$$

where X_t is a vector time series with the subscript i denoting the first and second period, $i = 1, 2$. The variance of the residuals is represented by Σ. The moving-average representation can be arrived at if $D_{i,j}$ is assumed to be the matrix of coefficients of the j^{th} lag in the matrix lag polynomial, hence $D_{i,j} = [I - \Phi_i(L)L]^{-1}$. This implies the variance of the k^{th} series in X_t in the i^{th} period is,

$$var(X_{kt}) = \left(\sum_{j=0}^{\infty} D_{ij}\Sigma_i D'_{ij} \right) = \sigma_k(\Phi_i, \Sigma_i)^2 \tag{2}$$

Equation (2) shows $\sigma_k(\Phi_i, \Sigma_i)$ to be the standard deviation of X_t in period i. From this one can calculate the counterfactual variance of X_{kt}. If for example $\sigma_k(\Phi_1, \Sigma_1)$, this would represent the standard deviation of X_{kt} in period 1. With this logic, $\sigma_k(\Phi_2, \Sigma_1)$ would be the standard deviation of X_{kt} if the lag dynamics had been those of the second period and the error covariance matrix been that of the first period. These expressions are based on the population parameters. The counterfactuals can be estimated by replacing the population parameters with sample estimators (Stock and Watson, 2002a). The results are presented in Table 5.

The results for GDP suggest that had the shocks of the 1980's occurred in the second time period, 1991-2005, the second period would have been as volatile as the first period. The counterfactual combination of second period dynamics and first period shocks, $\sigma(\Phi_2, \Sigma_1)$, produces an estimated standard deviation of 1.67, slightly lower than the first period standard deviation. In contrast, first period dynamics with second period shocks $\sigma(\Phi_1, \Sigma_2)$, produces a standard deviation result of 0.83. The results also suggest that monetary policy is much more reactive to shocks that hit the economy.

This result implies that had the shocks of the second period occurred in the first period, the first period, 1980 - 1990, would have been as quiescent as the second period. The changes in the covariance matrix of the unforecastable components of the VAR's - the impulses - account for a significant proportion of the reduction in the observed volatility of output. This result is supported by all the sensitivity analysis results in Table 5.B. This result is very similar to that found by Stock and Watson (2002a) for the US economy and support those conclusions made by Ahmed *et al.* (2002) and Boivin and Giannoni (2002), in which they conclude that the reduction in variance stems from smaller shocks, but also give particular weight to the changes in the propagation mechanism (40 to 60 percent respectively).

The results in Table 5 allow one to deduce that a significant part of the fall in the variance of the four-quarter growth of GDP for the Euro area, can be attributable to changes in the covariance matrix of the reduced form VAR innovations with an equal proportion also attributable to changes in the propagation mechanism ($\Phi_i(L)$).

Table 5: Implied Standard Deviations of Four-Quarter GDP Growth from Subsample VARs

Variable	Sample standard deviation		Standard deviation of four-quarter GDP growth in VAR model			
	1980-1990	*1991-2005*	$\sigma(\Phi_1, \Sigma_1)$	$\sigma(\Phi_2, \Sigma_2)$	$\sigma(\Phi_1, \Sigma_2)$	$\sigma(\Phi_2, \Sigma_1)$
y_t	1.93	1.38	1.98	1.28	0.83	1.67
π_t	1.15	0.65	1.53	0.71	0.74	1.41
r_t	1.68	1.33	2.20	1.30	1.01	2.60

B: Sensitivity Analysis

Deviation from Benchmark Specification	$\sigma(\Phi_1, \Sigma_1)$	$\sigma(\Phi_2, \Sigma_2)$	$\sigma(\Phi_1, \Sigma_2)$	$\sigma(\Phi_2, \Sigma_1)$
First Period – 1980-1990	1.67	1.02	0.65	2.74
VAR(6)	1.68	1.22	0.77	2.77
Levels data	1.56	1.19	0.70	2.43
Using the Long-term rate as a monetary policy indicator	1.70	1.25	0.96	2.61
Alternative commodity price index – Raw Materials Index	1.65	1.25	0.90	2.30
Commodity prices dropped	1.66	1.25	0.88	2.30
GDP replaced with production (total goods)$^{+}$	2.31	3.23	1.77	4.46
GDP replaced with private consumption	2.86	0.75	1.15	1.69
Replacing CPI with PPI^{+}	1.09	1.20	0.78	1.43

Notes: *The entries represent the square root of the variance of the four-quarter growth in GDP.*

These changes in the reduced-form VAR innovations could arise from reductions in the variance of certain structural innovations or from changes in the Euro area's economic ability to absorb such shocks, notably through changes in the priorities of monetary policy.

5 Explanations for the Great Moderation

This section considers four potential reasons to the above titled sectional heading. The first is sectoral shifts in the economy. While cyclically sensitive sectors such as manufacturing, which once constituted a large share of the G7 economies, those shares have fallen, coinciding with the rising importance of the service sector. As pointed out by Moore and Zarnowitz (1986), this shift should reduce the cyclical volatility of aggregate production. Secondly, the reduced form VAR impulse and propagation results, suggested a significant proportion of the decline in the variance of real GDP is attributable to changes in the covariance of the VAR innovations. The second category attempts to pinpoint the main types of shocks; money shocks, fiscal shocks, productivity/balanced growth shocks and oil/commodity price shocks. Thirdly, an investigation is undertaken that looks at the importance of improved monetary policy, through counterfactual simulation, for the moderation in GDP growth as suggested by Taylor (1998). Finally, the paper goes onto the examine the role of international shocks, utilising a common trends Factor Structural VAR (FSVAR), to gauge whether they have had a significant role in the reduction of variance in output growth[18].

5.1 Changes in the Sectoral Composition

It is in this subsection the analysis suffers from the lack of a data set in the tradition of the NIPA dataset for the US economy.

Table 6: The Effect of Changing Sectoral Composition on the Variance of GDP

	Standard Deviation		*Shares*	
	1991-1995	**1996-2005**	**1991**	**2005**
GDP (Actual)	0.1081	0.1005		
GDP (1995 shares)	0.1080	0.1003		
Agriculture	0.105	0.097	0.029	0.024
Manufacturing, Energy & Mining	0.120	0.096	0.235	0.216
Trade, Hotels, Transp, Comm.	0.108	0.098	0.210	0.211
Construction	0.117	0.113	0.063	0.057
Real Estate, Renting & Bus. Act	0.096	0.101	0.239	0.263
Pub.Admin,Education., Health & Oth. Services	0.111	0.104	0.220	0.225

Notes: *GDP is represented by Gross Value Added. The first row represents the standard deviation of the four quarter changes in the aggregate series. The preceding row shows the standard deviation of the 1995 share weighted share of four quarter changes in the disaggregated series shown in the other rows of the table.*

The data set is only available from 1991:1-2005:4. Even during this period, there has been a 0.8% reduction in the volatility of output growth. Without a dataset from the first sample period, 1980:1-1990:4, making an examination of the sectoral shift hypothesis remains difficult. However, Table 6 is included as an illustration to show that even during the past decade, a shift has been taking place towards the services sector of the economy, with the share of the service sector industries increasing as a percentage of GDP.

Stock and Watson (2003), using annual data spanning 1960-1997, found that in the case of Germany and France the shift in the structure of the economy had contributed 24 and 9 percent towards stabilisation respectively. In contrast, the result for Italy was minus nine percent. Since all three countries constitute over 70 percent of Euro area output, it can be deduced that the shift from manufacturing to services in the

[18] For a full summary see Stock and Watson (2003b) and Kose *et al.* (2001).

three constituent economies has been responsible for around 15 - 20 percent of the stabilisation in output growth.

5.2 Shocks and Surprises

The estimates from the previous section suggested that the decline in the variance of real GDP growth is partly attributable to changes in the covariance matrix of the VAR innovations. This has led many to claim that the volatility in output fluctuations in the seventies and eighties arose from misfortune like the oil price crises. Conversely, less pronouned shocks over the past decade are deemed to have contributed to the decline in economic activity. Resolution of this debate requires empirical tests. This subsection considers five types of shocks: money shocks, demand shocks, fiscal shocks, productivity shocks and oil/commodity price shocks.

Money Shocks Previous literature has tested a variety of models in hope of an accurate capture of a monetary policy shock. One of the most well known examples is Christiano *et al.* (1999). Using a Structural VAR (SVAR), the identification strategy for the Christiano *et al.* (1999) model is computed along with a sign restriction approach due to Uhlig (2005), with the computed strategy of Mountford (2005)[19], and a second sign restriction model due to Peersman and Straub (2004) with the same sign restrictions as the Mountford (2005) model.

The standard deviation of the Christiano *et al.* (1999) and sign restriction VAR monetary shocks in the 1991 - 2005 sample period, relative to the standard deviation in the earlier period, are reported. The results from both models suggest monetary shocks were more volatility in the first period relative to the second. The results from both models also infer that the reductions in the variance of monetary shocks have played a significant role in explaining the moderation of real output, asserting the importance of monetary shocks in determining output growth volatility. These assertions are also supported by the Peersman and Straub (2004) model.

[19]In Mountford's (2005) sign restriction VAR, a positive sign is placed upon the short-term interest rate and the exchange rate and a negative response on GDP deflator and money M1. The signs are in accord with the Mundell-Fleming Dornbusch model. No prejudgement is made with regards to output, hence no restriction is placed on output.

Demand Shocks Traditional Keynesian literature stressed the importance of demand-side innovations as significant contributors to fluctuations in economic activity. The results from the two models, Mountford (2005) and Peersman and Straub (2004), show that even though demand innovations are less volatility, such innovations have not played a contributing factor towards the stabilisation witnessed in the Euro area business cycle.

Fiscal Policy Shocks The first two rows in the fiscal policy shocks section, are calculated using a VAR sign restriction approach, with the restrictions in accord with Mountford and Uhlig (2005). The results for this model suggest a 25 - 30 percent reduction in fiscal policy shocks volatility. However, this model only predicts a very small contribution from fiscal policy shocks to GDP variance reduction. This result is similar to that found by Stock and Watson (2003) for the US economy, using the Blanchard and Perroti (2002) framework.

Table 7: Changes in the Standard Deviation of Various Macroeconomic Shocks

Shocks	Period 1	Period 2	$\frac{S_{period\,2}}{S_{period\,1}}$	Relative contribution to GDP variance reduction
Monetary Policy				
Mountford (2005) +	1981 -1990	1991 - 2002	0.94	0.68
Christiano-Eichenbaum-Evans[б] (1999)	1981 -1990	1991 - 2002	0.66	0.33
Peersman & Straub (2004)	1982 -1990	1991 - 2002	0.87	0.45
Demand Shocks				
Mountford+ (2005)	1982 -1990	1991 - 2002	0.57	-0.27
Peersman and Straub[µ] (2004)	1982 -1990	1991 - 2002	0.57	-0.12
Fiscal Policy				
Mountford & Uhlig (2005) – Spending *	1981 -1990	1991 - 2005	0.76	0.07
Mountford & Uhlig (2005) – Revenue*	1981 -1990	1991 - 2005	0.71	0.08
Productivity Shocks				
King *et al.* (1991)#	1981 - 1990	1991 - 2002	1.01	-0.08
Gali (1999, 2004)	1982 - 1990	1991 - 2002	0.84	-0.03
Blanchard-Quah (1989)	1981 - 1990	1991 - 2002	0.77	-0.10
Peersman & Straub (2004) – Labour	1982 - 1990	1991 - 2002	1.01	0.02
Peersman & Straub (2004) – Productivity	1982 - 1990	1991 - 2002	0.80	-0.52
Oil Prices				
Nominal Price	1980 -1990	1991 - 2005	1.02	-0.08
Real price	1980 -1990	1991 - 2005	1.01	-0.22
Hamilton (1996)	1980 -1990	1991 - 2001	0.99	-0.29
Commodity Prices				
All	1980 -1990	1991 - 2001	0.86	-0.02
Non-Fuel Primary Commodities	1980 -1990	1991 - 2005	0.79	0.16
Metals	1980 -1990	1991 - 2005	0.69	-0.11
Industry Materials Prices	1980 -1990	1991 - 2001	0.83	-0.09

Notes: +*The monetary shocks are derived from a sign restriction VAR model based on Mountford (2005). The restrictions are modelled on accepted priori beliefs of the effects of monetary shocks on the wider economy (also see Uhlig, 2005 and Leeper et al., 1996). The length of the shock, k, is set k=2 as in Uhlig (2005), an assumption which is also supported Christiano et al. (1999), who argues that monetary policy shocks do not usually last past one to two quarters. The demand shocks are modelled as in the standard macroeconomic textbook example, where a demand shock leads to a rise in output and prices.*

* *The fiscal shocks are derived in a fashion due to Mountford and Uhlig (2005), with the length of the shock, k, set to four. Two variables were used to derive the fiscal shocks, government expenditure and government revenue, as recommended by Mountford and Uhlig (2005).*

The balanced growth shock has been derived as in King et al. (1991), using a VECM model with long-run restrictions on y, c and i.

б *The relative contribution to GDP variance reduction result is calculated with a lag length seven, unlike the other results which are calculated with lag length 12. This is due to an exploding solution for the Christiano et al. (1999) model. The same also applies for the Hamilton oil price shock, which is estimated with lag length four.*

µ Peersman and Straub (2004) model a demand shock as a positive innovation in both output, prices and the policy interest rate.

Productivity Shocks Ever since Kydland and Prescott's (1982) seminal article traditional Real Business Cycle (RBC) theory has claimed a central role for exogenous variations in technology as a source of economic fluctuations in industrialised economies. However, standard measures of productivity shocks, such as the Solow residual, suffer from measurement problems, which include variations in capacity utilisation, imperfect competition and other sources (Stock and Watson, 2002a). Hence, this paper relies on four different models to capture productivity shocks. The first was suggested by King *et al.* (1991), which looked at balanced-growth innovations using a sign restriction vector error correction mechanism (VECM) model with assumptions from a one-sector real business cycle (RBC) model. The second is Gali (1999,2004), who imposed restrictions, in a SVAR framework with regards to a two-sector RBC model. The third model is that of Blanchard and Quah (1989), implementing a long-run restriction that demand shocks are neutral with respect to output. The final model follows Peersman and Straub (2004), in which they use a sign restriction VAR to capture labour supply and technology shocks, with technology shocks captured as a positive sign on output and wages with a corresponding negative sign on wages with labour supply shocks captured in a similar fashion except for wages falling.

Gali's (1999, 2004) productivity shock, which investigates the relationship between output and labour productivity per hour in a SVAR framework, shows a 16% reduction in volatility. However, Gali's (1999,2004) productivity shock has led to a very slight increase in real output volatility. In contrast to Gali's (1999) shock, the balanced-growth innovations from King *et al.* (1991), show a very slight rise in the volatility of productivity shocks. The result from the Gali (1999,2004) and the King *et al.* (1991) model, infer that productivity shocks have not played a positive role in the reduction of real output volatility. The same analysis and interpretation can also be applied to the shocks from the Blanchard and Quah (1989) model. This result would seem to suggest that productivity shocks, even though less frequent, have increased in magnitude with regards to their effect on real output. Lastly, the Peersman and Straub (2004) model finds that technology/productivity shocks have not played a positive role in moderating the cycle, despite there being a fall in volatility of technology shocks.

Oil Price Shocks The oil price shock section illustrates oil shocks calculated in real and nominal terms in quarterly growth rates. A third measure due to Hamilton (1996) is also included. Hamilton (1996), investigated the affects of asymmetric oil price shocks by measuring oil price innovations as the percentage difference between the current price and the maximum price during the previous year[20].

The nominal, real and Hamilton (1996) oil price shocks all declare near zero adjustment in the variability of oil shocks from the first period relative to the second. All oil price estimates suggest a negative relative contribution of oil price shocks to the reduction in the variance of real output. This is perhaps not surprising, since the second sample period includes the oil price hikes from the two Gulf war's and the very recent rises in crude oil prices due to rising demand from quickly growing developing economies like China and India. The first sample period was characterised by relatively stable oil prices compared to the seventies.

Other commodity price shocks The final section in the Table 7 show results for a wider variety of commodity prices, which include an aggregate of commodity prices, a non-fuel commodity price index which captures food prices changes, a metals and a wider industrial materials index. The estimates are calculated in the same fashion as the oil price shocks. The results suggest that the volatility in all four indices have fallen. Nonetheless, apart from the metals index, the commodities seem to have been a negative factor in the stabilisation of the Euro area cycle. The estimates suggest that although commodity price shocks are less frequent, their effect on Euro area output has increased in magnitude. In summation, despite a general fall in the volatility of commodity prices, they have had a negative effect on the relative contribution to real output variance reduction, suggesting that real output in the Euro area is more sensitive to commodity price changes.

As mentioned by Stock and Watson (2002a), it is tempting to add up the entries in the final column to produce a composite number, but this would be misleading. As is common understanding in structural shocks literature, it is often assumed with the innovations derived in the Table 7 that they are mutually uncorrelated. Yet as pointed out by Stock and Watson (2002a) and Rudebusch (1998), this is not always the case. There remains little consensus on whether these series are plausible proxies for the structural shocks they purport to estimate. Even so, ignoring the concerns just posited, it would appear that 30%[21] of the reduction in the conditional mean of output can be explained by the shocks above, implying by definition,

[20] The construction here ranges from 1980:1 2004:4 using the formula as in Hamilton (1996) = $\max(0, 100^* \{\ln(o_t) - \ln[\max(o_{t-1}, o_{t-2}, o_{t-3}, o_{t-4})]\}$, where o_t is the oil price variable.

[21] This result excludes demand shocks

that over 70% of the stabilisation in real output are *not* caused by the shocks in Table 7. This would advocate an examination of a much wider scope of innovations than that suggested in Table 7.

5.3 Institutional Change

Empirical studies, mainly on the US economy, have suggested that monetary policy change has played a significant role in reducing the fluctuations of output variability. An illustration is the case of the US economy where Clarida *et al.* (2000) estimate a large increase in the response to inflation of a Taylor-type monetary policy rule. There have been a number of studies investigating the extent to which a change in monetary policy has led to a reduction in the variance of output growth - see Clarida *et al.* (2000), Cogley and Sargent (2005), Boivin and Giannoni (2002), Gali *et al.* (2002) and Sims and Zha (2006)[22].

The general strategy in the literature has been to combine some structural intuition with VAR's that permit the model to fit the dynamic in the data, but within this general framework the details of the approach differ widely (Stock and Watson, 2003). As in Stock and Watson (2003), this paper uses a counterfactural policy evaluation performed using a SVAR with real GDP (y_t), GDP deflator inflation (π_t), a short-term interest rate (r_t) and a crude oil price commodity index (z_t).

The structural VAR identification is based on a model with an IS equation, a forward-looking New Keynesian Phillips Curve (NKPC), a forward looking Taylor-type monetary policy rule and a crude price index, which acts as an exogenous variable.

$$y_t = \theta r_t + \sum_{i=1}^{\infty} y_{t-j} + \epsilon_{y,t} \tag{3}$$

$$\pi_t = \gamma Y(\delta)_t + \sum_{i=1}^{\infty} \pi_{t-j} + \epsilon_{\pi,t} \tag{4}$$

$$r_t = \beta_\pi \overline{\pi}_{t+h/t} + \beta_y \overline{y}^{gap}_{t+h/t} + \epsilon_{r,t} \tag{5}$$

$$Z_t = \sum_{i=1}^{\infty} Z_{t-j} + \alpha_y \epsilon_{y,t} + \alpha_\pi \epsilon_{\pi,t} + \alpha_r \epsilon_{r,t} + \epsilon_{z,t} \tag{6}$$

where r_t represents the real rate of interest, which is defined as $r_t = i_t - \overline{\pi}_{t+k/t}$ in which $\overline{\pi}_{t+k/t}$ is the expected average inflation rate over the next k periods, where k is the term of the interest rate R_t

[22] See Mojon and Peersman (2001) and van Els, Locarno, Mojon and Morgan (2003) for a thorough review of the interaction of monetary policy and output in the Euro area.

and $Y(\delta)_t = \sum_{i=0}^{\infty} \delta^i y^{gap}_{t+1/t}$ is the discounted expected future output gap where $\overline{y}^{gap}_{t+h/t}$ is described as the expected future average output gap over the next h periods[23].

Equation (3) is an IS equation, with the following equation (4), a hybrid NKPC with a discount factor δ[24]. One arrives at (4) by solving this equation forward with $\delta = 1$. The NKPC allows for forward looking behaviour with δ interpreted as the weight on forward inflation (Gali and Gertler (1999), also see Gali *et al.*, (2001)). Equation (5) is a forward-looking real interest rate rule, a Taylor rule, where parameter h represents the horizon period, which is set at $h = 1$. Equation (5) contends the traditional trade-off between inflation and output stabilisation faced by central banks. The same short-term interest rate is used in both (3) and (6). Lastly, as is standard in SVAR analysis, the structural innovations ϵ_t, are assumed orthogonal.

Table 8: Implied Standard Deviation from Sample-Specific Structural VAR

A: Estimated Taylor Rule Coefficients, Benchmark Specification
θ=-0.2, δ=0.5, γ=0.3

	β_π	β_y
Sample Period 1	-0.891 (0.173)	0.353 (0.451)
Sample Period 2	0.416 (0.243)	0.432 (0.328)

B: Implied Standard Deviations of Four-Quarter GDP Growth, Benchmark Specification

Variable	Sample Standard Deviation		Standard deviations implied by VAR							
			VAR with $\Phi=\Phi_1$				*VAR with* $\Phi=\Phi_2$			
	1980-1990	1991-2005	Ω_1, A_1	Ω_1, A_2	Ω_2, A_1	Ω_2, A_2	Ω_1, A_1	Ω_1, A_2	Ω_2, A_1	Ω_2, A_2
GDP	1.33	1.16	1.22	1.43	0.66	0.78	2.02	1.89	1.49	1.19
Inflation	1.15	0.65	1.25	1.17	1.32	0.84	1.29	1.44	1.01	0.86
Monetary Policy Rate	1.68	1.33	1.67	1.42	1.90	1.00	2.16	2.21	1.91	1.29

C: Sensitivity Analysis: Alternative Parameter Values[+]

IS and Phillips curve Parameters			*Estimated Taylor Rule Coefficients*				*Standard deviations implied by VAR*					
			Period 1		Period 2		*VAR with* $\Phi=\Phi_1$			*VAR with* $\Phi=\Phi_2$		
θ	γ	δ	β_π	β_y	β_π	β_y	Ω_1, A_2	Ω_2, A_1	Frac Var_1	Ω_1, A_2	Ω_2, A_1	Frac Var_2
-0.20	0.30	0.90	0.15	-0.56	0.51	-0.78	1.77	1.17	-0.19	1.86	2.07	1.84
-0.20	0.30	0.10	-0.62	0.54	-0.53	0.18	1.63	0.88	0.11	2.22	1.34	0.21
-0.20	0.10	0.50	-0.63	0.57	-0.40	0.07	1.70	0.84	-0.03	2.41	1.22	0.01
-0.20	0.60	0.50	-0.15	-0.13	0.36	-0.65	1.63	0.99	0.11	2.19	1.48	0.47
-0.10	0.30	0.50	-0.46	0.32	0.10	-0.40	1.73	0.82	-0.11	2.85	1.01	-0.29
-0.50	0.30	0.50	2.12	0.70	1.72	-0.63	2.08	0.93	-0.98	3.22	1.71	0.94
-0.20	0.10	0.90	-0.12	-0.18	0.39	-0.67	1.62	1.04	0.13	2.08	1.54	0.58
-0.20	0.30	0.75	-0.08	-0.23	0.42	-0.70	1.63	1.14	0.12	1.97	1.80	1.14
-0.20	0.10	0.75	-0.45	0.30	0.11	-0.41	1.72	0.82	-0.08	2.83	1.02	-0.28
-0.50	0.10	0.75	0.69	1.18	2.01	-0.49	1.62	0.98	0.13	2.19	1.52	0.55
-0.00	0.30	0.50	-0.72	0.29	-0.37	-0.33	1.88	0.74	-0.45	3.31	0.91	-0.42

Notes: *Data series runs from 1980:1 till 2005:4. The two sample periods are 1980-1990 and 1991-2005. The Frac* Var_1 *is the ratio* $[\sigma^2(\Phi_1, \Omega_1, A_1) - \sigma^2(\Phi_1, \Omega_1, A_2)]/[\sigma^2_1 - \sigma^2_2]$ *and Frac* $Var_2 = [\sigma^2(\Phi_2, \Omega_2, A_1) - \sigma^2(\Phi_2, \Omega_2, A_2)]/[\sigma^2_1 - \sigma^2_2]$.

[23] It must be noted that this assumption assumes certain restrictions and technology and the labour market structure within a local neighbourhood of the steady state real marginal costs - see Rotemberg and Woodford (1997).

[24] With $\delta = 0$, it would represent the traditional New Keynesian Phillips curve.

Estimation of the model relies on *apriori* knowledge of the three key parameters θ (the slope of the IS curve), γ (slope of the Phillips relation) and δ (parameter governing the forward-looking properties of the Phillips curve relationship). It must be noted that there remains little agreement over the correct parameter values for θ, γ and δ[25]. For the Euro area, Gali *et al.* (2001) find δ to be 0.088 using a Calvo (1983) specification. Their study further finds that backward price setting has been a relatively unimportant factor behind the dynamics of Euro area inflation, which allowed Gali *et al.* (2001) to construe that backward looking behaviour is unimportant for the Euro area. This low discount rate figure stands in contrast to the high figure set by Tillmann (2005), who simulated a variety of models with δ set between 0.91 and 0.98. He finds the fit improves with lower values of δ. In further contrast to Gali *et al.* (2001), McAdam and William (2004) find a more balanced role for backward and forward looking components in the estimation of a NKPC for the Euro area.

For the Phillips curve relationship, O'Reilly and Whelan (2005) estimate $\gamma = 0.596 - 0.675$, which stands in contrast to the negative coefficients found by Gali *et al.* (2001) for the Euro area. A simple snapshot of the results for the Euro area reveals very little agreement over the correct calibration parameters. Hence, the benchmark model is calibrated with the conventional loadings, also used by Stock and Watson (2002a), where $\theta = -0.2$, $\gamma = 0.3$ and $\delta = 0.5$, which are assumed to remain constant over the sample period.

Estimation is undertaken by first running a reduced form VAR of the all the variables in the four equation system and replacing the variables by the reduced form VAR residuals. The reduced form VAR residuals can be interpreted as forecasts of the output gap and inflation. Next, innovations in the expected future gap are replaced with innovations in expected future output, which is plausible if one assumes that the forecast errors of trend output are negligible. This implies that with θ, γ and δ given, the innovations ϵ_y and ϵ_π follow equations (3) and (4). Finally, equation (6) is estimated by OLS.

The analysis here is similar to that conducted in Table 5. Table 8 is characterised by three sets of parameters; the VAR distributed lag coefficients Φ, the covariance matrix of the innovations $\Omega = (\epsilon_y, \epsilon_\pi, \epsilon_r, \epsilon_z)$ and finally A, which represents the structural coefficients $(\theta, \gamma, \delta, \beta_\pi, \beta_y, \alpha_y, \alpha_\pi, \alpha_r)$ that link the structural innovations and reduce form residuals. Hence, $\sigma(\Phi_i, \Omega_j, A_k)$ estimates are presented in Table 8, where i, j

[25] See Gali *et al.* (2002), Rudebusch (2002), Clarida *et al.* (2000) and Rudebusch and Svensson (1999).

and k represent the two sample periods.

Table 8 show the results for the model presented in Equations $(3)-(7)$. The Ω parameter - the covariance matrix - represents the change in the variability of output which can be attributable to shocks, with (A) corresponding to changes in the variability of output attributable to policy. The results are presented for the two sample periods, 1980-1990 and 1991-2005. The estimated Taylor rule coefficients in Table 8.A find that the inflation response in the first period is negative. The second period is characterised by a larger output coefficient β_y, and a positive inflation response. These results are consistent with those found for the US economy. Starting with (Φ_1, Ω_1, A_1), the standard deviation of the output growth is 1.22 in comparison to (Φ_2, Ω_2, A_2), which has a standard deviation value of 1.19. These estimates are calculated from the sample moments of GDP. These results infer that monetary policy has become more sensitive to changes in output. The results conjecture that changes in the monetary policy coefficients have infact contributed to increasing output variability by around 12%. A delineation implying that most of the reduction in variability in output is due to smaller shocks and not to changes in the monetary policy coefficients. The results for the other sets of calibrated parameter values are shown in Table 8.C.

5.3.1 Quantitative Evidence from Two Macro Models

The finding of changing monetary policy coefficients suggest that structural shifts in monetary policy have occurred. Consequently, this section investigates whether the long-term decline in volatility may be partly attributable to the gradual development of macroeconomic policy and policy makers' long and variable learning curve[26]. Here the effect of improved monetary policy on output volatility through counterfactual simulations of a changing monetary policy rule is estimated. Such an analysis will allow one to question Martin and Rowthorn (2005) and Stock and Watson (2003) contention that the quiescence of the past fifteen year could well be a hiatus before a return to more turbulent economic times. This is achieved by estimating what the standard deviation of output growth would have been under a counterfactual environment in which monetary and structural factors is of post-1993 but subjugated to pre-1990 shocks[27].

[26]The learning legacy is made up of lender of last resort facilities, deposit insurance, financial safety nets and automactic fiscal stabilisers.

[27]The analysis in this subsection will provide an insight into whether the stability currently enjoyed by the Euro area cycle, will endure for the long-term.

In addition to the Stock and Watson (2003) model from the previous section, the Rudebusch and Svensson (1999) model is also estimated counterfactually. The Rudebusch and Svensson (1999) model consists of three equations.

$$\Delta\pi_{t+1} = \alpha_0 + \alpha_{\pi 1}\Delta\pi_t + \alpha_{\pi 2}\Delta\pi_{t-1} + \alpha_{\pi 3}\Delta\pi_{t-2} + \alpha_y y_t^{gap} + \epsilon_{t+1} \quad (7)$$

$$y_t^{gap} = \beta_0 + \beta_{y1} y_t^{gap} + \beta_{y2} y_{t-1}^{gap} + \beta_r(\overline{R}_t - \overline{\pi}_t) + \eta_{t+1} \quad (8)$$

$$R_{t+1} = \phi_0 + \phi_{R1} R_t + \phi_{R2} R_{t-1} + \phi_\pi \overline{\pi}_{t+1} + \phi_{y1} y_{t+1}^{gap} + \phi_{y2} y_t^{gap} + \psi_{t+1} \quad (9)$$

Equation (7) represents a Phillips curve where π_t and y_t represent inflation and the output gap. Equation (8) represents the IS curve, where $\overline{R}_t$ and $\overline{\pi}_t$ are the four quarter averages of the short-term interest rate and inflation. The model is closed with Equation (9), which is a Taylor rule equation from Judd and Rudebusch (1998). The model, as before, is estimated in the two sample periods - full coefficient results for the Rudebusch and Svensson (1999) model are shown in Appendix B.

Table 9

A: The Effect of Improved Monetary Policy on Output Volatility

Model	Standard Deviations		Percent of Variance Reduction Explained
	Base Model	*Base+pre-1990 Monetary Policy*	
Rudebusch-Svensson	0.85*	0.90*	4%
Stock-Watson SVAR	0.85*	1.20*	30%
Smets & Wouters†	1.63	1.88	26%
Historical Values			
Period	1993-2002	1980-1990	
Standard Deviation	1.43	2.10	

B: The Effect of Smaller Shocks on Output Volatility

Rudebusch-Svensson	0.85	1.33	52%
Stock-Watson SVAR	0.85	0.97	10%
Historical Values			
Period	1993-2002	1980-1990	
Standard Deviation	1.43	2.10	

Notes: * *Based on Simulation from 1980:1-2002:4.*
† Source: Stock and Watson (2003) – *The results have been calculated using the Base + pre-1979 monetary policy shocks with the same calibration parameters as in Smet and Wouters (2003)*

The results from both models in Table 9.A are congruous. The Rudebusch-Svensson (1999) model reveals that monetary policy has had a positive impact in its relative contribution to output stabilisation, whereas the Stock and Watson (2003) model result reveals that up to 30% of the reduction in the variance of output growth is due to improved monetary policy[28], concurring with Rudebusch-Svensson (1999) model. As noted in Stock and Watson (2003), the two models tested here focus on the use of the short-term interest rate as a tool for achieving inflation and/or output stabilisation goals over the short to medium term. However, central banks have a much wider remit than that considered here. Such responsibilities include, short-term crisis management, such as providing liquidity and preventing financial crises. Hence, it is possible that the

[28] See Stock and Watson (2003a) for a more complete explanation.

reduced volatility of output is in part a result of better management by the monetary authorities, a channel not addressed by conventional models of monetary policy transmission[29].

Further, Table 9.B analyses output volatility under a 'big shock' counterfactual scenario. This is undertaken by estimating what the standard deviation of output would have been under a counterfactural scenario in which monetary policy and the economic structures are reflected in the post 1993 environment, with the economy subjected to shocks as large as those of pre-1990. This estimation is undertaken with both the Stock and Watson (2003) and the Rudebusch-Svensson (1999) models. The estimations suggest that in both the Stock and Watson (2003) and the Rudebusch-Svensson (1999) models, four quarter growth would have been larger than its actual post-1990 value. The Rudebusch-Svensson (1990) model indicates that the decreased shock volatility explains about 50% of the variance reduction from pre-1990 to post-1993. In both models, the output volatility increase arising from using pre-1990 shocks is much larger than the increase from using pre-1990 monetary policy, suggesting that shocks more disperse than monetary shocks are important in explaining the variance reduction in real output, supporting the assertions made in previous sections of this paper. The proposition that the stabilisation witnessed in the business cycle is as a result of missing shocks is consistent with the sectoral evidence presented in Table 3, which exhibited a widespread decline in volatility across sectors and other real activity measures. This pattern is coherent with what one would expect if little changed on the real side of the economy, except that the standard deviations of all economic shocks fell.

As noted by J. Bradford Delong, the results concerning the behavioural change of the monetary authorities to changes in economic activity do raise specific issues. The idea that policy is limited to systematic reactions by the monetary authorities that change interest rates in response to changes in inflation and output is a limited definition of policy. It will not capture the fact that central banks in the Euro area may have reacted

[29] Assertions that improved monetary policy is the cause of business cycle moderation concentrate around a few key hypotheses. The first being unstable equilibria. Monetary policy in the 1980s is characterised by stop-go monetary policies, in which the brakes on an over-heating economy were applied too hard and too late. As a result of this, inconjunction with the fact that the econometric models tested above being linear, the results above may not address the stop-go hypothesis. The second is the anchored inflation expectations hypothesis. The models tested above imply a fully credible central bank and the central banks long-term inflation target is known. However, whether inflation expectations are anchored, as is assumed by the models, it is difficult to assess directly, but what evidence there is suggests that if inflation expectations are anchored this is a quite recent phenomenon. It is difficult to argue that inflation expectations were anchored in the mid-eighties. This second hypothesis is not addressed by the two models either. Tests for non-linearity behaviour in the Taylor rule are shown in Appendix D. The linear Taylor rule is extended to contain nonlinear terms, such as a threshold once inflation reaches a certain level. Appendix D estimates a variety of nonlinear extensions of dynamic Taylor rules but finds scant evidence of nonlinearities, such as threshold effects, that match descriptions of stop-go policies. Even if there were a nonlinear, as a statistical matter the nonlinear policy seems to be well approximated by the linear Taylor-type rules summarized in Table 9.

in one way when it was worried about price stability, and reacted in a different way when faced with a similar economic situation when it was worried about maximum purchasing power.

5.4 International Shocks and Synchronisation

One important, relatively recent, branch of research now focuses on whether cross-country linkages in growth have shifted, perhaps in a way that can help rationalise the variance reduction (Doyle and Faust, 2005). There are many frameworks available for developing an econometric model, which permits the answering of how much as a fraction of a country's cyclical variance is due to international shocks and how these shocks have evolved over time. The econometric model needs to resolve the issue of how best to identify an international shock.

Stock and Watson (2003) identify four alternatives, and most commonly employed, econometric models which could be utilised to capture international shocks. Firstly, a world shock could be estimated as an innovation in a univariate time series model of world GDP growth. There are however, limitations to this framework. Since US output receives a great weight in the four economies being considered here, it may confound world shocks with US shocks and idiosyncratic shocks to other large economies. Assuming no common world shock or the presence of international trade, this identification scheme would nonetheless attribute a large fraction of US output fluctuations to a common shock as an arithmetic implication of its construction. The second modelling framework, which overcomes some of the flaws of a univariate model, utilises a parametric dynamic factor model as in Kose *et al.* (2001) and Waston (1994), where the number of shocks is greater than the number of series and the comovements across series at all leads and lags are attributed to the common shock. This results in an unobserved components model that can be estimated using Kalman filtering. Undertaking such a framework has one hypothetical advantage. In the case of no economic spillovers and no common shock, the estimation results would indicate no comovements with the common shocks being correctly identified as having zero variance. Yet due to the cross-dynamics being associated to the world shock, this approach is perhaps not best suited to identifying the separate effects of a common world shock and any spillovers arising through trade. The third approach focuses upon the use of non-parametric methods to estimate a dynamic factor model. As in Stock and Watson (2002b), if a large number of series have a dynamic factor structure, then the common component or the common dynamic factor can be estimated using principal components. This procedure has been used by Helbling and Bayoumi (2003) to estimate the importance of common factors in G7 fluctuations and also by Helg *et al.* (1995) to extract European industry and country specific shocks. The notion that the principle components/nonparametric approach has the advantages of the second approach without the disadvantage of assuming that all comovements stem from the common disturbance rather than through trade spillovers, is tainted by the fact that individual countries are sometimes necessarily heavily weighted, like the US, leading to the same disadvantage as the first approach.

The fourth approach, which is employed here, adopts a VAR framework allowing for lagged effects and the identification of world shocks as those that affect all economies within the same period. These economies include the UK, Japan and the US. A similar econometric model was also exploited by Altonji and Ham (1990) and Norrbin and Schlagenhauf (1996). The factor structure allows a decomposition of the $h-$step ahead forecast error for GDP growth into three sources; unforeseen common shocks, unforeseen domestic shocks and spillover effects arising from unforeseen domestic shocks to other countries in the model. The econometric model used here is succinctly nested in equation 10,

$$\Delta y_{1,t} = \alpha_1 \Delta y_{1,t-1} + b_1 \Delta y_{2,t-1} + \epsilon_{1,t} + c\,\eta_t \tag{10}$$

In Equation (10), $\Delta y_{1,t}$ is the growth of output in country 1, $\epsilon_{1,t}$ is the country-specific shock for country 1, and η_t is the common world shock. Similar equations characterise other countries in the model. If there is a world shock it will affect output growth in all countries, although the magnitude of that effect will differ from country to country. Country-specific shocks affect their own country directly, with spillovers due to international linkages between countries. In this framework, cross-country correlations depend on the magnitudes of the various shocks and their effect on the economies (Stock and Watson, 2003)[30].

Yet, equation (10) contains more shocks than observable variables. There are four countries and one common shock, which in total concurs five shocks. As a result, estimation requires factor models. In a similar fashion, Monfort *et al.* (2002) model the international linkages as arising entirely from current and

[30] Also see Helbling and Bayoumi (2003).

lagged effects of the common international shock. Supplementary to the international shock, all shocks which are country-specific, ω_t, have an international transmission requiring around one quarter i.e. 'spillovers'. The model considered in Equation (10) has the following econometric assumptions,

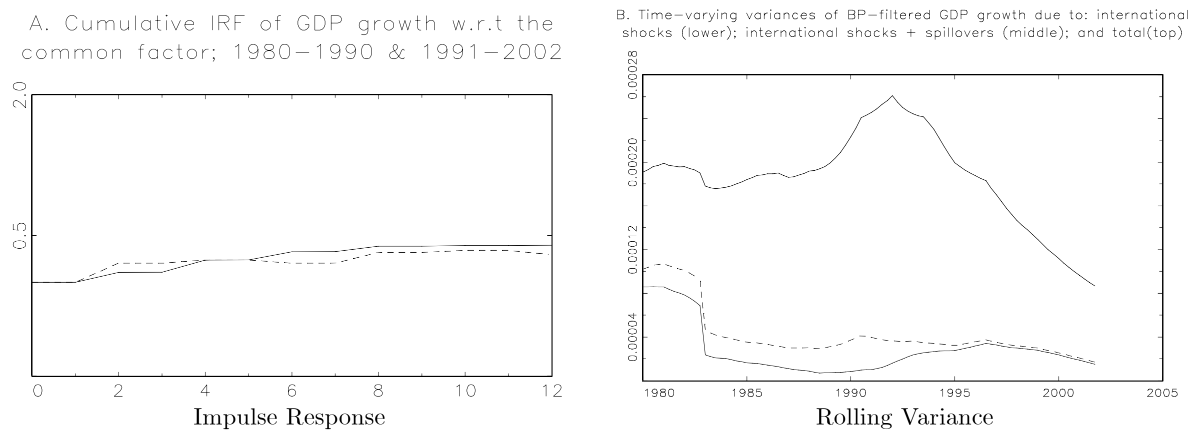
$$\epsilon_t = \Gamma f_t + \omega_t, \text{ where } E(f_t f_t') = \Sigma_{ff} = diag(\sigma_{f1}, ..., \sigma_{fk}) \; \forall \; t \text{ and } E(\omega_t \omega_t') = diag(\sigma_{\omega 1}, ..., \sigma_{\omega 4}) \tag{11}$$

where f_t is a $kx1$ vector that denotes the common international factors, secondly, Γ is the $4xk$ matrix of factor loadings and ω_t are the country-specific idiosyncractic country shocks, with the standard normalisation assumptions applied $E(\omega_t) = 0$ and $E(\omega_t \omega_s') = 0 \; \forall \; s \neq t$. In addition to the standard normalisation assumptions $E(f_t, \omega_s') = 0 \; \forall s, t$, which postulates that shocks in ω_t are contemporaneously and intertemporally uncorrelated and may have different variances. As a result, this decomposition in turn permits a decomposition of the variances of the $h-$step ahead forecast error. Equation (11) identifies international shocks are those shocks that affect output in multiple economies, within the quarter, contemporaneously. The FSVAR is estimated using Gaussian maximum likelihood.

Likelihood ratio tests are undertaken to determine the number of factor loadings. The results are presented in Appendix C. In both sample periods and the pooled sample, the hypothesis of $k = 1$ cannot be rejected against the alternative, that of the covariance matrix, Σ_ϵ, having full rank. In contrast, the null $k = 2$ can be rejected at the 99% significance level. The results advise that $k = 1$ is appropriate, so an adopted specification with one common international shock i.e. one common factor, is estimated.

Table 10 summarises the variance decomposition for GDP growth and for the band-pass filtered GDP for the Euro area. The comparative importance of international shocks, which are decomposed into common shocks or spillovers, can be measured as one minus the share of the forecast error variance attributed to domestic shocks. A general overview of the results suggest that international shocks are playing a more important role in influencing the Euro area cycle.

Figure 3– Impulse Response and Rolling Variance Estimate



At the eight quarter horizon, $h = 8$, international shocks are responsible for 25% of the fraction of the forecast error variance in Euro area output, with spillovers accounting for one-fifth at the same horizon period. As a result, one sees a fall in the importance of domestic shocks. This result ties in with Stock and Watson (2005), who find that for most G-7 economies, international shocks are playing a greater role at the expense of domestic shocks. The results are also supported by Helbling and Bayoumi (2003), who found evidence of international shocks explaining an increasing amount of the variation in output for the industrialised economies.

Table 10
A: Two Factor FSVAR: Common Shocks, Spillovers and Own Country Shock Variance Decomposition

		1980-1990				1991-2005			
		Forecast error	*Fraction of Forecast error variance due to*			Forecast error	*Fraction of Forecast error variance due to*		
Economy		standard deviation	Int'l Shocks	Spillovers	Own Shock	standard deviation	Int'l Shocks	Spillovers	Own Shock
Euro Area	1	1.90	0.06	0.00	0.94	1.41	0.24	0.00	0.76
	2	1.34	0.06	0.00	0.94	1.00	0.24	0.00	0.76
	4	1.11	0.04	0.04	0.91	0.68	0.23	0.15	0.62
	8	0.97	0.03	0.07	0.90	0.44	0.25	0.20	0.55
UK	1	1.12	0.71	0.00	0.29	0.45	0.70	0.00	0.30
	2	0.79	0.71	0.00	0.29	0.32	0.70	0.00	0.30
	4	0.67	0.74	0.06	0.20	0.26	0.62	0.11	0.27
	8	0.70	0.75	0.11	0.14	0.28	0.57	0.21	0.22
US	1	0.89	0.56	0.00	0.44	0.91	0.74	0.00	0.26
	2	0.63	0.56	0.00	0.44	0.65	0.74	0.00	0.26
	4	0.59	0.62	0.01	0.37	0.59	0.77	0.02	0.21
	8	0.67	0.67	0.03	0.30	0.52	0.77	0.10	0.13
Japan	1	0.92	0.49	0.00	0.51	0.38	0.44	0.00	0.56
	2	0.65	0.49	0.00	0.51	0.27	0.44	0.00	0.56
	4	0.60	0.61	0.04	0.35	0.27	0.35	0.05	0.61
	8	0.66	0.68	0.09	0.23	0.31	0.34	0.10	0.56

B: Band-Pass Filtered GDP

Euro Area	0.66	0.04	0.09	0.87	0.40	0.19	0.23	0.57
UK	0.36	0.69	0.16	0.15	0.14	0.47	0.29	0.23
US	0.31	0.61	0.06	0.32	0.34	0.65	0.11	0.24
Japan	0.33	0.57	0.16	0.27	0.16	0.31	0.18	0.51

C: Decomposition of Changes in the Variance of 4-quarter-ahead FSVAR Forecast Errors into Changing Impulses and Changing Propagation

	Variances			*Contribution of change in Shock spillover*				*Contribution of change in Impulse function*			
	1980-1990	1991-2005	change	Int'l	Spill over	own	total	Int'l	Spill over	own	total
Euro-Area	1.24 (0.30)	0.56 (0.17)	-0.68 (0.34)	-1.28 (0.40)	0.00 (0.01)	-0.55 (0.22)	-1.82 (0.52)	1.25 (0.51)	0.00 (0.02)	-0.10 (0.11)	1.15 (0.56)

Notes: *The results illustrate the standard deviation and three-way decomposition of variance of filtered versions of GPD. Panel (a) shows results for FSVAR forecast errors at the 1, 2, 4 and 8 quarter horizon. Panel (b) shows results for the ideal (infinite order) 6-32 quarter band pass filtered values of GDP. In Panel (c), the first three columns give the variance of BP-filtered GDP (in percentage points) in the first and second subsample, using the estimated FSVAR and their difference. The remaining columns decompose this difference into changes in the impulse response functions and changes in the variances of the shocks themselves. The sum of the "international," "spillover," and "own" columns equals the "total" column, and the sum of the two "total" columns equals the "change" column. Estimated standard errors are shown in parentheses.*

The variance for band-pass filtered GDP allows one to draw similar a conclusion to the variance decomposition results of GDP growth in Table 10.A. Figure 3.B presents the time-varying estimates of the variance decomposition of band-pass filtered output, which is based on rolling estimates of the one-factor FSVAR. Figure 3.B illustrates the time-varying estimates of the variance decomposition of bandpass-filtered GDP, based on rolling estimates of the one-factor FSVAR. The lower line in Figure 3.B is the contribution to the variance of the international shocks, the middle dotted line is the sum of the contributions of the international shocks and spillovers and the top line is the total variance. Hence, the gap between the top and middle lines is the contribution to the variance of domestic shocks. As found by Stock and Watson (2005) for Germany, Italy, UK and the US the recent decline in overall volatility for the Euro area, tracks the decline in the variance arising from international shocks along with a large historical decline in the variance associated with the importance of domestic shocks.

Table 10.C investigates the principle that the contribution of international shocks to output volatility could decrease because the variance of the international shocks has decreased, because a shock of a fixed magnitude has less of an effect on the economy, or both. Said differently, the variance of GDP growth in the Euro area may have changed because the magnitude of the shocks impinging on the Euro area economy have changed or because the effects of those shocks have changed. As in Stock and Watson (2003), the variance of output growth in the two sample periods, 1980-1990 and 1991-2002, are decomposed into changes in the magnitudes of the shocks (impulses) and changes in their effects on the economy (propagation). This is formally modelled as

$$V_p = V_{p,1} + ... + V_{p,5} \tag{12}$$

where V_p denotes the variance of the four-quarter-ahead forecast errors in a given country in period p to each of the five shocks. Thus the change in the variance between the two periods is $V_2 - V_1 = (V_{2,1} - V_{1,1}) + ... + (V_{2,5} - V_{1,5})$. In an identified SVAR, the variance component $V_{p,j}$ can be re-written as $a_{pj}\sigma_{pj}^2$ where a_{pj} is a term depending upon the squared cumulative impulse response of GDP to shock j in period p with σ_{pj}^2 is the variance of shock j in period p, leading to an expression where the contribution of the j^{th} shock can be decomposed exactly as,

$$V_{2,j} - V_{1,j} = \left(\frac{a_{1j} + a_{2j}}{2}\right)(\sigma_{2j}^2 - \sigma_{1j}^2) + \left(\frac{\sigma_{1j}^2 + \sigma_{2j}^2}{2}\right)(a_{2j} - a_{1j}) \tag{13}$$

Equation (13) decomposes the variance into the contribution from the change in the shock variance plus the contribution from the change in the impulse response. This decomposition requires that the covariance matrix of the factors $\Sigma_{ff} = diag(\sigma_{f1}, ..., \sigma_{fk})$ and the factor loadings, Γ, to be identified separately.

Table 10.C presents the decomposition of the change in variance of four quarter-ahead forecast errors in Euro area output. Changes in the variance of shocks led a large and statistically significant decline in volatility. Indeed, the decline in shock variances more than accounts for the drop in the variance of real output forecast errors. The results in Table 10.C also contend that the decline in variance is not attributed to changes in the propagation mechanism, but due to changes in the size of the shocks, which is partially supported by Figure 3.A, which examines whether there have been important changes in the effect of an international shock of a fixed magnitude on the Euro area cycle. The impulse response function in Figure

3.A, with respect to the common factor, suggests that the magnitude of the effect of the common shocks has changed little, with the estimated responses to the common factor shock being relatively close to zero.

Given the results in Table 10, it is perhaps a surprise to find that there has been no increase in the synchronisation of business cycles among the industrialised economies.

Table 11: Correlations of GDP Growth Across Countries
Four-quarter growth rates, simple correlation coefficients

	USA	Japan	UK	Euro Area
US	1.00			
Japan	0.45	1.00		
UK	0.47	0.54	1.00	
Euro Area	0.42	0.39	0.56	1.00
		1991-2005		
US	1.00			
Japan	0.13	1.00		
UK	0.23	-0.06	1.00	
Euro Area	0.53	0.07	0.52	1.00

The previous two decades have seen common international shocks to have risen and increased in importance as a determinant of output fluctuations. These common international shocks however, have become very slightly smaller in magnitude, implying that despite their increasing effect, the net result is that international correlations have seen little increasing synchronisation. This finding is similar to that of Stock and Watson (2005) and Doyle and Faust (2005), in that they emphasise the importance of the reduction in the variance of the shocks, in this case the common international shock, complementing the results in the previous section of this paper.

6 Conclusion

There is evidence of a decline in the volatility of economic activity measured by both broad aggregates and by a wide variety of other series that track specific facets of economic activity. For real output growth, the decline is best characterised as a break model, with a sharp drop from 1991 onwards. This decline in real output growth coincides with similar declines witnessed in consumption and GFCF investment. The short and long-term interest rates have shown a slight rise in volatility.

An explanation for the stabilisation in output growth finds many possible causes for the moderation. In addition, less volatile monetary policy has also played a role. However, this leaves more than half of the decline in volatility unaccounted for. Identifiable shocks, such as productivity and oil price shocks have played no role in the stabilisation of real output. The evidence from the reduced-form model in section 4 asserts that the stabilisation in output is associated with an increase in the precision of forecasts of output growth.

With improved monetary policy attaining little recognition, it would imply that the moderation in real output will continue even with a change in the policy regime. Further as seen in section 4, a significant proportion of the reduction seems to be due to good luck in the form of smaller economic disturbances, which also leaves the Euro area with the same unsettling conclusion of that found for the US economy by Stock and Watson (2002a), that the quiescence of the past two decades could well be a hiatus before a return to more turbulent economic times. In other words, the reduction in the output business cyclle is down to good luck and not skill: primarily to smaller shocks hitting the Euro area economy and secondary to better monetary policy.

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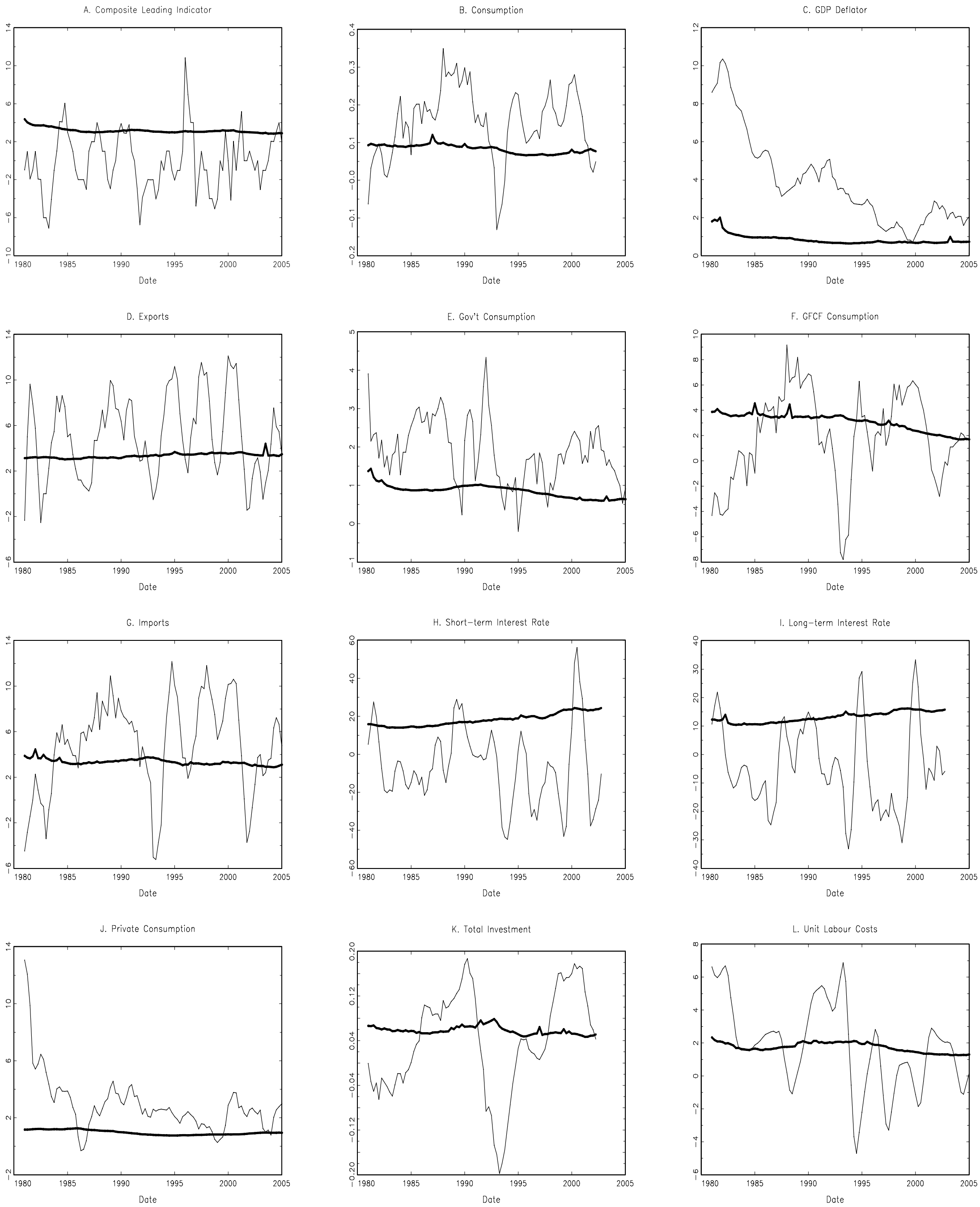
Appendix

A: Volatility results for the principal series



Appendix B
Parameter Estimates for the Rudebusch-Svensson Model

Parameter	1980:1 – 2005:4	1980:1 – 1990:4	1991:1 – 2005:4
α_0	-0.077 (0.155)		
$\alpha_{\pi 1}$	-0.303 (0.113)		
$\alpha_{\pi 2}$	-0.489 (0.096)		
$\alpha_{\pi 3}$	-0.168 (0.110)		
α_y	0.207 (0.075)		
β_0	0.110 (0.086)		
β_{y2}	1.174 (0.112)		
β_{y2}	-0.263 (0.106)		
β_r	-0.033 (0.022)		
φ_0		2.647 (0.772)	3.429 (0.644)
φ_{R1}		0.970 (0.156)	0.919 (0.094)
φ_{R2}		-0.335 (0.157)	-0.427 (0.097)
φ_π		0.012 (0.073)	0.347 (0.092)
φ_{y1}		0.230 (0.136)	0.023 (0.096)
φ_{y2}		0.190 (0.190)	0.171 (0.114)
σ_ε	0.617		
σ_η	1.515		
σ_ξ		0.557	0.444

Notes: *Heteroskedastic robust standard errors are given in the parenthesis.*

Appendix C: Tests of *k*-factor FSVAR vs. Unrestricted VAR

Number of factors (*k*)	d.f.	1980-2005 LR Statistic	1980-2005 *p*-value	1980-1990 LR Statistic	1980-1990 *p*-value	1991-2005 LR Statistic	1991-2005 *p*-value
1	2	3.86	0.14	1.23	0.54	4.30	0.11
2	1	0.005	0.00	0.714	0.00	1.5e-005	0.00

Appendix D: Tests for Nonlinearity

$$R_{t+1} = \phi_0 + \phi_{R1} R_t + \phi_\pi \overline{\pi}_{t+1} + \phi_{y1} y_{t+1}^{gap} + \phi_{y2} y_t^{gap} + \psi_{t+1}$$

Regressor	Base Model			
Baseline Regressors				
constant	2.16 (0.56)	2.76 (0.90)	1.52 (0.44)	1.17 (0.52)
R_{t-1}	1.10 (.0.01)	0.98 (0.12)	1.07 (0.14)	1.05 (0.22)
R_{t-2}	-0.51 (0.14)	-0.50 (0.18)	-0.42 (0.15)	-0.27 (0.12)
$\overline{\pi}$	0.36 (0.14)	0.47 (0.20)	0.34 (0.14)	0.15 (0.13)
y_t^{gap}	0.09 (0.12)	0.03 (0.14)	0.52 (0.23)	0.06 (0.10)
y_{t-1}^{gap}	0.11 (0.13)	1.17 (0.16)	-0.27 (0.19)	0.00 (0.12)
Additional Regressors				
$1(\overline{r}_{t-1} > F_{\overline{r},0.75})x\overline{r}_{t-1}$		-0.20 (0.52)		
$1(\overline{r}_{t-1} < F_{\overline{r},0.25})x\overline{r}_{t-1}$		0.93 (0.40)		
$1(\overline{r}_{t-1} > F_{\overline{r},0.75})$		1.28 (2.27)		
$1(\overline{r}_{t-1} < F_{\overline{r},0.25})$		-3.64 (1.60)		
$1(y_t^{gap} > F_{y^{gap},0.75})xy_t^{gap}$			-0.44 (0.29)	
$1(y_t^{gap} < F_{y^{gap},0.25})xy_t^{gap}$			-0.66 (0.47)	
$1(y_t^{gap} > F_{y^{gap},0.75})$			-0.71 (0.96)	
$1(y_t^{gap} < F_{y^{gap},0.25})$			1.50 (0.60)	
$1(\overline{\pi}_t - \overline{\pi}_{t-8} > F_{\overline{\pi}-\pi_{-8},0.75})$ x $(\overline{\pi}_t - \overline{\pi}_{t-8})$				8.52 (3.20)
$1(\overline{\pi}_t - \overline{\pi}_{t-8} > F_{\overline{\pi}-\pi_{-8},0.75})$ x $\overline{\pi}_t$				-1.83 (0.66)
$1(\overline{\pi}_t - \overline{\pi}_{t-8} > F_{\overline{\pi}-\pi_{-8},0.75})$				0.09 (0.50)
F-statistic (p-value) for exclusion of additional regressors		1.80 (0.12)	3.96 (0.00)	5.62 (0.00)

Notes: *Tests for nonlinearities were carried using the above equation estimated over 1980:1-1990:4. The tests were conducted by adding several "threshold" variables to the base specification. To define these threshold variables, let Fx,0.75 denote the 75th percentile of the empirical distribution of x over the 1980-1990 sample period, and let Fx, 0.25 be similarly defined. Let* $r_t = \overline{R}_t - \overline{\pi}_t$. *The table below shows results with additional variables, the estimated coefficients, standard errors and F-statistics for joint significance.*

Appendix E: Additional Results for the Rudebusch-Svensson and Structural VAR Models

	Rudebusch-Svensson			Stock and Waston		
	Base Model	Base+pre-1993 Monetary policy	Base+pre-1993 shocks	Base Model	Base+pre-1993 Monetary policy	Base+pre-1993 shocks
$\sigma(y_t - y_{t-4})$	0.85	0.90	1.33	0.85	1.20	0.97
$\sigma(\pi_t - \pi_{t-4})$	1.49	1.56	2.53	1.48	1.49	1.86
$\sigma(\overline{\pi})$	1.44	1.45	1.96	-	-	-
$\sigma(\overline{y})$	-	-	-	0.97	1.20	1.60

Notes: $\sigma(y_t - y_{t-4})$ denotes the standard deviation of $y_t - y_{t-4}$, *similarly for* $\pi_t - \pi_{t-4}$, $\sigma(\overline{y})$ *and* $\sigma(\overline{\pi})$.