# Output Fluctuations in the United States: What Has Changed Since the Early 1980s?

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**Abstract**

We document a structural break in the volatility of U.S. GDP growth in the first quarter of 1984, and provide evidence that this break emanates from a reduction in the volatility of durable goods production. We find no evidence of increased stability in the nondurables, services or structures sectors of the economy. In addition, no other G7 country experienced a contemporaneous reduction in output volatility. Finally, we show that the reduction in durables volatility corresponds to a decline in the share of durable goods accounted for by inventories.

# 1 Introduction

*From boardrooms to living rooms and from government offices to trading floors, a consensus is emerging: The big, bad business cycle has been tamed.*

-The Wall Street Journal, Nov 15, 1996

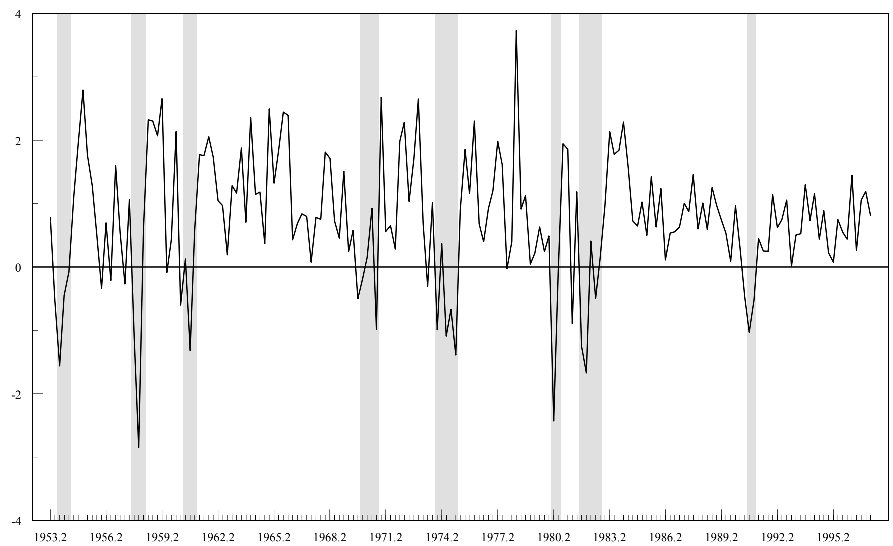
The business press is currently sprinkled with references to the 'death' or 'taming' of the business cycle in the United States. While such claims are undoubtedly premature, they are in part rooted in the apparent reduction in the volatility of U.S. output fluctuations over the period beginning in the early 1980s. Figure 1 plots the growth of real U.S. GDP over the period 1953:2 to 1997:2; the variance of output fluctuations over the period ending in 1983 is more than four times as large as the variance for the period since 1984.

In this paper, we document a structural break in the volatility of U.S. GDP growth in the first quarter of 1984. We begin with an example in which we show that a regime switching model of output growth fails to capture a business cycle signal when the model is augmented to allow both the mean and the variance of output to switch between states. To explain the absence of the business cycle signal, we appeal to the dominant effect of a one-time decline in the variance of GDP growth in the early 1980s, and support this claim by endogenously estimating a structural break in the residual variance of an AR specification for output growth in the first quarter of 1984.

As a means of understanding this dramatic reduction in U.S. output volatility, we first examine the output series of each of the other G7 countries for a break contemporaneous to the one which occurred in the U.S. We then decompose output growth into its component parts and provide evidence that the break emanates from a reduction in the volatility of durable goods production. We further show that the break in durables is roughly coincident with a break in the proportion of durables accounted for by inventories.

The break in output volatility affects the implementation of a range of simulation and econometric techniques. For example, one common method for taking theory to the data is to compare the moments of data generated from calibrated models with

Figure 1: Real U.S. GDP Growth: 1953:2 to 1997:2




the moments of actual data. The presence of a one-time reduction in output volatility in the early 1980s clearly affects the time horizon over which the second and higher moments of output growth should be computed.

On the empirical front, the volatility break implies that linear models for output growth over periods that span the break are misspecified. In addition, signal-to-noise ratios in state-space characterizations of business cycle fluctuations, such as dynamic factor or Markov-switching models, will be reduced when the variance is modeled as constant. Indeed, we present one important example of this in the paper. Finally, the reduction in the variance of output fluctuations should alter the interpretation policy makers place on a particular realization of quarterly GDP growth; what may have been considered a moderate decline in activity prior to the break may now be viewed as severe.

The paper proceeds as follows. In Section 2 we use both the empirical business cycle methodology and structural stability tests to characterize the changes in the process for output in recent years. Section 3 examines both international and disaggregate U.S. data in order to better understand the source of the break in output volatility. In Section 4 we outline and discuss a set of candidate explanations for the volatility decline. Section 5 concludes.

## 2 The Decline in U.S. Output Volatility

There is a large literature which explores the question of whether the magnitude or duration of economic fluctuations have changed across the pre- and post-WWII periods (examples include DeLong and Summers (1986), Romer (1986a, 1986b, 1989, 1994), Shapiro (1988), Diebold and Rudebusch (1992), Lebergott (1986) and (Watson (1994)). While the evidence on this particular issue is mixed (resulting in no small part from the difficulties associated with the construction of comparable data series across the two periods), the more general pursuit of documenting changes in the process governing output fluctuations is an important element of macro economic research. Such documentation is valuable both because it leads to a collection of macro economic stylized facts and because it may provide insight into whether such changes are likely to be permanent or temporary.

In this section we characterize recent changes in the process for U.S. output growth. We do so by focusing on quarter-to-quarter fluctuations in the growth rate of GDP, rather than on changes in the business cycle per se. In addition, since we are interested in understanding the rather dramatic reduction in output volatility in the most recent two decades relative to the previous three, we use only post-war data and thereby avoid the problems associated with pre-and post-war data comparability.

We begin by showing that the widely used regime switching framework is no longer a useful characterization of business cycle movements when we allow both the mean and the variance of output to switch between states. We then document a structural break in the residual variance of an AR specification for output growth in the first quarter of 1984 and show that there are no corresponding breaks in the autoregressive coefficients. Finally, we present an illustrative exercise in which we show that the ability of the switching-mean model to identify 1990-91 recession depends critically on the exclusion of the high-variance years from the estimation.

## 2.1 The Empirical Business Cycle

The starting point for our analysis is motivated by the empirical business cycle literature spawned by Hamilton (1989). In his paper, Hamilton uses a regime switching framework to show that by allowing the mean of the process to switch between states, one can capture the periodic shifts between positive and negative real GDP growth in the U.S.. He further shows that such shifts accord well with the NBER business cycle peaks and troughs. A number of researchers have since found this to be a useful approach to characterizing business cycles, including Lam (1990), Phillips (1991), Jefferson (1992), Ghysels (1993), Boldin (1994), Durland and McCurdy (1994), Filardo (1994), Kim (1994), and Diebold and Rudebusch (1996).

Following this literature, we estimate a Markov switching model for the rate of growth of real GDP for the period 1953.2 - 1997.2 by considering a latent variable, $S_t$, which represents different states of output growth.[1] Conditional on the value of $S_t$, the expected value for the rate of growth of GDP, denoted $\dot{y}_t$, is:

$$E(\dot{y}_t|S_t = i) = \mu_{S_t} \tag{1}$$

In addition, we assume that $S_t$ follows a first-order Markov chain, and therefore,

$$P(S_t = i|S_{t-1} = j, S_{t-2} = k, \ldots) = P(S_t = i|S_{t-1} = j) = p_{ij} \tag{2}$$

We can rewrite Equation 1 as:

$$\dot{y}_t = \mu_{S_t} + u_t \tag{3}$$

where $u_t$ represents other factors that affect the dynamics of $\dot{y}_t$ and $E(u_t) = 0$. As in Hamilton, we model $u_t$ as following an AR(p) process, but we slightly modify his original specification by allowing both the mean and the variance of the AR(p) model

[1]We use chain-weighted GDP data, as constructed by the Bureau of Economic Analysis.

to switch between states. Adding the AR(p) specification to Equation 3 we obtain:

$$\dot{y}_t = \mu_{S_t} + \sum_{j=1}^{p} \phi_j(\dot{y}_{t-j} - \mu_{S_{t-j}}) + \epsilon_t, \qquad \epsilon_t \sim N(0, \sigma_{S_t}) \tag{4}$$

where i can assume two alternative values: 1 or 2.[2] These two values are commonly interpreted as indicating periods of recession and expansion. Therefore, $\mu_{S_t=1}$ is the expected value of the rate of growth of GDP during recessions and $\mu_{S_t=2}$ is the expected value for expansions.

We test two alternative restricted models (Model 1 and Model 2) against the unrestricted model shown in Equation 4 (Model 3). In particular, Model 1 restricts $\sigma_{S_t=i} = \sigma$, but allows for different means. Model 2 imposes the restriction that $\mu_{S_t=i} = \mu$ but allows the variances to differ across states. Model 3 allows both the mean and the variance to switch across states.[3]

Table 1 reports the results of this exercise for an AR(1) specification for output growth.[4] The subscripts on the parameters indicate states. The transition probabilities $p_{11}$ and $p_{22}$ give estimated probabilities of moving from state 1 to state 1 and from state 2 to state 2, respectively. Looking first at the p-values on the tests of the restrictions imposed by Models 1 and 2, we find that we can reject the constant mean model (Model 1) in favor of Model 3, but that we can not reject the constant variance model (Model 2) in favor of Model 3.[5]

Given that the switching mean typically captures the business cycle signal, and that we can reject the switching mean model, but not the switching variance model, what is the nature of the signal being captured by our estimation? To answer this,

[2]Equation 4 shows that the value of $\dot{y}_t$ depends not only on the state of the economy in period t, but also on the state of the economy in periods t-1…t-p. We therefore have $2^{p+1}$ states of the economy.

[3]We also test Model 3 against a linear specification for GDP growth. This test is discussed below.

[4]We do not arbitrarily impose the AR(1) structure on the data. Instead, in this exercise and throughout the paper, we explicitly test for the best AR characterization of the data. In this case, we report the results of the specification test against the AR(4) in the last line of Table 1 (because Hamilton estimates an AR(4)). Our result that the AR(1) is the best model is consistent with the findings of Hess and Iwata (1997).

[5]Though not reported, the qualitative nature of the results are unchanged for the AR(4).

Table 1: Markov-Switching Model: U.S. Real GDP Growth

| **AR(1) Specification - 1953:2 to 1997:2** | | | | | | |
|---|---|---|---|---|---|---|
| | **Model 1** | | **Model 2** | | **Model 3** | |
| $\mu_1$ | -1.09 | (0.49) | 0.70 | (0.09) | 0.65 | (0.11) |
| $\mu_2$ | 0.90 | (0.13) | | | 0.79 | (0.15) |
| $\phi$ | 0.35 | (0.10) | 0.35 | (0.07) | 0.33 | (0.07) |
| $\sigma_1^2$ | 0.65 | (0.11) | 0.24 | (0.05) | 0.23 | (0.05) |
| $\sigma_2^2$ | | | 1.18 | (0.16) | 1.18 | (0.15) |
| $p_{11}$ | 0.40 | (0.29) | 0.99 | (0.02) | 0.99 | (0.01) |
| $p_{22}$ | 0.95 | (0.03) | 0.99 | (0.01) | 0.99 | (0.01) |
| LL value | -238.13 | | -227.71 | | -227.45 | |
| p-value (vs. Model 3) | 0.00 | | 0.48 | | | |
| p-value (vs. AR(4)) | 0.11 | | 0.31 | | 0.31 | |

Note: Standard errors appear in parentheses below coefficient estimates. Model 3 is the unrestricted model. Model 1 restricts the variance to be constant, Model 2 restricts the mean to be constant. We report the AR(1) specification since we cannot reject the AR(1) in favor of the AR(4) for any of the models. The p-values for the test of the AR(1) versus the AR(4) are reported in the last line of the table.

we plot GDP growth along with the smoothed state 1 probabilities from each of the three models. These plots are shown in Figure 2. State 1 probabilities indicate the probability of being in the low mean state in the case of Model 1, the low variance state for Model 2 and the low mean-low variance state for Model 3.

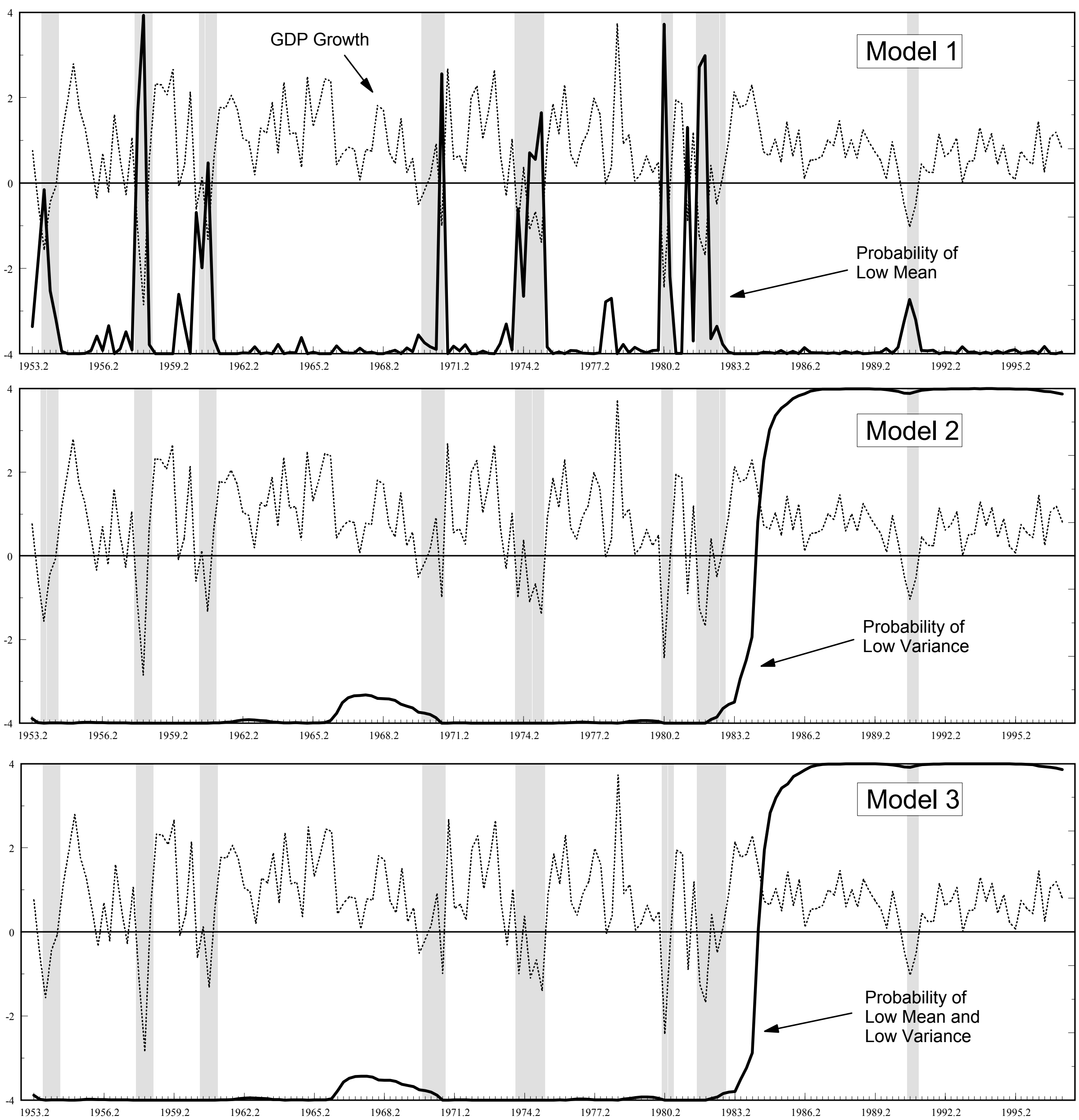
The smoothed probabilities for Model 1 are shown in the top panel of Figure 2. The pattern corresponds closely to the business cycle, as measured by the NBER turning points (we will return to the implications of the model's difficulty in picking out the 1990-91 recession). The second panel plots the probabilities from Model 2. There is a one-time switch to the low-variance state in the early 1980s. The last panel plots the probabilities from Model 3. Recall that state 1 for this model is the low mean-low variance state. The close correspondence between the patterns in panels 2 and 3 makes it easy to understand our failure to reject Model 2; the business cycle signal in the data is virtually swamped by the dramatic reduction in the variance in the early 1980s, and thus, modeling this signal adds very little to the data characterization.[6]

Up to this point we have imposed the switching specification on the data. It is important to test explicitly for whether we can reject a linear model (constant mean and constant variance) for GDP growth in favor of the switching model. Since under the null of a linear model the regularity conditions necessary to conduct the LM, LR or Wald tests are not met, we use the approximation proposed by Hansen (1992, 1994).[7]

[6] Note that when we allow both the mean and the variance to switch between states (see Table 1), we find that the low mean and low variance states occur together. Ramey and Ramey (1996) use a panel data set across countries to show that higher volatility in the rate of growth of GDP is associated with slower average rate of growth. Thus our time series finding does not accord well with their cross sectional result.

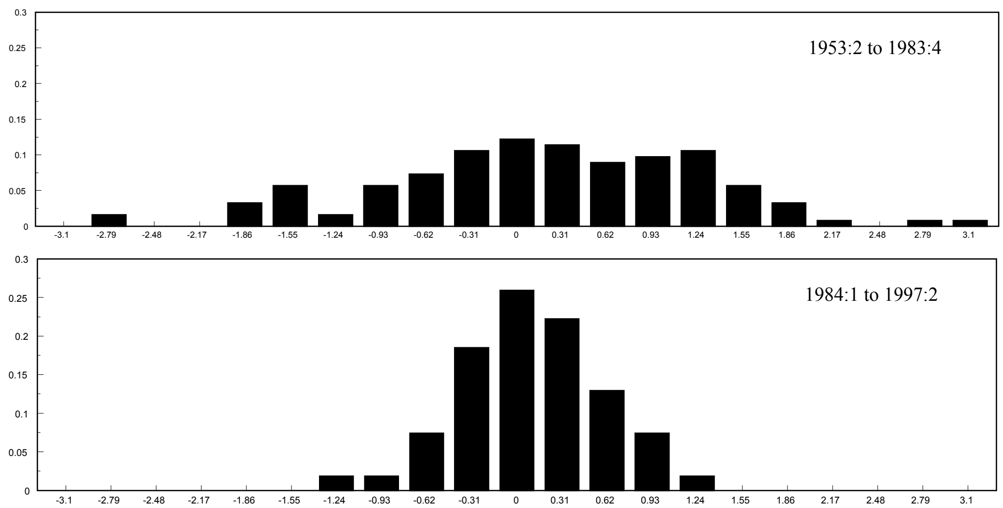
[7] To do this, we define $\alpha = (\mu_{S_t=2} - \mu_{S_t=1})$, $\beta = (\sigma_{S_t=2} - \sigma_{S_t=1})$, $\mu = \mu_{S_t=1}$ and $\sigma = \sigma_{S_t=1}$, and rewriting Equation 4, we obtain: $\dot{y}_t = (\mu + \alpha D_t) + \sum_{j=1}^{p} \phi_j(\dot{y}_{t-j} - \alpha D_{t-j}) + \epsilon_t$, with $\epsilon_t \sim N(0, (\sigma + \beta D_t)^2)$, where $D_t$ is an indicator variable that is equal to 0 when $S_t = 1$ and equal to 1 when $S_t = 2$. The test requires one to compute the constrained estimates of the likelihood function over a grid of possible values for the set of parameters, $\Theta$, that under the null hypothesis of the linear model do not converge to any fixed population parameters. In our case, $\Theta = (\alpha, \beta, p_{11}, p_{22})$. We define the grid of values for the elements of $\Theta$ in the following way: $\alpha$ ={0.01 to 0.20 in intervals of 0.01}, $\beta$ ={0.05 to 1.0 in intervals of 0.05}, $p_{11}$ ={0.981 to 0.997 in intervals of 0.004} and

Figure 2: Smoothed State 1 Probabilities: Models 1, 2 and 3




Note: Figure plots GDP growth along with the smoothed State 1 probabilities from the estimation of Models 1,2, and 3. State 1 corresponds
low mean state for Model 1, the low variance state for Model 2, and the low mean-low variance state for Model 3.

Figure 3: Distribution of Residuals from Linear Specification for GDP Growth




We find that we can reject the null of the linear model in favor of the switching specification with a p-value of 0.003.[8] The intuition behind this result is illustrated in Figure 3. This figure shows two histograms which are produced by dividing the residual variance from the linear model into the two subperiods suggested by the plot of the smoothed probabilities (i.e., we split the sample in the fourth quarter of 1983). The reason for the rejection of the linear model is obvious; the variance for the period 1951 to 1983 is more than four times larger than the variance for the period 1984 to 1997.

## 2.2 Structural Change

The pattern of the smoothed probabilities shown in Figure 2 suggests the possibility that GDP growth is better characterized by a process with a structural break in the

$p_{22}$ ={0.899 to 0.999 in intervals of 0.002}. This grid implies that the space for $\Theta$ is partitioned into 5625 points.

[8]As suggested in Hansen (1994), we use a consistent kernel estimator to account for serial correlation. We use a bandwidth size of four, but the results are robust to bandwidth sizes of three, two, one and zero.

variance in the early 1980s than by a switching regime. In this section, then, we use structural stability tests to endogenously estimate a break date for the variance of GDP growth and provide a measure of the statistical significance of our estimated breaks.

Drawing on our previous result that GDP growth is best characterized by an AR(1), we test for a structural break in the residual variance from the following specification for GDP growth:

$$\dot{y}_t = \mu + \phi \dot{y}_{t-1} + \epsilon_t \tag{5}$$

If $\epsilon_t$ follows a normal distribution, $\sqrt{\frac{\pi}{2}}|\hat{\epsilon}_t|$ is an unbiased estimator of the standard deviation of $\epsilon_t$. Therefore, we look for a break in an equation of the form:

$$\sqrt{\frac{\pi}{2}}|\hat{\epsilon}_t| = \alpha + \mu_t \tag{6}$$

where $\alpha$ is the estimator of the standard deviation.[9]

We estimate a break point by jointly estimating the following system using GMM:

$$\dot{y}_t = \mu + \phi \dot{y}_{t-1} + \epsilon_t \tag{7}$$

$$\sqrt{\frac{\pi}{2}}|\hat{\epsilon}_t| = \alpha_1 D_{1t} + \alpha_2 D_{2t} + \mu_t \tag{8}$$

where

$$D_{1t} = \begin{cases} 0 & \text{if } t \leq T \\ 1 & \text{if } t > T \end{cases}$$

$$D_{2t} = \begin{cases} 1 & \text{if } t \leq T \\ 0 & \text{if } t > T \end{cases}$$

[9] In the absence of the normality assumption, $|\hat{\epsilon}_t|$ in Equation 8 can be interpreted as an estimator of the standard deviation.

and T is the estimated break point, and $\alpha_1$ and $\alpha_2$ are the corresponding estimators of the standard deviation. The list of instruments for each period $t$ is as follows: a constant, $\dot{y}_{t-1}$, $D_{1t}$, and $D_{2t}$.[10]

The appearance of the parameter T under the alternative hypothesis but not under the null implies, as in the case of the Markov switching versus the linear model, that the LM, LR and Wald tests of equality of the coefficients $\alpha_1$ and $\alpha_2$ do not have standard asymptotic properties.

Andrews (1992) and Andrews and Ploberger (1994) develop tests for cases such as this, when a nuisance parameter is present under the alternative but not under the null. They consider the function, $F_n(T)$, where n is the number of observations, which is defined as the Wald, LM or LR statistic of the hypothesis that $\alpha_1 = \alpha_2$, for each possible value of T. We assume that T lies in a range $T_1, T_2$.[11] Andrews (1992) shows the asymptotic properties of the statistic:

$$\sup_{T_1 \le T \le T_2} F_n = \sup F_n(T) \tag{9}$$

and reports the asymptotic critical values. In this test, the T that maximizes $F_n(T)$ will be the estimated date of the break point.

However, Andrews and Ploberger (1994) show that this test is not optimal and instead propose the following statistics:

$$\exp F_n = \ln(1/(T_2 - T_1 + 1)) * \sum_{T=T_1}^{T_2} \exp(1/2 * F_n(T)) \tag{10}$$

and

$$\text{ave}\, F_n = (1/(T_2 \_ T_1 + 1)) * \sum_{T=T_1}^{T_2} F_n(T). \tag{11}$$

They prove the optimality of these statistics for the case in which a nuisance parameter is present under only the alternative hypothesis. The p-values associated with these

[10] The results for $\hat{\epsilon}_t^2$, the estimator of the variance, are very similar to those reported below.

[11] Following Andrews and Andrews and Ploberger, we set $T_1 = .15 * n$ and $T_2 = .85 * n$.

statistics are computed using the approximation suggested by Hansen (1997).[12]

The results of the tests for structural change in the residual variance of the process for the growth rate of GDP are reported in the top panel of Table 2. Each of the three test statistics presented indicates a strong rejection of the null that $\sigma_1 = \sigma_2$, and the estimated break date occurs in the first quarter of 1984. The timing of this break corresponds closely with that suggested by the smoothed probabilities displayed in Figure 2.[13]

We now consider the possibility that the break in the residual variance results from a break in the AR coefficients. In particular, we estimate:

$$\dot{y}_t = \mu_1 D_1 + \mu_2 D_2 + \phi_1 \dot{y}_{t-1} D_1 + \phi_2 \dot{y}_{t-1} D_2 + \epsilon_t \tag{12}$$

where $D_1$ and $D_2$ are as defined above.

We first test jointly for a break in the mean and the coefficient on lagged GDP growth, and then for a break in each of the mean and the lag coefficients separately.[14] These results are reported in the bottom panel of Table 2. In all cases, we cannot reject the null of no break. In fact when we conduct a Chow test and actually impose the estimated break date of 1984:1, we still cannot reject the null of no break. The p-value associated with the LR statistic for the Chow test is reported in the last column of Table 2. We therefore have strong evidence that the break in the variance in 1984:1 is not due to a change in the AR components of the model.[15]

[12]The optimality results are valid only for the case of a Wald or LM test, but not the LR test. In terms of a comparison between the Average and Exponential statistics, the Average has better properties for alternatives close to the null, while more distant alternatives are better tested by the Exponential.

[13]An alternative possibility is that the reduction in output volatility was gradual rather than discrete. The plot of GDP growth shown in Figure 1, however, suggests that the extreme movements in output growth in the late 1970s and early 1980s make it unlikely that we would reject our discrete break in favor of a gradual change. Given this, and the fact that the choice of functional forms would in any case be arbitrary, we view it as a simplifying assumption to model the change as discrete rather than gradual, and do not rule out the possibility that the change actually took place over several quarters.

[14]For simplicity we use a LM test rather than a Wald. Andrews and Ploberger (1994) prove that these two tests are equivalent.

[15]This result suggests that there has been a change in the amplitude of output fluctuations, but

Table 2: Structural Break Tests: U.S. Real GDP Growth - 1953:2 to 1997:2

| **Specification:** $\dot{y}_t = \mu + \phi \dot{y}_{t-1} + \epsilon_t$ | | | | |
|---|---|---|---|---|
| $\epsilon_t \sim N(0, \sigma_t^2)$, **where** $\sigma_t^2 = \sigma_1^2$ **if** $t \leq T$, **and** $\sigma_t^2 = \sigma_2^2$ **if** $t > T$ | | | | |
| **Residual Variance** | | | | |
| **Null** | **Sup** | **Exp** | **Ave** | |
| $\sigma_1^2 = \sigma_2^2$ | 15.43 | 5.12 | 4.96 | |
| | (0.00) | (0.00) | (0.00) | |
| Estimated break date: **1984:1** | | | | |
| **AR Coefficients** | | | | |
| **Null** | **Sup** | **Exp** | **Ave** | **Chow** |
| $\mu_1 = \mu_2$, $\phi_1 = \phi_2$ | 2.77 | 0.64 | 1.20 | 0.32 |
| | (0.99) | (0.85) | (0.75) | (0.85) |
| Estimated break date: **none** | | | | |
| $\mu_1 = \mu_2$ | 2.09 | 0.40 | 0.72 | |
| | (0.93) | (0.61) | (0.53) | |
| Estimated break date: **none** | | | | |
| $\phi_1 = \phi_2$ | 2.73 | 0.46 | 0.81 | |
| | (0.82) | (0.55) | (0.47) | |
| Estimated break date: **none** | | | | |

Note: p-values appear in parentheses below test statistics. Chow test imposes the 1984:1 estimated break date (for the residual variance) on the AR(1) coefficients.

Table 3: U.S. Real GDP Growth: AR(1)

| **Specification:** $\dot{y}_t = \mu + \phi \dot{y}_{t-1} + \epsilon_t$ | | | |
|---|---|---|---|
| Sample | $\hat{\mu}$ | $\hat{\phi}$ | $\hat{\sigma}^2$ |
| 1953:2 to 1997:2 | 0.50 | 0.34 | 0.9055 |
| | (0.09) | (0.07) | |
| 1953:2 to 1983:4 | 0.53 | 0.33 | 1.1974 |
| | (0.12) | (0.08) | |
| 1984:1 to 1997:2 | 0.40 | 0.41 | 0.2552 |
| | (0.11) | (0.12) | |

Note: standard errors appear in parentheses below coefficient estimates.

Table 3 reports the estimated parameters for the full sample and for the two subsamples implied by the estimated breakdate for the residual variance. The similarity of the AR estimates across subsamples, along with the difference in the variance estimates, is not surprising in light of the results of the tests for structural change.

Finally, we test for the presence of additional breaks, conditional on having found the first break in 1984:1, by repeating the break tests for the two subsamples implied by that breakdate (1953:2 to 1983:4 and 1984:1 to 1997:2). For each of the Exponential, Average and Supreme tests we cannot reject the hypothesis of no break in either the 1953:2 to 1983:4 or 1984:1 to 1997:2 subsample.[16]

## 2.3 The Empirical Business Cycle Revisited

As a means of summarizing our empirical findings, we return briefly to the regime-switching models of Section 2.1. This time however, we use our estimated break date

not the frequency.

[16]One might hypothesize that the break in 1984:1 is simply due to a return to stability after the highly volatile 1970s. Our finding of no additional breaks indicates that this is not the case. This accords well with the pattern of smoothed probabilities from Model 2, as shown in the second panel of Figure 2. These probabilities trace out a one-time shift to the low-variance state in the early 1980s, rather than, for example, a pattern in which we see a high probability of the low variance state on either side of the 1970s.

Table 4: Markov-Switching Models: U.S. Real GDP Growth - Split Sample

| **AR(1) Specification - SAMPLE 1: 1953:2 to 1983:4** | | | | | | |
|---|---|---|---|---|---|---|
| | **Model 1** | | **Model 2** | | **Model 3** | |
| $\mu_1$ | 1.16 | (0.28) | 0.84 | (0.16) | 1.31 | (0.21) |
| $\mu_2$ | -0.38 | (0.61) | | | 0.37 | (0.35) |
| $\phi$ | 0.21 | (0.19) | 0.34 | (0.09) | 0.25 | (0.11) |
| $\sigma_1^2$ | 0.84 | (0.15) | 0.56 | (0.35) | 0.54 | (0.18) |
| $\sigma_2^2$ | | | 1.41 | (0.40) | 1.43 | (0.30) |
| $p_{11}$ | 0.89 | (0.07) | 0.83 | (0.22) | 0.86 | (0.09) |
| $p_{22}$ | 0.68 | (0.27) | 0.94 | (0.13) | 0.89 | (0.11) |
| LL value | -183.39 | | -184.17 | | -181.94 | |
| p-value (vs. Model 3) | 0.09 | | 0.03 | | | |
| **AR(1) Specification - SAMPLE 2: 1984:1 to 1997:2** | | | | | | |
| | **Model 1** | | **Model 2** | | **Model 3** | |
| $\mu_1$ | 0.76 | (0.07) | | | 0.76 | (0.07) |
| $\mu_2$ | -0.46 | (0.42) | | | -0.61 | (0.24) |
| $\phi$ | 0.19 | (0.15) | | | 0.15 | (0.16) |
| $\sigma_1^2$ | 0.19 | (0.04) | | | 0.19 | (0.04) |
| $\sigma_2$ | | | | | 0.09 | (0.11) |
| $p_{11}$ | 0.98 | (0.02) | | | 0.98 | (0.02) |
| $p_{22}$ | 0.67 | (0.28) | | | 0.64 | (0.28) |
| LL value | -37.56 | | | | -37.47 | |
| p-value (vs. Model 3) | 0.67 | | | | | |

Note: Standard errors appear in parentheses below coefficient estimates. Model 3 is the unrestricted model. Model 1 restricts the variance to be constant, Model 2 restricts the mean to be constant. We do not report estimates of Model 2 for the second subsample because this model converged to a linear specification for a wide range of initial values.

of 1984:1 to split the data into two subsamples and re-estimate Models 1, 2 and 3. We report the results of this exercise in Table 4. The p-values for the tests of the restrictions imposed by Model 1 and Model 2 indicate that for each subsample we cannot reject Model 1 in favor of Model 3 (using a five percent test).[17]

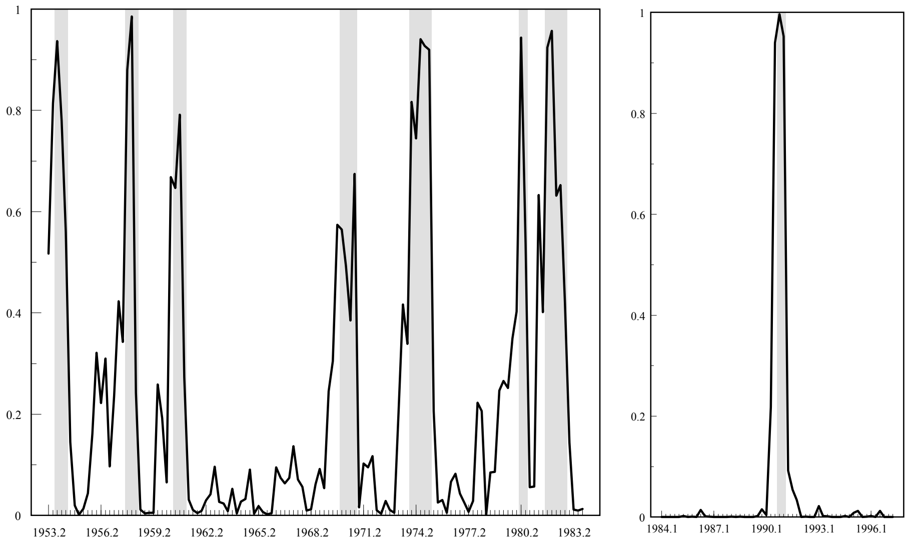
In Figure 4, we plot the smoothed probabilities of the low mean state obtained from the estimation of Model 1 over each of the subsamples. It is useful to compare this figure with the top panel of Figure 2, which plots the estimated probabilities from Model 1 over the full sample. Note that the probabilities from the full sample estimation do not pick out the early 1990s as a period of recession, but that the probabilities from the subsample estimation clearly do.

This exercise illustrates the importance of accounting for the volatility break in this type of state-space characterization of business cycle fluctuations. When the model is estimated over the full sample, the signal from the 1990-91 recession is simply too weak to register this period as a recession. However, once the relatively high-variance years of 1953 though 1983 are excluded from the estimation, the signal to noise ratio of the model is increased, and as a consequence, the model is now able to distinguish the 1990-91 recession.

The implications of the volatility break extend beyond those for econometric modeling. In particular, the break implies that policy makers and economic analysts should update their posterior distribution of quarter-to-quarter GDP growth to reflect the fact that extreme movements in output are much less likely to occur today than they were twenty or thirty years ago. For example, in the period from 1953 to 1983, approximately 30% of quarterly GDP growth rates were in excess of 1.5%, while for the period beginning in 1984, only 0.6% of observations were as large as 1.5%. On the other end of the distribution, realizations of output growth below 0% accounted for 22% of the total in the early period, but for only 6.6% for the period since 1984.

[17]We also redo this estimation using Hamilton's exact sample, 1952:1 to 1984:4. Not surprisingly, we find that even if Hamilton had estimated the more general model, he would not have rejected the constant variance specification used in his paper.

Figure 4: Probability of Low Mean: 1953:2 to 1983:4, 1984:1 to 1997:2




# 3 Sources of the Decline in Output Volatility

## 3.1 Is the Break Unique to the U.S.?

To understand more fully the source of the reduction in the variability of output fluctuations in the U.S., we begin by conducting structural break tests on the residual variance and autoregressive coefficients from the output series of the other G7 countries. A contemporaneous decline in the volatility of other countries' output would suggest a change in the frequency or magnitude of some shock which is common across countries.

The results of the residual variance break tests are reported in Table 5. For all countries, there is no break in the AR coefficients and hence these results are omitted from the table. We find a break in Great Britain's residual variance in the fourth quarter of 1987, in Canada's in the second quarter of 1991 and in Japan's in the second quarter of 1976. We find no breaks in the output processes for France, Germany or Italy.

Table 5: Residual Variance Break Tests: Other G7 Countries

| **Specification:** $x_t = \alpha + \sum_{i=1}^{p} \beta_i x_{t-i} + \epsilon_t$ | | | | | |
|---|---|---|---|---|---|
| $x_t$ = Output Growth | | | | | |
| Country | Sample | $p =$ | Date | Expo | Ave |
| Canada | 1961:1 - 1997:1 | 1 | 1991:2 | 0.0012 | 0.0115 |
| France | 1970.1 - 1997:1 | 2 | none | 0.9937 | 0.9385 |
| Germany | 1960:1- 1990:4 | 4 | none | 0.2065 | 0.1115 |
| Great Britain | 1955:1 - 1996:4 | 1 | 1987:4 | 0.0060 | 0.0087 |
| Italy | 1970:1 - 1997:1 | 3 | none | 0.2837 | 0.1995 |
| Japan | 1955:2 - 1997:1 | 3 | 1976:2 | 0.0150 | 0.0020 |

Note: We report only the p-values from the Exponential and Average tests for breaks in the residual variance. Estimated break dates are reported only when either or both of these tests indicate significance at the 5 percent level. The tests for breaks in the AR coefficients are omitted since they uniformly fail to reject the null of no break in the AR coefficients. The data for Germany ends with the German unification.

We interpret the absence of contemporaneous breaks in other countries' output series as evidence that the source of the break U.S. output volatility in 1984 is likely unique to the U.S. economy. In light of this result, we proceed by further disaggregating U.S. output into its component parts and examining these parts for breaks.

## 3.2 A Closer Look at the U.S. Data

In this section we look for breaks in disaggregate U.S. output data as a means of better understanding the decline in aggregate volatility. We examine two alternative cuts of the data. We label the first as DECOMP1, where the components of DECOMP1 are consumption, investment, government spending, exports and imports. Our second decomposition, which we refer to as DECOMP2, breaks GDP into goods, services, and structures.

For each decomposition, we fit an AR model to both the growth rate and growth contribution of each component, and following the methodology of the previous section, we test for breaks in the residual variance and the AR coefficients from this estimation.[18] Since GDP growth is essentially the sum of the growth contributions of its components, tests for breaks in the growth contributions will reveal the extent to which an individual component is responsible for the break in the variance of GDP growth. It is further necessary to test for a break in the growth rate of a particular component, however, to determine whether the break in the growth contribution is emanating from increased stability *within* that sector, or whether there has instead been a change in the share of output accounted for by that sector.[19]

The results for DECOMP1 are presented in Table 6. We do not report the results of the breaks tests for the AR coefficients because we uniformly fail to reject the null of no break in these coefficients. We report, however, the order of the lag polynomial (indicated by p) in the second column of the table. We find evidence of breaks in export and import growth in 1982:4 and 1986:2, respectively, but no evidence of breaks in their growth contributions. In addition, there is weak evidence of a break in the growth contribution of investment in 1988:1.

How should we interpret these results? The breaks in the import and export growth rates, while perhaps an important part of the story, cannot themselves explain the reduction in output volatility in the absence of associated breaks in their growth contributions. In fact, the absence of breaks in the growth contribution of any of the components in the early 1980s suggests that no one component is responsible for the break in output volatility.

We therefore shift our focus to DECOMP2, which reflects the decomposition of

[18]Throughout our analysis, we will compute growth contributions as the product of the share of nominal GDP accounted for by a particular component in period t-1 and the real growth rate of that component in period t. The BEA uses a slightly more complicated method to compute the quarterly growth contributions. We used annual data, however, to compare our method with the BEA's, and the correlation between the BEA's growth contributions and those computed using the lagged nominal weights is greater than 0.99.

[19]In the results presented in this section, we have omitted the covariance terms that would obviously be present if one where to write out the full expression for the variance of GDP growth. We do so because in the cases in which we find no breaks in the variance terms, we also find no breaks in the covariances. However, for the cases in which we find breaks in the variance terms, the covariances provide little additional information.

Table 6: Residual Variance Break Tests: DECOMP1 - 1953:2 to 1997:2

| **Specification:** $x_t = \alpha + \sum_{i=1}^{p} \beta_i x_{t-i} + \epsilon_t$ | | | | |
|---|---|---|---|---|
| $x_t$ = Growth Rates | | | | |
| Component | $p =$ | Date | Expo | Ave |
| Consumption | 1 | none | 0.47 | 0.42 |
| Investment | 1 | none | 0.09 | 0.11 |
| Government | 3 | 1960:3 | 0.09 | 0.04 |
| Exports | 1 | 1982:4 | 0.00 | 0.00 |
| Imports | 1 | 1986:2 | 0.00 | 0.02 |
| $x_t$ = Growth Contributions | | | | |
| Component | $p =$ | Date | Expo | Ave |
| Consumption | 1 | none | 0.70 | 0.65 |
| Investment | 1 | 1988:1 | 0.02 | 0.07 |
| Government | 3 | 1960:3 | 0.04 | 0.01 |
| Exports | 1 | none | 0.92 | 0.89 |
| Imports | 1 | 1968:2 | 0.02 | 0.02 |

Note: We report only the p-values from the Exponential and Average tests for breaks in the residual variance. Estimated break dates are reported only when either or both of these tests indicate significance at the 5 percent level. The tests for breaks in the AR coefficients are omitted since they uniformly fail to reject the null of no break in the AR coefficients.

Table 7: Residual Variance Break Tests: DECOMP2 - 1953:2 to 1997:2

| **Specification:** $x_t = \alpha + \sum_{i=1}^{p} \beta_i x_{t-i} + \epsilon_t$ | | | | |
|---|---|---|---|---|
| $x_t =$ Growth Rates | | | | |
| Component | $p =$ | Date | Expo | Ave |
| Goods | 1 | 1984:1 | 0.00 | 0.02 |
| Services | 1 | 1967.1 | 0.02 | 0.00 |
| Structures | 1 | none | 0.37 | 0.37 |
| $x_t =$ Growth Contributions | | | | |
| Component | $p =$ | Date | Expo | Ave |
| Goods | 1 | 1984.1 | 0.00 | 0.00 |
| Services | 1 | none | 0.13 | 0.10 |
| Structures | 1 | 1984:2 | 0.03 | 0.09 |

Note: We report only the p-values from the Exponential and Average tests for breaks in the residual variance. Estimated break dates are reported only when either or both of these tests indicate significance at the 5 percent level. The tests for breaks in the AR coefficients are omitted since they uniformly fail to reject the null of no break in the AR coefficients.

GDP expenditures by major type of product, rather than expenditure category. The results of these break tests are reported in Table 7. We find strong evidence of a break in the variance of goods and its growth contribution, and the break date corresponds to that found for aggregate output growth: 1984:1. In addition, there is no break in the volatility of services or its contribution to growth. Finally, while there is no evident break in the volatility of the structures sector, there is a break in its growth contribution. These suggest that the break in output is emanating from either the goods or structures sectors of the economy (or both). We explore each of these possibilities in turn, starting with the structures sector.

The break in the variance of the growth contribution of structures, without a corresponding break in the growth rate itself, prompts us to consider the role of the proportion of output accounted for by structures in the decline in the volatility of aggregate output. The average proportions of GDP accounted for by each of the

Table 8: Residual Variance Break Tests: Structures Experiment - 1953:2 to 1997:2

| **Specification:** $x_t = \alpha + \sum_{i=1}^{p} \beta_i x_{t-i} + \epsilon_t$ | | | | |
|---|---|---|---|---|
| | $x_t$ = Growth Rates | | | |
| | $p =$ | Date | Expo | Ave |
| GDP1 | 1 | 1984:1 | 0.02 | 0.03 |

Note: We report only the p-values from the Exponential and Average tests for breaks in the residual variance. Estimated break dates are reported only when either or both of these tests indicate significance at the 5 percent level.

components in the pre-1984 portion of our sample are 0.36, 0.52 and 0.12 for goods, services and structures, respectively. The corresponding proportions for the post-1984 period are 0.38, 0.53 and 0.09. Thus there has been a decline in the proportion of structures, and this decline has been fairly evenly distributed across the other two sectors.

Given that services is less volatile than structures, the sectoral shift away from structures and towards services may explain the reduction in output volatility. To evaluate this possibility, we conduct a simple experiment in which we hold the proportion for each sector constant at its sample wide average, thereby not allowing the ratio of structures to output to decline. A new output series (labeled GDP1) is generated under this counter factual assumption, and this series is tested for a structural break.

Table 8 shows that we obtain the same break date for our simulated data as was found for actual output. Thus while there is a reduction in the growth contribution of structures in the early 1980s, this reduction is simply not large enough to account for the magnitude of the reduction in output volatility that occurred in 1984.

We therefore turn our attention to the break in the growth contribution of goods. We make the simplifying assumption that the proportion of output accounted for by goods is a constant and couch the remainder of the analysis in terms of the growth

Table 9: Residual Variance Break Tests: Goods - 1953:2 to 1997:2

| **Specification:** $x_t = \alpha + \sum_{i=1}^{p} \beta_i x_{t-i} + \epsilon_t$ | | | | |
|---|---|---|---|---|
| $x_t$ = Growth Rates | | | | |
| Component | $p =$ | Date | Expo | Ave |
| Durables | 1 | 1985:1 | 0.01 | 0.01 |
| Nondurables | 1 | none | 0.41 | 0.34 |
| $x_t$ = Growth Contributions | | | | |
| Component | $p =$ | Date | Expo | Ave |
| Durables | 1 | 1985:1 | 0.00 | 0.00 |
| Nondurables | 1 | none | 0.19 | 0.26 |

Note: We report only the p-values from the Exponential and Average tests for breaks in the residual variance. Estimated break dates are reported only when either or both of these tests indicate significance at the 5 percent level. The tests for breaks in the AR coefficients are omitted since they uniformly fail to reject the null of no break in the AR coefficients.

rate, rather than growth contribution, of goods.[20]

The growth rate of goods can be further decomposed into the contributions from durables and nondurables growth. We test for breaks in each of these quantities and find that both the growth rate and contribution of durables break in the first quarter of 1985. We find no evidence of a break in the corresponding quantities for nondurable goods.

To assess the role of the decline in durables volatility in the reduction in aggregate volatility, we undertake an exercise similar to the one used to examine the role of structures. We generate a new durables series by holding the volatility constant at its pre-1984 average throughout the whole sample and we use this series to construct an output series, which we refer to as GDP2. Tests for parameter constancy on this

[20] This assumption allows us to avoid the problems associated with analyzing the variance of the product of two random variables (one would need to impose more structure on the problem by making distributional assumptions) and is defensible on the grounds that the average proportion of total output accounted for by goods is 0.36 in the pre-1984 period and 0.38 in the post-1984 period.

Table 10: Residual Variance Break Tests: Durables Experiment - 1953:2 to 1997:2

| **Specification:** $x_t = \alpha + \sum_{i=1}^{p} \beta_i x_{t-i} + \epsilon_t$ | | | | |
|---|---|---|---|---|
| $x_t =$ Growth Rates | | | | |
| Series | $p =$ | Date | Expo | Ave |
| GDP2 | 1 | none | 0.47 | 0.40 |

Note: We report only the p-values from the Exponential and Average tests for breaks in the residual variance. Estimated break dates are reported only when either or both of these tests indicate significance at the 5 percent level.

new series are reported in Table 10.

This table shows that by simply not allowing the variance of durables to decline in the way that it actually did, we have constructed an aggregate output series for which there is no volatility break. Thus, the magnitude of the decline in durables volatility alone is sufficient to account for the break in the volatility of aggregate output.[21]

# 4 Discussion

The previous two sections provide evidence that the magnitude of quarter-to-quarter fluctuations in real aggregate U.S. GDP growth declined in a statistically significant way in the early 1980s, and that this decline can be explained by a corresponding reduction in the volatility of durable goods production.[22] Why did output volatility decline so dramatically in the early 1980s?

In this section, we outline a set of candidate explanations for the volatility decline. In doing so, however, we emphasize that the purpose of our empirical analysis thus far

[21]This experiment is not strictly correct in that we should allow the weights to change each period as the growth rate of durables changes. Our omission of this portion of the exercise, however, disadvantages our hypothesis that the reduction in the volatility of durables alone can account for the reduction in the volatility of GDP.

[22]Strictly speaking, our results do not correspond to production, but instead to total output. However, we also conducted structural break tests on both aggregate and durables industrial production and find that both these series have breaks in 1984.

has been to characterize, rather than explain, the recent changes in output fluctuations. Nonetheless, it is useful to examine the extent to which each of these alternative explanations is compatible with the empirical facts presented in the previous sections.

## 4.1 Changes in Composition of U.S. Economy

One commonly held notion is that the increased stability is owed to a shift in the composition of output from manufacturing to services. First, the product decomposition used in this paper shows that there has been almost no change in the proportion of services relative to goods.[23] Second, even if the stability of these proportions is an artifact of the particular definition of services used in this paper, it is difficult to see why a compositional shift would lead to a decline in volatility *within* the goods sector of the economy. The break in durables volatility in the early 1980s seems to weaken the case for the compositional shift story.

## 4.2 Monetary Policy

Another potential explanation is that monetary policy has succeeded in stabilizing output fluctuations. In particular, it has been argued that in 1979 the conduct of monetary policy changed in such a way as to become a more stabilizing influence relative its pre-1979 counterpart (see for example, Clarida, Galí and Gertler (1997)). Though the timing of this explanation is appealing, it is not easily reconciled with two of the empirical facts presented in this paper. First, we find a break in the volatility of durables production, but no corresponding decline in the volatility of nondurables, services or structures. It seems likely that monetary policy ultimately affects all sectors of the economy, and thus we should see its impact in these other sectors. Second, and perhaps more importantly, even if policy is likely to first affect

[23] Filardo (1997) points out that while there has been a significant shift in the composition of the labor force towards services and away from manufacturing, there has been an offsetting increase in productivity in the manufacturing sector. Thus there has been very little change in the composition of output.

Table 11: Residual Variance Break Tests - 1953:2 to 1997:2

| **Specification:** $x_t = \alpha + \sum_{i=1}^{p} \beta_i x_{t-i} + \epsilon_t$ | | | | |
|---|---|---|---|---|
| | $x_t$ = Growth Rates | | | |
| | $p =$ | Date | Expo | Ave |
| GDB | 1 | 1984:1 | 0.00 | 0.00 |
| FSD | 1 | none | 0.24 | 0.18 |

Note: We report only the p-values from the Exponential and Average tests for breaks in the residual variance. Estimated break dates are reported only when either or both of these tests indicate significance at the 5 percent level.

an interest sensitive sector of the economy, such as durables, one might expect to see this effect in *sales* of durable goods rather than production, per se. In fact, we see no break in the volatility of sales.

## 4.3 Changing Trade Patterns

Another possibility is that changes in trade patterns are responsible for the decline in output volatility. The decomposition used in this paper would not detect such a change since our definition of output is net of imports. If we are subtracting out a volatile component of gross domestic purchases and if the proportion of output accounted for by this component rose sharply in the early 1980s, this might account for the break in volatility.

As a first pass at determining whether this is the case, we conducted structural break tests on gross domestic purchases of goods and services (denoted GDB), which is GDP with exports subtracted out and imports added back in. The first row of Table 11 shows that we find a break in this series in 1984:1. This break indicates that the U.S. economy is not simply 'exporting its business cycle', since it is domestic purchases of goods and services that has changed, not just the return to domestic factors of production.

### 4.4 Inventories

Given that the decline in durables volatility is of a magnitude sufficient to account for the aggregate volatility break, we proceed to look more closely at changes in the durables sector. We decompose the growth of durable goods output in the following way:

$$\dot{dur}_t = \dot{sal}_t(\frac{sal}{dur})_{t-1} + \Delta \dot{inv}_t(\frac{\Delta inv}{dur})_{t-1} \tag{13}$$

where $sal$ is real sales of durable goods and $\Delta inv$ is the change in real inventories. To understand the break in durables volatility, we focus on two variables. The first is $var(\dot{sal})$, and the second is $|\frac{\Delta inv}{dur}|$.[24] We look at the absolute value of $\frac{\Delta inv}{dur}$ because we are interested in determining whether inventory movements, either positive or negative, have become a smaller fraction of durables production.

The results of break tests for these variables are reported in Table 12. The top panel shows that there is no evidence of a break in $var(\dot{sal})$. On the other hand, we see strong evidence of a break in $|\frac{\Delta inv}{dur}|$ in the third quarter of 1984, a date which corresponds closely to that found for aggregate output.

One potential explanation for the declining share of inventories is the introduction of inventory management techniques such as just-in-time. These techniques began to be widely used in the U.S. in the early to mid-1980s, mainly in response to increased global trade and the high inventory carrying costs brought on by the exceptionally high interest rates of the early 1980s.[25] The timing of the estimated breaks in output and durables volatility documented in this paper corresponds with the introduction of these methods in U.S. manufacturing. It is interesting to note that Japanese firms began to use just-in-time methods earlier than did U.S. firms, and that we find a break in Japanese output in the mid-1970s.[26]

[24] The other variable that could affect durables is the growth rate of the change in inventories. However, we do not analyze this variable because, in addition to the computational difficulties associated with this quantity, it lacks an obvious economic interpretation.

[25] We are distinguishing the effects of the general conduct of monetary policy from the effects of the policy induced high real interest rates of the early 1980s. We discuss the former above, and the latter here.

[26] One might also expect to see declines in the inventory-to-sales ratio since the early 1980s. West

Table 12: Break Tests: Durables - 1953:2 to 1997:2

| **Final Sales of Durables: Growth Rate** | | | |
|---|---|---|---|
| Specification: $x_t = \alpha + \sum_{i=1}^{p} \beta_i x_{t-i} + \epsilon_t$ | | | |
| $p =$ | Date | Expo | Ave |
| 1 | none | 0.83 | 0.86 |
| **Absolute Value of ($\Delta I/Dur$)** | | | |
| | Date | Expo | Ave |
| | 1984:3 | 0.00 | 0.00 |

Note: We report only the p-values from the Exponential and Average tests for breaks in the residual variance. Estimated break dates are reported only when either or both of these tests indicate significance at the 5 percent level. The tests for breaks in the AR coefficients are omitted since they uniformly fail to reject the null of no break in the AR coefficients.

An alternative possibility is that changes have occurred in the industry composition of inventories and that the decline in the proportion of output accounted for by inventories may reflect a shift toward less inventory intensive industries. Similarly, there may have been changes in the composition of goods by stage of processing that have caused the fraction of inventories to shrink.

Since inventories traditionally account for a large fraction of the variability of aggregate output, the declining share of inventories could have substantial effects on the volatility of output fluctuations.[27] As a final exercise, we subtract inventories from the GDB series used above (inventories in this case include both imported and domestically produced inventories) from GDB, and obtain domestic final sales of goods and services (denoted FSD). In the bottom row of Table 11 we see that once

(1992) notes that there is a decline in the aggregate inventory-to-sales ratio in Japan beginning in the early to mid-1970s. He also notes that there is no evidence of similar decline in the U.S.. His data, however, end in the late 1980s. A plot of the inventory-to-sales ratio (not shown) suggests that at least for U.S. durables manufacturing, there has been a downward trend in the inventories-to-sales ratio in the period since the mid-1980s.

[27]The link between inventories and output stability has been explored by other authors, for example, Morgan (1991), Allen (1995), Filardo (1995), and Ramey and West (1997).

we subtract purchases of inventories from total purchases, we have eliminated the volatility break.

Further research is needed to sort through the evidence that inventories are an important factor in producing the recent stability. In particular, it would be useful to determine which industries make the most use of just-in-time techniques and to assess their contribution to the decline in volatility. It would also be interesting to examine the extent to which these methods have been used in the other G7 countries and to relate this to the existence or lack of breaks in the output processes for these countries.

# 5 Conclusions

This paper documents a break in the volatility of U.S. output in the early 1980s. This break has important implications for widely used theoretical and empirical techniques, examples of which include model calibration and the estimation of state-space models of business cycle fluctuations. In addition, since the break implies that we are now much less likely to see extreme movements in GDP growth, it affects the interpretation policy makers place on particular growth rate realizations.

In order to provide a comprehensive characterization of the break in output volatility, we examine international as well as disaggregate U.S. output data for similar breaks. Our findings suggest that no other G7 country shared a contemporaneous break in output. We also find that the break in U.S. output emanates from a break in the volatility of durable goods production, and that the timing of these breaks corresponds to a reduction in the proportion of durables accounted for by inventories. A precise characterization of the changes that led to the decline in the proportion of inventories is the subject of further research.

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