

OUT PUT FLUCTUATIONS IN THE UNITED STATES: WHAT HAS
CHANGED SINCE THE EARLY 1980s?

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

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Abstract

In this paper, 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. Further, the reduction in durables volatility corresponds to a decline in the share of durable goods accounted for by inventories. We find no evidence of increased stability in the nondurables, services or structures sectors of the economy. Our evidence is compatible with a scenario in which changes in inventory management techniques in the durable goods sector have reduced the variability of aggregate output.

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

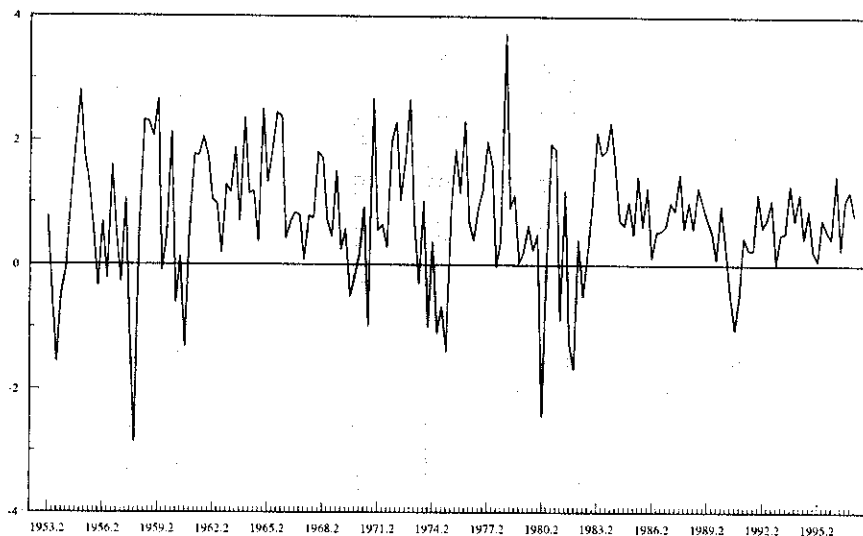
The business press is currently sprinkled with references to the ‘death’ or ‘taming’ of the business cycle in the United States. While such claims are undoubtedly premature, they are in part rooted in the apparent reduction in the volatility of U.S. output fluctuations over the period beginning in the early 1980s. Figure 1 plots the growth of U.S. GDP over the period 1953:2 to 1997:2; the variance of output fluctuations over the period ending in 1983 is almost five 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. This break 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 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 the estimation of linear models for output growth over periods that span the break is 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. Finally, the reduction in the variance of output fluctuations should alter the interpretation that 1 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.

As a means of understanding the dramatic reduction in output volatility in the early 1980s, we decompose output growth into its component parts and provide evidence that the break emanates from a reduction in the volatility of durable goods

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



production. We further show that the timing of this reduction coincides with a break in the proportion of durables accounted for by inventories. The evidence presented here is compatible with a scenario in which changes in inventory management techniques have reduced the variability of aggregate output. Such a scenario suggests that it is unlikely that output growth will return to its pre-1984 volatility levels.

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 a set of candidate explanations for the volatility decline and propose that the stylized facts set forth in this paper accord best with a scenario in which changes in inventory management techniques have served to stabilize output fluctuations. 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, b, 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 macroeconomic research. Such documentation is valuable both because it leads to a collection of macroeconomic 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 with an illustrative exercise in which we show 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 return to our illustrative exercise and show that by using the estimated breakdate to split the sample we can recapture the strength of the business cycle signal even when we allow both the mean and the variance of the output process to switch between states.

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., and 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 growth¹. Conditional on the value of S_t , the expected value for the rate of growth of GDP, denoted \dot{y}_t , is:

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

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

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

We can rewrite Equation 1 as:

$$\dot{y}_t = \mu_{S_t=i} + u_t \quad (3)$$

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 the variance of the AR(p) model to switch between

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

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

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

where i can assume two alternative values: 0 or 1². These two values are commonly interpreted as indicating periods of recession and expansion. Therefore, $\mu_{S_t=0}$ is the expected value of the rate of growth of GDP during recessions and $\mu_{S_t=1}$ 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.

Table 1 reports the results of this exercise for an AR(1) specification for output growth³. 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)⁴.

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, 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

²Equation 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 $t-1 \dots t-p$. We therefore have 2^{p+1} states of the economy.

³We 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).

⁴Though 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	
μ_1	-1.13	(0.55)	0.71	(0.09)	0.65	(0.11)
μ_2	0.90	(0.13)			0.81	(0.14)
ϕ	0.35	(0.09)	0.35	(0.07)	0.33	(0.07)
σ_1	0.69	(0.11)	0.24	(0.05)	0.23	(0.05)
σ_2			1.20	(0.16)	1.19	(0.15)
p_{11}	0.96	(0.03)	0.99	(0.02)	0.99	(0.01)
p_{22}	0.36	(0.30)	0.99	(0.01)	0.99	(0.01)
LL value	-245.57		-234.45		-234.10	
p-value (vs. Model 3)	0.00		0.40			
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.

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 make 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 characterization⁵.

Up to this point we have imposed the switching specification on the data. It is important to test explicitly 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 Hansen (1992, 1994) approximation.

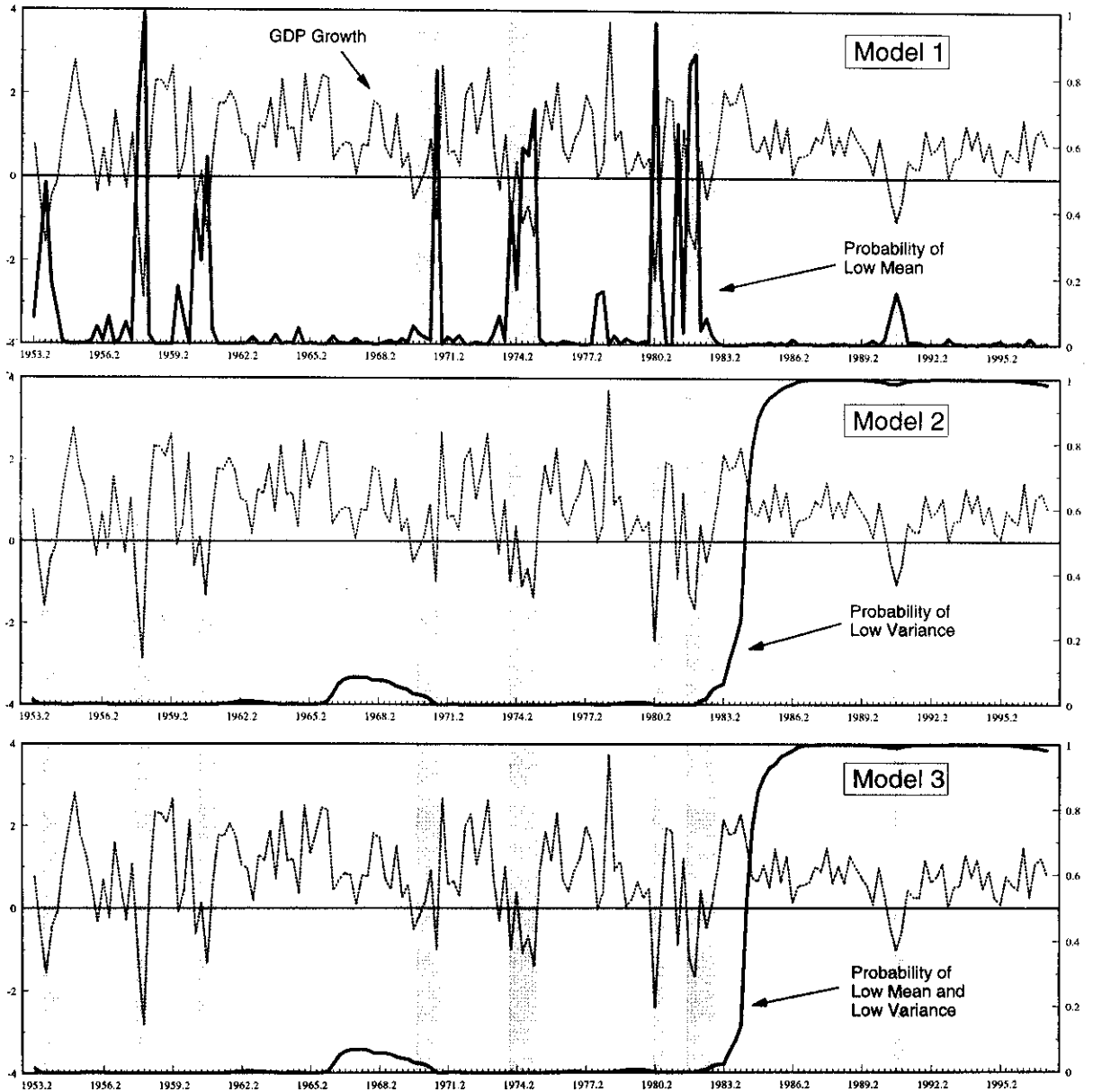
To do this, we define $\alpha = (\mu_{S_t=1} - \mu_{S_t=0})$, $\beta = (\sigma_{S_t=1} - \sigma_{S_t=0})$, $\mu = \mu_{S_t=0}$ and $\sigma = \sigma_{S_t=0}$, and rewriting Equation 4, we obtain:

$$\dot{y}_t = (\mu + \alpha S_t) + \sum \phi_j (\dot{y}_{t-j} - \alpha S_{t-j}) + \epsilon_t, \quad \epsilon_t \sim N(0, \sigma + \beta S_t) \quad (5)$$

The test requires one to compute the constrained estimates of the likelihood function over a grid of possible values for the set of parameters, Θ , that under the null hypothesis of the linear model do not converge to any fixed population parameters.

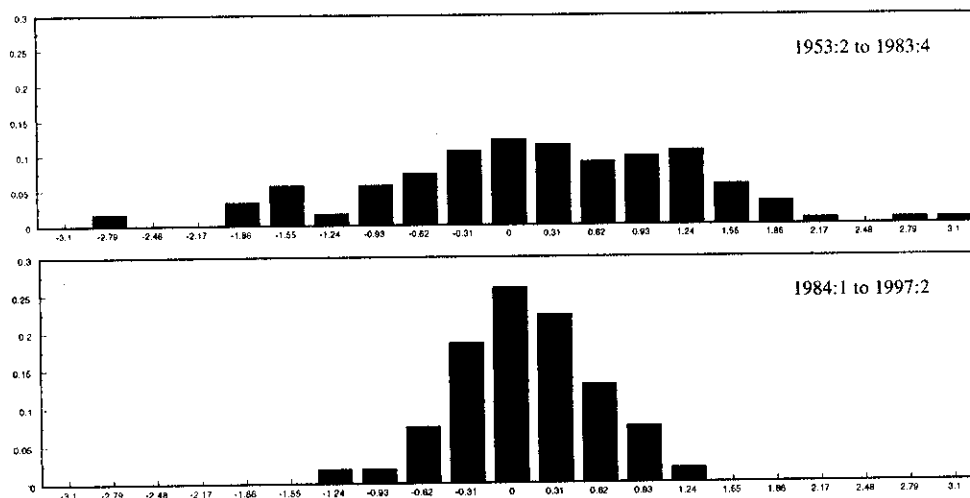
⁵Note that when we allow both the mean and the variance to switch between states (Model 3), we find that the low mean and low variance states occur together. Ramey and Ramey (1996) use a panel data set across countries to show that higher volatility in the rate of growth of GDP is associated with slower average rate of growth. Thus our time series finding does not accord well with this cross sectional result.

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 to 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



In our case, $\Theta = (\alpha, \beta, p_{11}, p_{22})$. We define the grid of values for the elements of Θ in the following way: $\alpha = \{0.05 \text{ to } 0.20 \text{ in intervals of } 0.01\}$, $\beta = \{0.5 \text{ to } 2.0 \text{ in intervals of } 0.1\}$, $p_{11} = \{0.899 \text{ to } 0.999 \text{ in intervals of } 0.002\}$ and $p_{22} = \{0.899 \text{ to } 0.999 \text{ in intervals of } 0.002\}$. This grid implies that the space for Θ is partitioned into 5625 points.

The results from this test allow us to reject the null of the linear model in favor of the switching specification (p-value = 0.02)⁶. The intuition behind this result is illustrated in Figure 3. This figure shows the two histograms 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 four times larger than the variance for the period from 1984 to 1997.

⁶We use a bandwidth size of five to account for serial correlation as suggested Hansen (1994).

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 variance in the early 1980s than by a switching regime. In this section, we estimate breakdates for both the residual variance and the autoregressive components of an AR specification for GDP growth and provide a measure of its statistical significance of our estimated dates.

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:

$$\hat{y}_t = \mu + \phi \hat{y}_{t-1} + \epsilon_t. \quad (6)$$

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

$$\sqrt{\frac{\pi}{2}}|\hat{\epsilon}_t| = \alpha + \mu_t \quad (7)$$

where α is the estimator of the standard deviation⁷.

In order to find the break, we jointly estimate the following system by GMM:

$$\hat{y}_t = \mu + \phi \hat{y}_{t-1} + \epsilon_t \quad (8)$$

$$\sqrt{\frac{\pi}{2}}|\hat{\epsilon}_t| = \alpha_1 D_{1t} + \alpha_2 D_{2t} + \mu_t \quad (9)$$

where

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

⁷In the absence of the normality assumption, $|\hat{\epsilon}_t|$ in Equation 9 can be interpreted as an estimator of the standard deviation.

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

and T is the estimated break point, and α_1 and α_2 are the corresponding estimators of the standard deviation. The list of instruments is as follows: a constant, y_{t-1} , D_{1t} , and D_{2t} ⁸.

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 α_1 and α_2 do not have standard asymptotic properties.

Andrew (1992) and Andrew and Ploberger (1994) develop tests for cases such as this one, 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 defined as the Wald, LM or LR statistic of the hypothesis that $\alpha_1 = \alpha_2$, for each possible value of T . The only information known is that T lies in a range T_1, T_2 ⁹. Andrew (1992) shows the asymptotic properties of the statistic:

$$\sup_{T_1 \leq T \leq T_2} F_n = \sup F_n(T) \quad (10)$$

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

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

$$\exp F_n = \ln(1/(T_2 - T_1 + 1)) * \sum_{T=T_1}^{T_2} \exp(1/2 * F_n(T)) \quad (11)$$

⁸The results for $\hat{\epsilon}_t^2$, the estimator of the variance, are very similar to those reported below.

⁹Following Andrew and Andrew and Ploberger, we set $T_1 = .15 * n$ and $T_2 = .85 * n$.

and

$$\text{ave } F_n = (1/(T_2 - T_1 + 1)) * \sum_{T=T_1}^{T_2} F_n(T) \quad (12)$$

and prove their optimality for the case in which a nuisance parameter is present under only the alternative hypothesis. The p-values associated with these statistics are computed using the approximation suggested by Hansen (1997)¹⁰.

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.

We now consider the possibility that the break in the residual variance is a result of the break in the AR coefficients. In particular we estimate:

$$\dot{y}_t = \mu_1 D_1 + \mu_2 D_2 + \phi_1 \dot{y}_{t-1} D_1 + \phi_2 \dot{y}_{t-1} D_2 + \epsilon_t \quad (13)$$

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 coefficient separately¹¹. 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 thus have strong evidence that the break in the variance in the first quarter of 1984 is not due to a change in the AR

¹⁰The 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 under the Exponential.

¹¹For simplicity we use a LM test rather than a Wald. Andrew and Ploberger (1994) prove that these two tests are equivalent.

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

Specification: $\dot{y}_t = \mu + \phi \dot{y}_{t-1} + \epsilon_t$				
$\epsilon_t \sim N(0, \sigma_t)$, where $\sigma_t = \sigma_1$ if $t \leq T$, and $\sigma_t = \sigma_2$ if $t > T$				
Residual Variance				
Null	Sup	Exp	Ave	
$\sigma_1 = \sigma_2$	15.43 (0.00)	5.12 (0.00)	4.96 (0.00)	
Estimated break date: 1984:1				
AR Coefficients				
Null	Sup	Exp	Ave	Chow
$\mu_1 = \mu_2, \phi_1 = \phi_2$	2.77 (0.99)	0.64 (0.85)	1.20 (0.75)	0.32 (0.85)
Estimated break date: none				
$\mu_1 = \mu_2$	2.09 (0.93)	0.40 (0.61)	0.72 (0.53)	
Estimated break date: none				
$\phi_1 = \phi_2$	2.73 (0.82)	0.46 (0.55)	0.81 (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	$\hat{\mu}$	$\hat{\phi}$	$\hat{\sigma}$
1953:2 to 1997:2	0.50 (0.00)	0.34 (0.00)	0.0096
1953:2 to 1983:4	0.53 (0.00)	0.33 (0.00)	0.0107
1984:1 to 1997:2	0.40 (0.00)	0.41 (0.00)	0.0048

Note: p-values appear in parentheses below coefficient estimates.

components of the model¹².

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 standard deviation 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 subsamples¹³.

¹²This result suggests that there has been a change in the amplitude of output fluctuations, but not their frequency.

¹³One 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	
μ_1	1.16	(0.28)	0.84	(0.16)	1.31	(0.21)
μ_2	-0.38	(0.61)			0.37	(0.35)
ϕ	0.21	(0.19)	0.34	(0.09)	0.25	(0.11)
σ_1	0.84	(0.15)	0.56	(0.35)	0.54	(0.18)
σ_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	-1.83.39		-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	
μ_1	0.76	(0.07)			0.76	(0.07)
μ_2	-0.46	(0.42)			-0.61	(0.24)
ϕ	0.19	(0.15)			0.15	(0.16)
σ_1	0.19	(0.04)			0.19	(0.04)
σ_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.

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 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. For each subsample, we find that we cannot reject Model 1 in favor of Model 3¹⁴.

In Figure 4, we plot the smoothed probabilities obtained from estimating Model 1 separately for each subsample. It is illustrative to compare this figure with the probabilities from the full sample estimation of Model 1, as shown in the top panel of Figure 2. Note that when we use the full sample, the signal from the 1990-91 recession is too weak to register this period as a recession. However, when we omit the high-variance years of 1953 through 1983 from the estimation, we obtain the result that the estimated probability of being in the low-mean state is close to one for the 1990-91 recession.

3 Sources of the Decline in Output Volatility

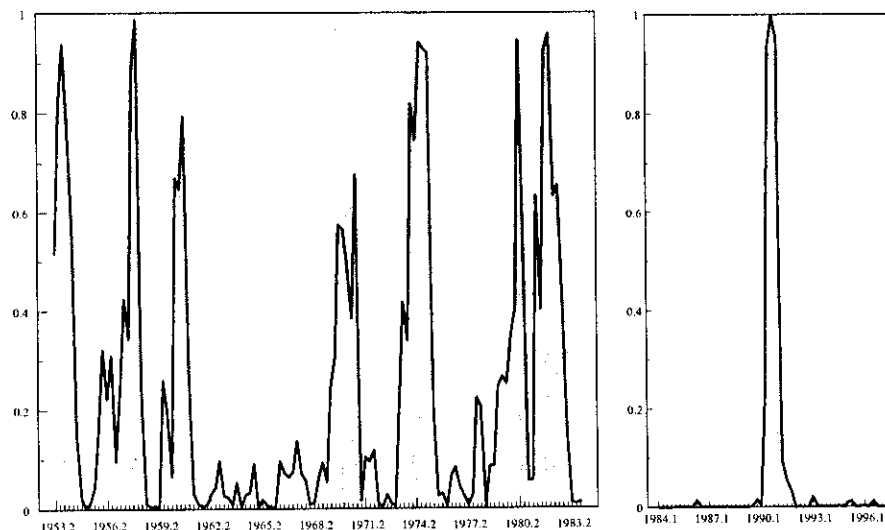
3.1 Is the Break Unique to the U.S.?

To understand more fully the source of the reduction in 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 these tests are reported in Table 5. For all countries, there is no break in the AR coefficients. 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

¹⁴We 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



in the second quarter of 1976. We find no breaks in the output processes for France, Germany or Italy.

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

Table 5: Residual Variance Break Tests: Other G7 Countries

Specification: $x_t = \alpha + \sum_{i=1}^p \beta_i x_{t-i} + \epsilon_t$					
$x_t = \text{Output Growth}$					
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.

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 estimation¹⁵. 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 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 sector¹⁶.

¹⁵Throughout 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.

¹⁶We have omitted the covariance terms from the results presented in this section. We do so because in the cases in which we find no breaks in the variance terms, we also find no breaks in the

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. In the second column of the table however, we report the order of the lag polynomial (indicated by p). We find evidence of breaks in export and import growth in 1982:4 and 1986:2, respectively, and 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 now focus on the components of DECOMP2. 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 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

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.

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.

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 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 counterfactual assumption, and this series is tested for a structural break.

The results of this experiment are reported in Table 8; we find there is still a structural break in output (as measured by GDP1) volatility in the first quarter of 1984. 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 rate, rather than growth contribution, of goods¹⁷.

¹⁷This assumption allows us to avoid the problems associated with analyzing the variance of the

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.

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 new series are reported in Table 10.

This table shows that by simply not allowing the variance of durables to decline in

product of two random variables (one needs 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.

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 output¹⁸.

4 Discussion

The empirical work in the previous two sections provides evidence of a change in the process characterizing output fluctuations in the U.S. over the last two decades. In particular, while we estimate no break in the coefficients of an AR(1) specification for GDP growth, we find strong evidence of a structural break in the residual variance in the first quarter of 1984. No other G7 country appears to have experienced a contemporaneous reduction in output volatility. In addition, we find that only the goods sector, and more precisely, the durable goods sector, experienced a corresponding reduction in volatility.

In this section, we show that the reduction in the volatility of durables production occurs at roughly the same time as a break in the proportion of durables out-

¹⁸This 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, disadvantages our hypothesis that the reduction in the volatility of durables alone can account for the reduction in the volatility of GDP.

put accounted for by inventories, a finding which suggests that changes in inventory management techniques in the early 1980s played an important role in the volatility decline. In addition, we discuss that there to be little support for a number of alternative possible explanations, including monetary policy, changing trade patterns or shifts in the sectoral composition of output. We will briefly discuss each of these candidate explanations.

4.1 Inventories

Given that the decline in goods volatility corresponds (within one year) to a decline in durables volatility, we undertake one final decomposition. In particular, we can write:

$$dur_t = sal_t \left(\frac{sal}{dur} \right)_{t-1} + \Delta inv_t \left(\frac{\Delta inv}{dur} \right)_{t-1} \quad (14)$$

where sal is real sales of durable goods and Δinv is the change in real inventories. To understand the break in durables volatility, we focus on two variables. The first is $var(sal)$, and the second is $|\frac{\Delta inv}{dur}|$ ¹⁹. 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 11. The top panel shows that there is no evidence of a break in $var(sal)$. On the other hand, there is strong evidence of a break in $|\frac{\Delta inv}{dur}|$ in the third quarter of 1984, a date which corresponds very closely to that found for aggregate output.

Figures 5 and 6 illustrate the changes which occurred in the durable goods sector in the early 1980s. Figure 5 plots real durable goods production and sales over the period 1953:2 to 1997:2. Note that while the variance of sales is fairly constant over the full sample, the variance of production declines rather dramatically in the early

¹⁹The 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.

Table 11: 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.

1980s²⁰. Figure 6 is a plot of the variable $|\frac{\Delta inv}{dur}|$ over our sample. There is a clear reduction in the mean of this proportion.

Figure 7 plots the inventory-to-sales ratio for durables and nondurables manufacturing. There is a clear change from a positive to a negative trend in the ratio for durables in the early 1980s. No such change in trend is apparent in the nondurables sector. The change in trend, as well as the decline in the share of durables output accounted for by inventories (shown in Figure 6) is consistent with the hypothesis that firms are holding less inventories. Since inventories traditionally account for a large fraction of the variability of aggregate output, such a phenomenon could have substantial effects on the volatility of output fluctuations.

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). Morgan (1991) and Filardo (1995) each discuss the implications of improved inventory control methods, generally referred to as *just-in-time* methods,

²⁰The point estimate of the correlation between the two series, however, is actually lower in the second period (0.76 for the pre-1984 period and 0.44 for the post-1984 period).

Figure 5: Growth of Production and Sales of Durables: 1953:2 to 1997:2

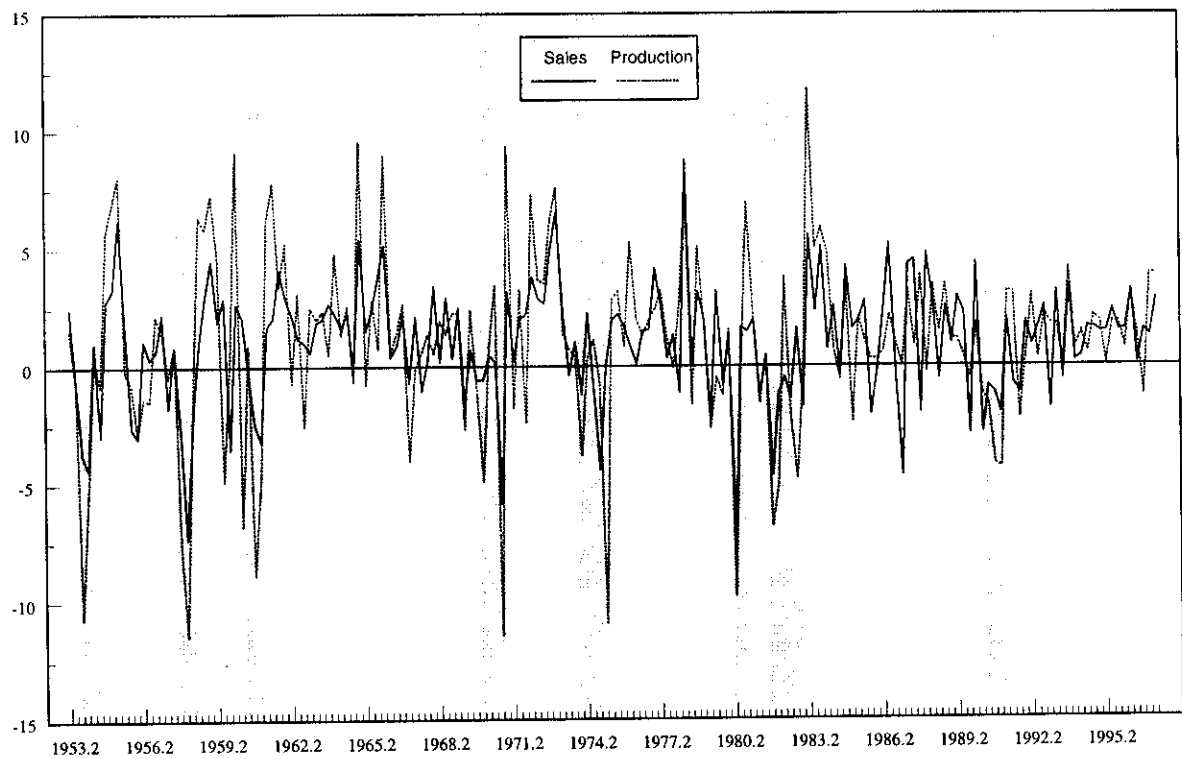


Figure 6: Absolute Value of $\Delta Inv/Dur$: 1953:2 to 1997:2

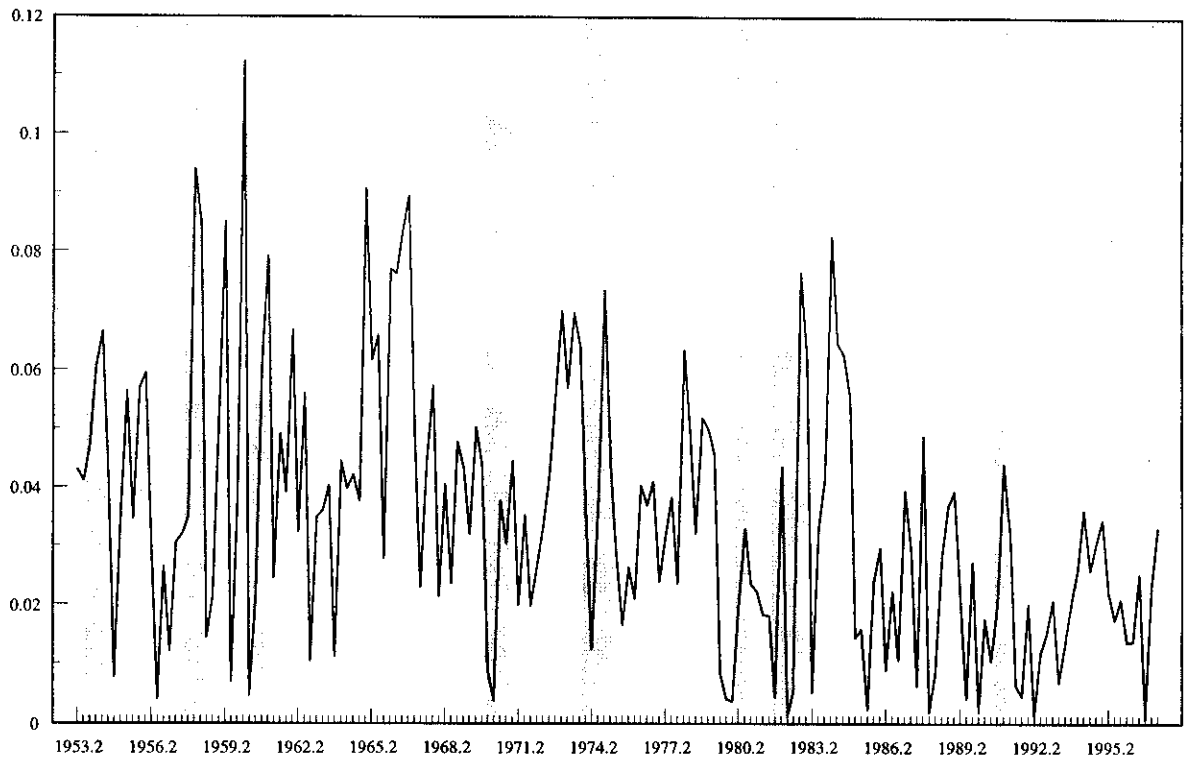
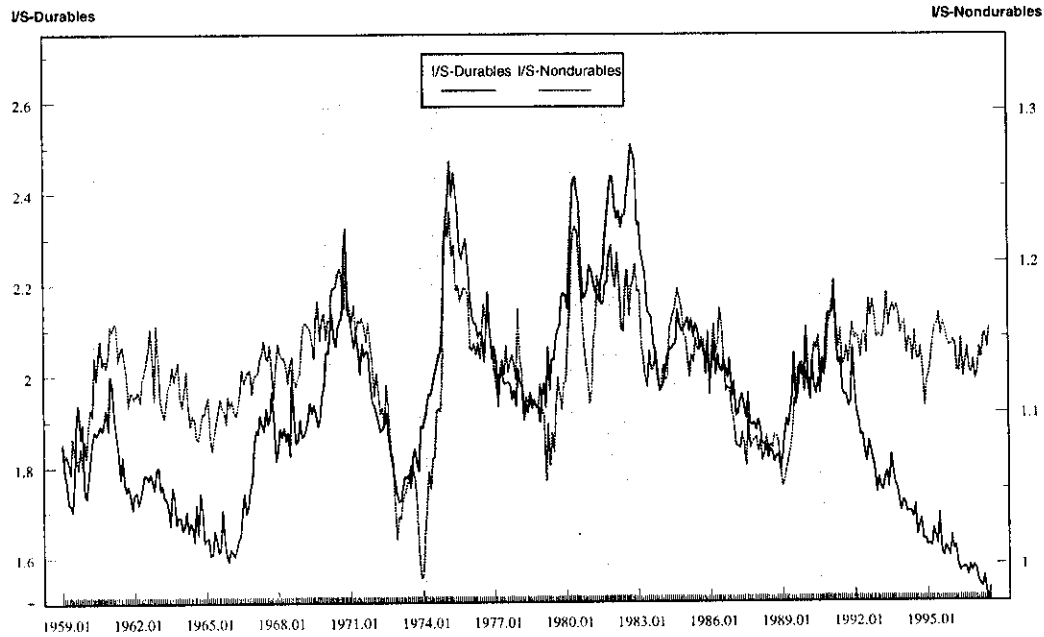


Figure 7: Inventory-to-Sales Ratios: Durable and Nondurable Goods - 1959 to 1997



for output stability. They also discuss that 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.

The timing of the introduction of these methods in U.S. manufacturing corresponds with the break in output and durables volatility documented in this paper. This literature also notes that Japanese firms began to use just-in-time methods earlier than did U.S. firms. Interestingly, we find a break in Japanese output in the mid-1970s²¹.

Further research is needed to sort through the evidence that inventories are an

²¹West (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. As noted above, Figure 7 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.

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.

4.2 Alternative Explanations

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 goods²². 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 composition shift would lead to a break in the volatility of goods production. Any competing theory would need to explain this stylized fact.

Another potential explanation is that monetary policy has succeeded in stabilizing output fluctuations. This explanation 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 affects all sectors of the economy, and thus we should see its impact in these other sectors. Second, even if policy is likely to first affect 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.

A final possibility is that changes in trade patterns in the durable goods industries are responsible for the decline in output volatility. The decomposition used in this

²²Filardo (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 12: 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.

paper would not detect such a change since our definition of durables is net of imports of durables. If we are subtracting out a volatile component of gross domestic purchases of durables and if the proportion of durables accounted for by this component rose sharply in the early 1980s, this might account for the break in output 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 (GDB)²³. This variable is essentially GDP with exports subtracted out and imports added back in. The first row of Table 12 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. Finally, we subtract out inventories (this measure of inventories includes both imported and domestically produced inventories) from GDB in order to obtain final sales of domestic goods and services (FSD) and we test this series for a break. In the bottom row of Table 12 we report that once we subtract purchases of inventories from total purchases, we have no break in volatility. This final decomposition provides additional support for the inventories hypothesis.

²³It would be preferable to conduct the following analysis on durables alone, rather than on aggregate output, but we were unable to obtain the appropriate data.

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, the volatility decline constitutes a stylized fact of the U.S. economy, and as such, it should be explainable by any complete model of the macro economy.

In order to provide a more 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. While our results are suggestive of a role for improved inventory management techniques, further research will explore this question more thoroughly.

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