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Breaking trends and the money-output correlation

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BREAKING TRENDS AND THE MONEY–OUTPUT CORRELATION

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Abstract—This paper examines the impact on the money–output correlation of a univariate specification that allows time series to be characterized as stationary around a broken trend function. Though pretesting suggests that U.S. real output (industrial production) can be described as broken-trend stationary, this result has only limited impact on the money–output correlation. Before 1985 there is a strong Granger causal relationship between money and broken-detrended output (but not first-differenced output), even when different short-term interest rates are used as regressors. However, after 1985 this relationship weakens significantly, whether or not one determines that output has a unit root.

I. Introduction

Does monetary policy affect the real economy? Friedman and Schwartz (1963) were the first to document extensively that ''the impact of changes in the stock of money on the rest of the economy appears to have been highly stable.'' More formal statistical evidence from distributed lag regressions (Anderson and Jordan (1968)) and Sims' (1972) work using vector autoregressions helped to forge a pre-1980s consensus that movements in money did convey useful information about future income movements. However, during the 1980s this empirical regularity has been seriously challenged on several levels. First, if one first-differences the data to achieve stationarity (following Nelson and Plosser's (1982) advice), the predictive power of money is reduced substantially (Christiano and Ljungqvist (1988)). Second, the importance of money basically disappears when the system is expanded to include a short-term interest rate variable (Sims (1980) and Litterman and Weiss (1985)). Third, the relationship is sensitive to the time period chosen (Eichenbaum and Singleton (1986)).

Stock and Watson (1989b) attempt to address these empirical ''puzzles'' and, in the process, rescue the money–output correlation. Their key result is that proper, systematic handling of the trends in the data (using a series of pretests) yields a stable relationship between money and real activity, even in the presence of a short-term interest rate, that endures up to the end of 1985. However, Friedman and Kuttner (1993) argue that Stock and Watson may have created more puzzles than they solved. Friedman and Kuttner run sensitivity checks that reveal two potential problems: first, the finding that fluctuations in money help explain subsequent output movements disappears when the commercial paper rate is used instead of the Treasury bill rate; second, the nonneutrality result again vanishes when the sample period is extended past December 1985.

This paper follows in the spirit of Stock and Watson in the sense that careful attention is first paid to the univariate trend properties of the data before examining the money–output correlation. The contribution of this paper is to expand the list of pretests. In particular, we test whether individual series can be characterized as stationary processes around a linear trend function that has a single change in its slope. Perron (1989, 1994) has argued that most macroeconomic time series are best characterized as stationary fluctuations around such a broken deterministic trend function. For example, a decline in the slope of the linear trend of (log) postwar U.S. real output would correspond to a slowdown in the growth path of the series.

The impact of this alternative characterization on the study of the statistical relationship between money and output could be significant. Specifically, if the difference-stationary characterization for real output used by many authors is rejected in favor of this broken-trend stationary alternative, then the output regressions used in their analyses are misspecified.

II. Trend and Integration Properties of the Data

The first step in the statistical analysis of the relationship between money and economic activity is to determine the univariate characteristics of the data through a series of pretests. The standard approach is first to test each variable for the presence of a unit root against the alternative hypothesis that the series is stationary around a linear trend. Table 1 reports augmented Dickey–Fuller statistics for log levels and first-differences of industrial production, nominal *M*1, wholesale prices, and the levels of the 3-month Treasury bill yield (*TB*3), the 6-month commercial paper rate (*CP*6), and the spread between them. Following Stock and Watson, we also test for the presence of a constant and a time trend in the differenced series. Mirroring their findings, output is characterized as having a single unit root with drift; the money supply has a single unit root with drift and a trend, inflation has a unit root;[1] both interest rates have a single unit root without drift; and the interest rate spread is stationary. A concern when combinations of these variables are analyzed together is that the series may have stochastic trends in common, interacting in such a way that their residuals are no longer integrated. However, cointegration tests will be postponed at this stage so that an alternative test for stationarity can be introduced and implemented. If any of the level series is found to be stationary, then there exists cointegration between them with a trivial cointegrating vector.

III. Broken Trends

An alternative to characterizing real output as containing a unit autoregressive root is that the log series may be stationary around a deterministic trend function that has one slope in the first part of the sample and a different slope in the second.[2] Though this model allows for only a single break in the time trend of a series at a single point in time, it can be thought of as proxying for possibly many events that

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[1] The statistics in table 1 are based on regression in which the lag length is chosen by the data-dependent t-sig method (see Ng and Perron (1995) and appendix A). Stock and Watson's findings also suggest the possibility that inflation is $I(1)$; they provide evidence that by controlling for the substantial heteroskedasticity exhibited by WPI during the first oil shock, inflation appears not to be integrated. Even if it were, the usual asymptotic distribution for the F-tests in section IV would still apply (see Sims et al. (1990)).

[2] The seminal papers are by Perron (1989) and Rappoport and Reichlin (1989).

TABLE 1.—TESTS FOR INTEGRATION AND TIME TRENDS, 1959–1994

Series	Augmented Dickey–Fuller Statistics		t-Statistics for Regression of Differences on:	
	Levels	Differences	Time[a]	Constant[b]
log IP	−2.90	−6.68[c]	−0.99	3.73[c]
log $M1$	−2.73	−4.35[c]	2.04[d]	4.13[c]
log prices	−1.72	−3.05	−0.09	2.79[c]
T-bill rate TB	−2.21	−7.76[c]	−0.69	0.23
Commercial paper rate CP	−2.68	−5.86[c]	−0.57	0.23
CP–TB spread	−5.30[c]	—	—	—

Notes: For Dickey–Fuller statistics, lag length selected using t-sig criterion starting with $k_{max} = 10$ (Ng and Perron (1995)); critical values, tabulated in Fuller (1976), are 10% = −3.13, 5% = −3.42, 1% = −3.96.
[a] t-statistic on the coefficient on the trend term in a regression of the first-differenced series on a constant, time trend, and six own lags.
[b] t-statistic for the constant in a similar regression without a time trend.
[c] Significant at the 1% level.
[d] Significant at the 5% level.

TABLE 2.—INTEGRATION TEST FOR OUTPUT ALLOWING FOR BREAKING-TREND FUNCTION, 1959–1994

A. Step 1: Detrending Regression[a]

$y_t = \mu + \beta t + \gamma(t - T_B)1(t > T_B) + \tilde{y}_t$

Break Date	Constant	Trend	Dummy
June 1968	3.54	0.0050	−0.0029
	(531.01)	(61.48)	(−32.91)

B. Step 2: Integration Test for Residuals

$\tilde{y}_t = \alpha\tilde{y}_{t-1} + \Sigma_{i=0}^{k} \beta_i \Delta\tilde{y}_{t-i} + \epsilon_t$

α	Last β	Lag Length
0.949	0.091	8
(−4.51)	(1.94)	

Notes: t-statistics are in parentheses. Under Step 2, t-statistic for testing $\alpha = 1$ is reported. Break date is selected by minimizing the t-statistic on γ in Step 1. Lag length selected using t-sig criterion starting with $k_{max} = 10$. Finite-sample critical values for t-statistic on α, tabulated in Perron (1994), are 10% = −3.83, 5% = −4.22, 2.5% = −4.50, 1% = −4.77.
[a] $1(\cdot)$ is the indicator function.

alter and determine the growth path of output over several decades. A broken-trend stationary characterization of a time series provides the econometrician with a simple modeling tool that takes into account these important events, while preserving the property that output exhibits mean reversion over business-cycle horizons.

A definitive answer to the question of whether U.S. real output can be characterized as being stationary around a broken trend is not provided by the current literature; several authors using different methods find different results. Christiano (1992) and Banerjee et al. (1992) fail to find evidence of broken-trend stationarity in U.S. output, while Perron (1994) finds the opposite. It should be noted that these authors analyze real gross national product (GNP) at quarterly or annual frequencies over long spans. (For example, Perron uses the Nelson and Plosser (1982) data set in which the real GNP series begins in 1909.) This paper follows the recent literature on the money–output correlation and focuses on the monthly industrial production as the measure of output after 1959.

Among the several methods that test whether or not a series can be characterized as broken-trend stationary when the break date is unknown a priori, we use one particularly simple procedure proposed by Perron (1994). This procedure allows for a single change in the slope of the trend function that occurs at one point in time, with both segments of the trend function joined at the break point T_B.

This strategy can be broken down into two steps. First, ''detrend'' the time series, x_t by running the following ordinary least-squares (OLS) regression:

$$x_t = \mu + \beta t + \gamma(t - T_B)1(t > T_B) + \tilde{x}_t \tag{1}$$

where $1(\cdot)$ is the indicator function and T_B is the date of the change in the slope of the trend function. Estimate this regression over all possible break dates and select the single break date that minimizes the t-statistic on γ, the parameter associated with the change in slope.[3] Second, test the residuals of the regression for the presence of a unit root using an augmented Dickey–Fuller test, which involves the t-statistic for $\alpha = 1$ in the regression,

$$\tilde{x}_t = \alpha\tilde{x}_{t-1} + \sum_{i=1}^{k} C_i \Delta\tilde{x}_{t-i} + \epsilon_t. \tag{2}$$

Notice that the broken-trend model uses the same two-step procedure as an augmented Dickey–Fuller test, except that the indicator function is included in equation (1), the detrending regression.[4]

The results of the two-step testing procedure are displayed in table 2. Panel A fits the best two connected trend lines to the log series and picks June 1968 as the date of the kink in the trend function.[5] Panel B reports a rejection, at the 2.5% significance level, of the unit-root null in favor of the hypothesis that postwar U.S. output is a stationary process around a broken linear trend. The logic behind the broken-trend stationary methodology is conveyed well in a simple picture. In figure 1, the log level of U.S. industrial production is plotted along with the broken-trend function determined in equation (1). The model tells us that from January 1959 to June 1968, industrial production grew at an annualized rate of 6.22%, and it slowed to 2.67% in the remaining period.

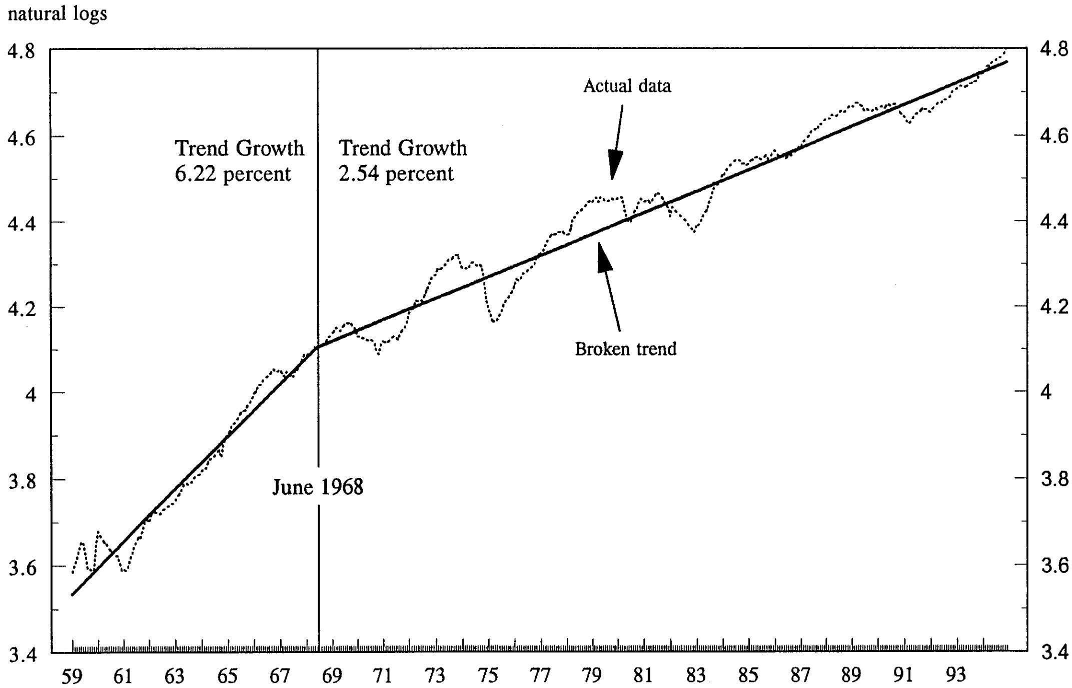
This finding implies that the log-differenced specification for postwar industrial production, suggested in section II, is incorrect. Similar unit-root tests, allowing for a broken-trend function, on the money supply, producer prices, and interest rates, all failed to reject the null hypothesis. Therefore the ''expanded'' list of univariate pretests

[3] Choosing the minimum value imposes a mild a priori restriction on the direction of the slope change (thereby allowing only for a ''slowdown'') and generally gives the test more power (Perron (1994)). Perron describes two break-date selection criteria that do not impose such a restriction. These tests also yield rejections of the unit-root null at the 10% level or better (see appendix A).

[4] The critical values for any unit-root test are dependent on the deterministic regressors included in equation (1) because the trend function must also be estimated (see Campbell and Perron (1991)). The finite-sample critical values for the t-statistic on $\hat{\alpha} = 1$, where the break date is unknown a priori, are reported in Perron (1994). These finite-sample critical values are also not invariant to the data-based procedure used to select the lag length k in equation (2). Here the t-sig criterion is employed, but using another lag-length criterion yields similar results (see appendix A).

[5] This break date occurs earlier than the second quarter 1974 break date for real GNP found by Perron (1994). This discrepancy is not unreasonable since the series being tested, its frequency, and the sample period are all different. Nevertheless, the 1968 break date roughly coincides with the year concerns over slow productivity growth first surfaced, and median family real incomes began their long period of stagnation (Krugman (1990)). In any case, the exact dating of the break point is not of central importance here. Instead, emphasis is placed on the integration properties of the time series.

FIGURE 1—U.S. INDUSTRIAL PRODUCTION INDEX, 1959–1994



suggests that the variables of interest should be specified as follows:

$$\tilde{y}_t = \qquad\qquad \epsilon_{yt}$$

$$\Delta m_t = \mu_m + \beta_{mt} t + \epsilon_{mt}$$

$$\Delta p_t = \mu_p + \qquad \epsilon_{pt}$$

$$\Delta r_t = \qquad\qquad \epsilon_{rt}$$

where all of the series ϵ_t are mean-zero stationary processes.

IV. The Money–Output Correlation

A. Evidence before 1985

The results of the pretests reported in sections II and III are now used to assess empirically the monetarist proposition that the money supply is of central importance in the business cycle. The metric for measuring money's importance, Granger causality, involves OLS estimation of

$$\tilde{y}_t = \mu + \beta t + \sum_{i=1}^{j} \delta_i \Delta m_{t-i} + \sum_{i=1}^{k} \gamma_i \tilde{y}_{t-i} + \sum_{i=1}^{k} \beta_i \Delta p_{t-i} + \sum_{i=1}^{k} \theta_i \Delta r_{t-i} + \epsilon_t \tag{3}$$

and tests the hypothesis that the coefficients on lagged money growth in equation (3) equal zero. A rejection of the null hypothesis is a rejection of the proposition that money is neutral.[6]

[6] A time trend is included in equation (3) because the analysis of section II suggests that we should examine the predictive power of *detrended* money growth. Equivalently, we could run a preliminary detrending regression for money growth and include these residuals in the system,

TABLE 3.—MONEY–OUTPUT CAUSALITY TESTS 1960–1985

Specification	Six Lags of Money: y, m	Six Lags of Money: y, m, p, r	Twelve Lags of Money: y, m	Twelve Lags of Money: y, m, p, r
A. First-Differenced Output, Δy_t				
1. Constant	2.16	2.53	1.31	1.40
	(0.047)	(0.021)	(0.212)	(0.164)
2. Constant, time	4.14	3.15	2.27	1.78
	(5.2×10^{-4})	(0.005)	(0.009)	(0.052)
B. Broken-Detrended Output, $\tilde{y}_t$				
1. Constant	3.56	4.32	1.86	2.34
	(0.002)	(3.5×10^{-4})	(0.039)	(0.007)
2. Constant, time	3.99	3.84	2.38	2.10
	(7.5×10^{-4})	(0.001)	(0.006)	(0.017)

Notes: Table reports F-statistics that test the hypothesis that all of the coefficients on the lags of Δm_t in the respective output equations are zero; p-values are in parentheses. Regressions include either six or twelve lags of money and twelve lags of other variables. Sample period begins in February 1960.

Table 3 shows the F-statistics and p-values from this equation estimated over the Stock and Watson sample period (February 1960 to December 1985). As a benchmark, we first examine the causality statistics using first-differenced output, which are reported in panel A. First, the predictive power of money is substantially enhanced when the time trend is included.[7] Second, using more than six lags of money dilutes the Granger causal relationship. Third, concentrating only on the specification that correctly includes a time trend, the Granger causal relationship from detrended money growth to output is robust to the presence of the Treasury bill rate. These results confirm the conclusions reached by Stock and Watson. Panel B reports the same statistics, but from regressions using broken-detrended output. These results yield similar rejections of the neutrality proposition. For example, using Stock and Watson's preferred specification (time trend included, six lags of money), the null hypothesis is rejected at the 0.1%

[7] Stock and Watson and Friedman and Kuttner report results including a quadratic time trend. Such a specification is not reported since it is not justified by the univariate pretests and adds little to the results.

TABLE 4.—CAUSALITY TESTS USING DIFFERENT INTEREST RATES, 1960–1985

	Six Lags of Money			Twelve Lags of Money		
	*TB*3	*CP*6	*CP*6 and *CP*6–*TB*3 Spread	*TB*3	*CP*6	*CP*6 and *CP*6–*TB*3 Spread
	A. First-Differenced Output $\Delta y_t = f(\Delta\, y_{t-i}, \Delta m_{t-i}, \Delta p_{t-i}, \Delta r_{t-i})$					
F-statistic for money	3.15 (0.005)	1.71 (0.119)	1.95 (0.073)	1.78 (0.052)	1.50 (0.125)	1.37 (0.182)
F-statistic for interest rate	0.83 (0.615)	1.28 (0.229)	0.97 (0.482)	0.70 (0.755)	1.56 (0.103)	0.94 (0.508)
F-statistic for spread			4.82 (4.0×10^{-7})			4.46 (1.8×10^{-6})
	B. Broken-Detrended Output $\tilde{y}_t = f(\tilde{y}_{t-i}, \Delta m_{t-i}, \Delta p_{t-i}, \Delta r_{t-i})$					
F-statistic for money	3.84 (0.001)	2.46 (0.025)	2.94 (0.008)	2.10 (0.017)	1.80 (0.049)	1.76 (0.055)
F-statistic for interest rate	1.25 (0.252)	1.24 (0.258)	1.03 (0.425)	1.07 (0.383)	1.44 (0.147)	0.92 (0.530)
F-statistic for spread			4.52 (1.3×10^{-6})			4.13 (6.6×10^{-6})

Note: See footnotes to table 3.

TABLE 5.—CAUSALITY TESTS USING DIFFERENT INTEREST RATES, 1960–1994

	Six Lags of Money			Twelve Lags of Money		
	*TB*3	*CP*6	*CP*6 and *CP*6–*TB*3 Spread	*TB*3	*CP*6	*CP*6 and *CP*6–*TB*3 Spread
	A. First-Differenced Output $\Delta y_t = f(\Delta y_{t-i}, \Delta m_{t-i}, \Delta p_{t-i}, \Delta r_{t-i})$					
F-statistic for money	1.84 (0.090)	0.91 (0.484)	0.84 (0.543)	1.10 (0.358)	0.70 (0.750)	0.79 (0.657)
F-statistic for interest rate	1.21 (0.274)	1.97 (0.026)	1.63 (0.083)	1.23 (0.259)	2.04 (0.020)	1.62 (0.084)
F-statistic for spread			4.31 (2.1×10^{-6})			4.39 (1.5×10^{-6})
	B. Broken-Detrended Output $\tilde{y}_t = f(\tilde{y}_{t-i}, \Delta m_{t-i}, \Delta p_{t-i}, \Delta r_{t-i})$					
F-statistic for money	2.12 (0.050)	1.17 (0.322)	1.19 (0.311)	1.35 (0.187)	0.85 (0.595)	0.93 (0.514)
F-statistic for interest rate	1.55 (0.104)	1.73 (0.059)	1.68 (0.069)	1.67 (0.071)	1.82 (0.043)	1.68 (0.070)
F-statistic for spread			3.82 (1.7×10^{-5})			3.85 (1.5×10^{-5})

Note: See footnotes to table 3.

level, even in the presence of the Treasury bill rate, no matter which univariate characterization of output is used.

Recall that the short-term interest rate is included in the system because we expect financial *prices* to contain information about the financial market not reflected in financial *quantities* such as the money supply.[8,9] If that information is relevant to the channels through which financial markets affect the real economy, then a test of the effect of money on output should control for interest rate effects. The question then arises, which short-term interest rate should be included? The original Sims (1980) study used the 6-month commercial paper rate (*CP*6), whereas Litterman and Weiss (1985) and Eichenbaum and Singleton (1986) use the Treasury bill rate (*TB*3). Friedman and Kuttner (1992a) argue that the information contained in *CP*6 is particularly relevant for private-sector business expenditures and, therefore, for forecasts of economic activity. Stock and Watson (1989a), Bernanke (1990), and Friedman and Kuttner (1992b) believe that the spread between *CP*6 and *TB*3 predicts future economic activity because it acts as an indicator of the stance of monetary policy and changing perceptions of default risk. Why is this discussion about short-term rates important? Friedman and Kuttner show that the predictive power of money growth for future output growth is substantially reduced if, instead of *TB*3, one includes *CP*6 or the spread between them.

Again focusing on the sample period before December 1985, table 4 displays causality statistics alternatively using *TB*3, *CP*6, and *CP*6 plus the spread. Panel A confirms Friedman and Kuttner's claim that

[8] Note that *M*1, the narrow money aggregate used here, is easily outperformed by other financial quantity variables such as *M*2 or divisia aggregates (see Rotemberg et al. (1995)).

[9] Prices are included to determine whether money helps in forecasting another indicator of economic activity. Though not reported here, the *F*-statistics for money in the producer price equation *always* fail to reject the null hypothesis.

the Granger causal link between money and first-differenced output is weakened if we substitute *CP*6 for *TB*3; even using Stock and Watson's preferred specification, the *p*-value rises from 0.5 to over 10%. How does the broken-detrended output react to the presence of *CP*6? Panel B shows that the firm finding of Granger causality using broken-detrended output is diluted somewhat, but the rejection remains significant below the 5% level using both six and twelve lags of money. When both *CP*6 and the *CP*6–*TB*3 spread are included, the results using first-differenced output depend on the lag length of money. In contrast, when using broken-detrended output, the rejection of the neutrality proposition is robust to the number of lags of money included.

B. Evidence after 1985

A second criticism mounted by Friedman and Kuttner against the Stock and Watson study is that the evidence that the Granger causal relationship between money and output is substantially weakened when the sample period is extended past 1985. Table 5 recalculates the causality statistics from table 4 using data ending in December 1994. In the extended sample, the only robustly positive results are that *CP*6 and the *CP*6–*TB*3 spread continue to be important forecasting variables. In contrast, the nonneutrality result is clearly sensitive to both the choice of interest rate and the lag length of money. Most importantly, the choice between first-differenced and broken-detrended output does not alter the Granger causality results.

Of course the finding that the relationship between narrow money and output has weakened after 1985 should not be too surprising. The Federal Reserve openly discussed its concerns about the reliability of *M*1 as an indicator of monetary policy as early as 1982 and completely abandoned setting target growth ranges for the narrow aggregate by 1987. These results confirm that the Granger causal relationship between *M*1 and output has weakened since 1985 (though note that, using Stock and Watson's preferred specification, it has not entirely broken down), and this conclusion is unchanged whether or not one decides that output has a unit root.

V. Conclusion

The results presented in this paper suggest that the debate over the univariate characterization of U.S. output has only limited impact on the money–output correlation. Including a technique introduced by Perron (1994) as part of a series of univariate pretests, we are able to reject the unit-root null hypothesis for real output (measured by the industrial production index) in favor of a broken-trend stationary alternative. Before 1985, when this broken-detrended (and not first-differenced) output series is used, we find that the positive Granger causal relationship between money and output is robust to both the inclusion of different interest rates and different lag lengths of money. However, when data after 1985 are included, the relationship is not robust to these changes, no matter which univariate characterization of output is used.

Nevertheless, these results suggest that a prudent analysis of many macroeconomic time series, and the interactions between them, should allow for the possibility of a regime change in the deterministic trend function of each series. This is especially true when one of the time series is real output, though the broken-trend stationary characterization may be applicable to other macroeconomic time series. For example, Perron (1994) finds that most of the data examined by Nelson and Plosser (1982) can also be characterized as broken-trend stationary processes, whereas Perron and Vogelsang (1992) use the broken-trend stationary model for measures of purchasing power parity. Therefore macroeconometric analysis that simply detrends the data through first-differencing without first testing for the possibility that the series may be stationary around a broken-trend function may be needlessly disposing of useful information.

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APPENDIX A

The *t*-sig criterion for selecting the lag length k in step 2 requires the *t*-statistic on the coefficient of the last included lag in the autoregression to be significant while the coefficient on the last lag in higher order autoregressions is insignificant, up to some maximum order k_{max}. Throughout the paper, $k_{max} = 10$ is used. When selecting the break date by minimizing the *t*-statistic on the coefficient associated with the slope break (as is done in the paper), the finite-sample ($T = 200$) critical values for rejecting the unit-root null hypothesis using the *t*-sig criteria are reported in table 2. Using an alternative lag-length criterion (*F*-sig described in Perron (1994)) yields the same break date and the same *t*-statistic on α of -4.51. This is also a rejection at the 2.5% level using the appropriate critical values. Asymptotically, the critical values are invariant to the data-based procedure used to select the lag length; the asymptotic critical values are 10% = -3.77, 5% = -4.08, 2.5% = -4.36, 1% = -4.67.

Perron (1994) describes two break-date selection criteria that do not impose an a priori restriction on the direction of the slope change. The first directly minimizes the *t*-statistic on α in equation (2) in the text. Finite-sample ($T = 200$) critical values using this dating procedure and the *t*-sig criterion are 10% = -4.38, 5% = -4.65, 2.5% = -4.96, 1% = -5.28, while the asymptotic critical values are 10% = -4.07, 5% = -4.36, 2.5% = -4.62, 1% = -4.91. Using this dating procedure, we date the slope break at January 1968 and obtain a *t*-statistic for $\alpha = 1$ of -4.54. So the unit-root null hypothesis is rejected at the 10% level. The second maximizes the absolute value of t_γ. Finite-sample ($T = 200$) critical values using this dating procedure and the *t*-sig criterion are 10% = -4.17, 5% = -4.41, 2.5% = -4.75, 1% = -5.02, while the asymptotic critical values are 10% = -4.04, 5% = -4.34, 2.5% = -4.58, 1% = -4.87. This dating procedure yields the same break date and *t*-statistic on α as that reported in the text. Using the above critical values, the unit-root null hypothesis is rejected at the 5% level.

APPENDIX B

Industrial production *IP*	Index 1987 = 100, seasonally adjusted
Producer prices	Index 1982 = 100, not seasonally adjusted
Money supply *M*1	U.S.$, billion, seasonally adjusted
3-month Treasury bill *TB*3	Percent
6-month commercial paper *CP*6	Percent

NONPARAMETRIC DEMAND ANALYSIS OF U.K. PERSONAL SECTOR DECISIONS ON CONSUMPTION, LEISURE, AND MONETARY ASSETS: A REAPPRAISAL

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Abstract—This paper utilizes a new data set to test for utility maximizing behavior and weakly separable subutility functions in the context of a utility function comprising durables, nondurables, services, leisure, and monetary asset holdings for the U.K. personal sector. All the data sets analyzed demonstrated consistency with respect to utility maximizing behavior. The weak separability results prove to be relatively invariant to the degree of aggregation over goods but highly sensitive to the assumption made regarding the representative consumer. Per-household scaling of the data produced a utility function that is weakly separable in goods, services and leisure, and in monetary assets.

I. Introduction

In many branches of empirical economics analysts often make implicit weak separability assumptions in order to allow them to focus on subsets of the consumer choice problem. Relatively little empirical work has been undertaken, however, to investigate the structure of consumer preferences in the context of fully specified utility functions containing consumer goods, services, leisure, and monetary assets. Exceptions include Swofford and Whitney (1987, 1988) for the United States and Patterson (1991) for the U.K. personal sector. Patterson found relatively few specifications of goods that did not produce violations of the generalized axiom of revealed preference (GARP) for the entire data set, however, making it very difficult to undertake wide-ranging tests of weak separability. This paper reexamines the issue of weak separability across these goods for the U.K. personal sector using Varian's nonparametric technique (Varian (1983)) in the context of a disaggregated set of monetary assets. The sensitivity of the results is tested with respect to both representative consumer scaling[1] and the aggregation or disaggregation of consumer goods. Sensitivity analysis is also conducted into the weak separability results themselves using recently developed goodness-of-fit measures for revealed preference tests (Varian (1991) and Ward (1994)).

II. Data

The specified utility function consists of nondurable goods, services, durable goods, leisure, and monetary assets (see table 1). For nondurables and services, the appropriate quantity series are taken to be the real expenditure flows (constant prices) per quarter.[2] The appropriate quantity for durable goods, however, is the net stock and not the expenditure flow over a given period.[3] Leisure hours per quarter are defined as [98 hours − (average hours worked per week during the quarter)] × 13.[4] Data on the nominal stocks of each asset

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[1] Data on the U.K. population aged over 18 and the number of households were provided by the Office of Population and Census Statistics (OPCS).
[2] Quarterly data on real expenditure flows and implicit price deflators were obtained from Datastream.
[3] The net stock is calculated using data on real expenditure flows together with unpublished CSO data on average life lengths for the main components of durables. The latter were used to calculate appropriate depreciation rates.
[4] Average weekly hours worked are derived from the annual (April) New Earnings Survey. Quarterly data were derived using linear interpolation.