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Nominal Stylized Facts of U.S. Business Cycles

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This paper investigates the basic nominal stylized facts of business cycles in the United States using monthly data from 1960:1 to 1993:4 and the methodology suggested by Kydland and Prescott (1990). Comparisons are made among simple-sum and Divisia aggregates using the Thornton and Yue (1992) series of Divisia monetary aggregates. The robustness of the results to (relevant) nonstochastic stationarity-inducing transformations is also investigated.

Kydland and Prescott (1990) argue that business cycle research took a wrong turn when researchers abandoned the effort to account for the cyclical behavior of aggregate data following Koopmans's (1947) criticism of the methodology developed by Burns and Mitchell (1946) as being "measurement without theory." Crediting Lucas (1977) with reviving interest in business cycle research, Kydland and Prescott initiated a line of research that builds on the growth theory literature. Part of it involves an effort to assemble business cycle facts. This boils down to investigating whether deviations of macroeconomic aggregates from their trends are correlated with the cycle, and if so, at what leads and lags.

Kydland and Prescott (1990) report some original evidence for the U.S. economy and conclude that several accepted nominal facts, such as the procyclical movements of money and prices, appear to be business cycle myths. In contrast to conventional wisdom, they argue that the price level (whether measured by the implicit GNP deflator or by the consumer price index), is countercyclical. Although the monetary base and M1 are both procyclical, neither leads the cycle. This evidence counters Mankiw's (1989) criticism of real business cycle models on the grounds that they do not predict procyclical variation in prices. Moreover, the evidence of countercyclical price behavior has been confirmed by Cooley and Ohanian (1991), Backus and Kehoe (1992), Smith (1992), and Chadha and Prasad (1994).

The cyclical behavior of money and prices has important implications for the sources of business cycles and therefore for discriminating among competing models. Initially it was argued, for example, that procyclical prices will be consistent with demand-driven models of the cycle, whereas countercyclical prices would be consistent with predictions of supply-determined models, including real business cycle models. Subsequently, however, Hall (1995) has shown that adding more detail to traditional demand-driven models can produce countercyclical prices, whereas Gavin and Kydland (1995) have shown that alternative money supply rules can generate either procyclical or countercyclical prices in a real business cycle setting.

The objective of this paper is to reexamine the cyclical behavior of money and prices using monthly U.S. data. For comparison purposes, the methodology used is mainly that of Kydland and Prescott (1990). Therefore in accordance with the real business cycle approach to economic fluctuations, we define the growth of a variable as its smoothed trend and the cycle components of a variable as the deviation of the actual values of the variable from the smoothed trend. However, we investigate robustness of the results to alternative (relevant) nonstochastic stationarity-inducing transformations.

To highlight the influence of money measurement on statistical inference [as in Belongia (1996)], comparisons are made among simple-sum and Divisia monetary aggregates (of M1A, M1, M2, M3, and L) —see Barnett, Fisher, and Serletis (1992) regarding the state of the art in monetary aggregation. The money measures employed are monthly simple-sum and Divisia indexes (from 1960:1 to 1993:4), as described in Thornton and Yue (1992), and were obtained from the Federal Reserve Economic Data (FRED) bulletin board of the Federal Reserve Bank of St. Louis.

The paper is organized as follows. Section 1 briefly discusses the Hodrick Prescott (HP) filtering procedure for decomposing time series into long-run and business cycle components. Section 2 presents HP empirical correlations of money, prices, and nominal interest rates with industrial production. In section 3 we investigate the robustness of our results to alternative stationarity-inducing transformations, and in the last section we summarize the main results and conclude.

METHODOLOGY

For a description of the stylized facts, we follow the current practice of detrending the data with the HP filter—see Prescott (1986). For the logarithm of a time series X_t, for $t = 1,2,...,T$, this procedure defines the trend or growth component, denoted τ_t, for $t = 1,2,...,T$, as the solution to the following minimization problem

$$\min_{\tau_t} \sum_{t=1}^{T} (X_t - \tau_t)^{2+\mu} \sum_{t=2}^{T-1} \left[(\tau_{t+1} - \tau_t) - (\tau_t - \tau_{t-1}) \right]^2$$

so $X_t - \tau_t$ is the filtered series. The larger the μ, the smoother the trend path, and when $\mu = \infty$, a linear trend results. In our computations, we set $\mu = 129{,}600$, as it has been suggested for monthly data. Note that the monthly cyclical components defined by $\mu = 129{,}600$ approximately average to the quarterly components defined by $\mu = 1{,}600$ which is commonly used to define business cycle fluctuations in research literature.

We measure the degree of co-movement of a series with the pertinent cyclical variable by the magnitude of the correlation coefficient $\rho(j)$, $j \in \{0, \pm 1, \pm 2, ...\}$. The contemporaneous correlation coefficient—$\rho(0)$—gives information on the degree of contemporaneous co-movement between the series and the pertinent cyclical variable. In particular, if $\rho(0)$ is positive, zero, or negative, we say that the series is procyclical, acyclical, or countercyclical, respectively. In fact, for $0.23 \leq |\rho(0)| < 1$, $0.10 \leq |\rho(0)| < 0.23$, and $0 \leq |\rho(0)| < 0.10$, we say that the series is strongly contemporaneously correlated, weakly contemporaneously correlated, and contemporaneously uncorrelated with the cycle, respectively. Following Fiorito and Kollintzas (1994) in our sample of 400 observations, the cutoff point 0.1 is close to the value 0.097 that is required to reject the null hypothesis, $H_0 : \rho(0) = 0$, at the 5 percent level in a two-sided test for bivariate normal random variables. Also, the cutoff point 0.23 is close to the value of 0.229 that is required to reject the null hypothesis H_0: $|\rho(0)| \leq 0.5$, in the corresponding one-tailed test. Also, $\rho(j)$, $j \in \{\pm 1, \pm 2, ...\}$—the cross correlation coefficient—gives information on the phase-shift of the series relative to the cycle. If $|\rho(j)|$ is maximum for a negative, zero, or positive j, we say that the series is leading the cycle by j periods, is synchronous, or is lagging the cycle by j periods, respectively.

HODRICK-PRESCOTT STYLIZED FACTS

In Table 1 we report contemporaneous correlations, as well as cross correlations (at lags and leads of one through six months) between the cyclical components of money and the cyclical component of industrial production. We see that all the monetary aggregates are strongly procyclical. With a minor exception for M1A, for both Divisia and simple-sum measures, the

Table 1

Correlations of HP-Filtered Sum and Divisia Monetary Aggregates with Industrial Production*

Variable, x	Volatility	Correlation Coefficients of Industrial Production with												
		x_{t-6}	x_{t-5}	x_{t-4}	x_{t-3}	x_{t-2}	x_{t-1}	x_t	x_{t+1}	x_{t+2}	x_{t+3}	x_{t+4}	x_{t+5}	x_{t+6}
Sum M1A	2.09	0.43	0.43	0.43	0.43	0.42	0.40	0.38	0.35	0.31	0.28	0.25	0.24	0.22
Sum M1	1.93	0.37	0.37	0.37	0.36	0.35	0.32	0.28	0.24	0.19	0.15	0.11	0.08	0.05
Sum M2	1.41	0.71	0.70	0.66	0.62	0.56	0.49	0.40	0.32	0.24	0.16	0.09	0.03	−0.03
Sum M3	1.48	0.50	0.52	0.53	0.53	0.52	0.50	0.47	0.44	0.41	0.38	0.35	0.32	0.29
Sum L	1.11	0.33	0.39	0.44	0.49	0.52	0.55	0.57	0.58	0.59	0.58	0.58	0.56	0.55
Divisia M1A	1.74	0.39	0.40	0.40	0.40	0.39	0.37	0.35	0.32	0.29	0.27	0.26	0.25	0.24
Divisia M1	1.50	0.28	0.28	0.29	0.28	0.27	0.24	0.21	0.18	0.14	0.10	0.08	0.05	0.03
Divisia M2	1.81	0.67	0.65	0.62	0.59	0.54	0.47	0.40	0.33	0.25	0.19	0.13	0.08	0.03
Divisia M3	1.78	0.68	0.67	0.66	0.64	0.60	0.56	0.50	0.45	0.39	0.34	0.29	0.25	0.21
Divisia L	1.58	0.62	0.63	0.64	0.64	0.62	0.60	0.57	0.53	0.49	0.45	0.41	0.37	0.33

* Monthly data from sample period 1960:1–1993:4.

broader the aggregate the more procyclical it is. There is also evidence that M2 money, however defined, leads the cycle by more than the other aggregates and, if anything, Sum L is slightly lagging. These results suggest the only major differences among simple-sum and Divisia monetary aggregates occur in the stronger correlation at leads for the broad Divisia aggregates, M3 and L.

We interpret these results as being generally consistent with the cyclical money behavior in the United States reported (using quarterly data) by Kydland and Prescott (1990) and Belongia (1996). Unlike Belongia, who like Kydland and Prescott, uses quarterly data and only the simple-sum and Divisia measures of M1 and M2, we find no significant differences across narrow simple-sum and Divisia monetary aggregates. We find strong contemporaneous correlations between broad-sum and Divisia money and the cyclical indicator. Divisia L, however, is leading the cycle, and Sum L is slightly lagging the cycle. This result seems to be consistent with the evidence reported by Barnett, Offenbacher, and Spindt (1984), who found that Divisia L was the best aggregate in terms of causality tests, produced the most stable demand-for-money function, and provided the best reduced-form results.

Next we turn to the statistical properties of the cyclical components of the price level (measured by the consumer price index) and two short-term nominal interest rates (to deal with anomalies that arise because of different ways of measuring financial market price information)—the Treasury bill rate and the commercial paper rate. The Treasury bill rate is the interest rate on short-term, unsecured borrowing by the U.S. government, whereas the commercial paper rate is the interest rate on short-term, unsecured borrowing by corporations. As Friedman and Kuttner (1993, p. 194) argue, the commercial paper rate is superior in capturing the information in financial prices because "the commercial paper rate more directly reflects the cost of finance corresponding to potentially interest-sensitive expenditure flows than does the Treasury bill rate."

Table 2 reports HP cyclical correlations of prices and short-term nominal interest rates with industrial production. We see that the price level is strongly countercyclical, whereas both the Treasury bill rate and the commercial paper rate are strongly procyclical and lag the cycle. These results provide strong confirmation

Table 2

Correlations of HP-Filtered Prices and Short-Term Nominal Interest Rates with Industrial Production*

Variable, x	Volatility	Correlation Coefficients of Industrial Production with												
		x_{t-6}	x_{t-5}	x_{t-4}	x_{t-3}	x_{t-2}	x_{t-1}	x_t	x_{t+1}	x_{t+2}	x_{t+3}	x_{t+4}	x_{t+5}	x_{t+6}
Consumer Price Index	1.46	−0.73	−0.71	−0.68	−0.65	−0.60	−0.55	−0.48	−0.43	−0.37	−0.31	−0.25	−0.20	−0.15
Treasury Bill Rate	1.66	−0.17	−0.09	0.01	0.11	0.22	0.32	0.40	0.44	0.46	0.47	0.47	0.48	0.48
Commercial Paper Rate	1.44	−0.12	−0.03	0.05	0.15	0.25	0.33	0.39	0.42	0.43	0.43	0.43	0.43	0.43

* Monthly data from sample period 1960:1–1993:4.

for the countercyclical price behavior in the United States reported by Kydland and Prescott (1990), Cooley and Ohanian (1991), Backus and Kehoe (1992), Smith (1992), and Chadha and Prasad (1994). They clearly support the Kydland and Prescott (1990) claim that the perceived fact of procyclical prices is but a myth.

ROBUSTNESS TO STATIONARITY-INDUCING TRANSFORMATIONS

We have characterized the key nominal features of U.S. business cycles using a modern counterpart of the methods developed by Burns and Mitchell (1946)—HP cyclical components. The HP filter is used almost universally in the real business cycle research program and extracts a long-run component from the data, rendering stationary series that are integrated up to the fourth order. HP filtering, however, has recently been questioned as a unique method of trend elimination. For example, King and Rebelo (1993) argue that HP filtering may seriously change measures of persistence, variability, and co-movement. They also give a number of examples that demonstrate that the dynamics of HP filtered data can differ significantly from the dynamics of differenced or detrended data.

Also, Cogley and Nason (1995), in analyzing the effect of HP filtering on trend- and difference-stationary time series, argue that the interpretation of HP stylized facts depends on assumptions about the time series properties of the original data. For example, when the original data are trend stationary, the HP filter operates like a high-pass filter. That is, it removes the low frequency components and allows the high frequency components to pass through. When the original data are difference stationary, however, the HP filter does not operate like a high-pass filter. In this case, HP stylized facts about periodicity and co-movement are determined primarily by the filter and reveal very little about the dynamic properties of the original data.

More recently, however, Baxter and King (1995) argue that HP filtering can produce reasonable approximations of an ideal business cycle filter. Though we believe that the results based on the HP filter are reasonably robust across business cycle filters, we believe it is useful to compare what we are doing with alternative popular methods of detrending the data. Once, however, we abstract from growth theory, we need to make some assumption about the trend. In particular, deterministic detrending will be the appropriate stationarity-inducing transformation under trend stationarity and differencing under difference stationarity.

Results reported in Koustas and Serletis (1996), based on augmented Dickey-Fuller–type regressions, indicate that the null hypothesis of a unit root in levels cannot be rejected for any of the

Table 3

Correlations of First Differences of Sum and Divisia Money with First Differences of Industrial Production*

Variable, x	Volatility	Correlation Coefficients of Industrial Production with												
		x_{t-6}	x_{t-5}	x_{t-4}	x_{t-3}	x_{t-2}	x_{t-1}	x_t	x_{t+1}	x_{t+2}	x_{t+3}	x_{t+4}	x_{t+5}	x_{t+6}
Sum M1A	0.005	0.09	0.06	0.05	0.17	0.14	0.12	0.10	0.06	−0.04	−0.08	−0.08	−0.03	−0.06
Sum M1	0.004	0.09	0.08	0.07	0.17	0.14	0.12	0.05	0.04	−0.05	−0.12	−0.08	−0.04	−0.05
Sum M2	0.003	0.25	0.23	0.21	0.27	0.23	0.16	0.11	0.04	−0.07	−0.07	−0.07	−0.05	−0.04
Sum M3	0.003	0.16	0.18	0.17	0.21	0.17	0.13	0.11	0.10	0.04	0.03	0.05	0.04	0.01
Sum L	0.003	0.10	0.11	0.07	0.12	0.10	0.14	0.15	0.17	0.11	0.11	0.11	0.08	0.09
Divisia M1A	0.005	0.04	0.03	0.01	0.14	0.11	0.09	0.04	0.02	−0.07	−0.08	−0.06	−0.02	−0.02
Divisia M1	0.004	0.04	0.05	0.03	0.14	0.10	0.08	−0.02	−0.01	−0.06	−0.10	−0.04	−0.01	−0.04
Divisia M2	0.004	0.18	0.20	0.17	0.28	0.25	0.20	0.07	0.02	−0.09	−0.10	−0.07	−0.05	−0.04
Divisia M3	0.003	0.17	0.20	0.18	0.27	0.24	0.19	0.08	0.06	−0.03	−0.04	0.00	0.01	0.02
Divisia L	0.003	0.14	0.16	0.13	0.24	0.21	0.21	0.12	0.12	0.02	0.02	0.04	0.05	0.07

*Monthly data from sample period, 1960:1–1993:4.

Table 4

Correlations of First Differences of Prices and Short-Term Nominal Interest Rates with First Differences of Industrial Production*

Variable, x	Volatility	Correlation Coefficients of Industrial Production												
		x_{t-6}	x_{t-5}	x_{t-4}	x_{t-3}	x_{t-2}	x_{t-1}	x_t	x_{t+1}	x_{t+2}	x_{t+3}	x_{t+4}	x_{t+5}	x_{t+6}
Consumer Price Index	0.003	−0.23	−0.22	−0.30	−0.23	−0.23	−0.16	−0.08	−0.07	−0.07	0.00	−0.01	−0.04	−0.07
Treasury Bill Rate	0.006	−0.10	−0.06	−0.06	−0.08	0.13	0.20	0.30	0.24	0.18	0.04	−0.00	0.02	−0.00
Commercial Paper Rate	0.006	−0.03	−0.02	−0.08	−0.04	0.14	0.21	0.23	0.23	0.11	0.03	−0.03	0.02	0.02

*Monthly data from sample period 1960:1–1993:4.

series used here, whereas the null hypothesis of a second unit root is rejected except for Sum M3, Sum L, and the price level which appear to be integrated of order 2 [or I(2) in Engle and Granger (1987) terminology]. Based on this evidence, in Tables 3 and 4 we report correlations (in the same fashion as in Tables 1 and 2) based on differenced data, keeping in mind that although differencing yields stationary series, these stationary series do not in general correspond to cyclical components. See, for example, Baxter and King (1995). These results are generally supportive of the hypothesis of acyclical money and price behavior. Nominal interest rates appear to be strongly procyclical and lagging slightly.

CONCLUSION

In this paper we investigated the cyclical behavior of U.S. money, prices, and short-term nominal interest rates, using monthly data from 1960:1 to 1993:4 and the methodology of Kydland and Prescott (1990). Based on stationary HP cyclical deviations, our results fully match recent evidence on the countercyclicality of the price level. We also found that short-term nominal interest rates are strongly procyclical and that money is in general procyclical. Furthermore, the evidence suggests that there are only slight differences across narrow simple-sum and Divisia money measures.

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