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LONGTERM BEHAVIOUR OF TERMS OF TRADE OF PRIMARY PRODUCTS VIS-A-VIS MANUFACTURES : A CRITICAL REVIEW OF RECENT DEBATE

Prabirjit Sarkar

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CENTRE FOR STUDIES IN SOCIAL SCIENCES, CALCUTTA
10, Lake Terrace, Calcutta - 700 029

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Abstract

After the publication of some long period data on terms of trade of primary products by Grilli and Yang(1988), a new debate started - whether the terms of trade series followed a deterministic path of decline or not. This paper critically examines this debate and concludes that the new series supports the Prebisch-Singer hypothesis. The terms of trade of primary products vis-a-vis manufactures have a statistically significant declining trend over the period, 1900-86 and the eruption of debt crisis in the 1980s accentuated the situation. This is a deterministic trend, not simply the permanent effect of a temporary shock.

I

The classical writers believed in a secular improvement in the terms of trade (Commodity Terms of Trade, CTT or the Net Barter Terms of Trade, NBTT) of primary products vis-a-vis manufactures.[1] The policy implication of the classical proposition is that an agricultural country need not industrialize to enjoy the fruits of technical progress in manufactures; free play of international market forces will distribute the gains from technical progress of the industrial countries to the agricultural countries by turning the terms of trade to the favour of primary product exporting countries.

However, during the middle of the present century, evidence to the contrary of the classical proposition was provided (League of Nations, 1945, United Nations, 1949). On the basis of this evidence, Prebisch (1950, 1959, 1964) and Singer (1950, 1975) launched the hypothesis of secular decline in the terms of trade of primary products and the primary producing less developed countries.[2] The policy conclusion that followed was 'inward looking' industrialisation by suspension of free play of market forces (see also Myrdal, 1956).[3]

The Prebisch-Singer(P-S) hypothesis and its policy conclusion created great controversy in the academic world. In particular, the data base of the Prebisch-Singer hypothesis was vehemently challenged. In the first half of the last decade, some attempts were made to refute all these criticisms (see Spraos, 1980, 1983; Thirlwall and Bergevin, 1985; Sapsford, 1985; Sarkar, 1986a and specially 1986b). However, the debate has not subsided. Rather the publication of some long period data by Grilli and Yang, (G-Y 1988) and their support of the Prebisch-Singer hypothesis has generated further controversy. Perhaps an easy access to a long period series provoked a number of economists to apply the computer packages incorporating the latest developments in the field of time series trend study. Thus starting from an elementary time series analysis on the basis of quinquennial averages (League of Nations, 1945; United Nations, 1949; Prebisch, 1950), the Prebisch-Singer controversy has reached the stage of 'hi-tech' statistical debate. The object of this paper is to critically examine this hi-tech debate.

II

In the early days, the standard empirical approach was to model a time series such as terms of trade as stationary around a deterministic trend. Consider a long-linear trend,

$$1Y(t) = a + b.t + u(t) \qquad (1)$$

where $1Y(t)$ is a time series in natural log, a is the intercept, b is the slope (the rate of growth of $Y(t)$), t = time and $u(t)$ is the error term.

Ignoring shortrun fluctuations inherent in the error process, u (t), the path of $1Y(t)$ in the longrun is assumed to be completely deterministic :

> Neither current nor past events will alter long-term expectations ... uncertainty is bounded, even in the infinitely distant future (Nelson & Plosser, 1982, p-142).

Since the error process is assumed to be stationary, all shocks are necessarily temporary or cyclical in nature. This type of 'determimistic' trend model has been termed 'trend-stationary' (TS) by Nelson and Plosser (1982).

Until recently, the tests of the Prebisch-Singer hypothesis were done by fitting a deterministic trend model (Spraos, 1980, 1983; Sarkar, 1986a, b; Grilli and Yang, 1988). However, as Cuddington-Urzua, (1989, p. 432) argued,

> it is difficult to justify *a priori* the assumption that the trend line for NBTT is completely deterministic. Presumably, real as well as nominal shocks affect the supplies and demand for commodities over time. Some of the shocks reflect permanent, secular or structural shifts (e.g. resource discoveries or technical innovations), while other shocks reflect

transitory or cyclical phenomena (e.g., monetary policy shocks, or crop failures).

If this is the case, the error process exhibits non-stationarity. Then the growth path of a time series is not deterministic but stochastic so that the entire growth path shifts upwards or downwards over time as shocks occur.

> Shocks in the growth path represent permanent (secular) effects. Fluctuations around the (shifting) growth path are cyclical effects (Cuddington and Urzua, 1989).

Use of a deterministic trend model (Eq. 1) when the true model is stochastic, is a misspecification where the conventional tests of a time trend are entirely biased towards finding a trend when none is present (Nelson & Kang, 1984). The procedure suggested here is to difference the data (to make the error process 'stationary') and fit an a alternative model which was termed as 'difference-stationary' (DS) model by Nelson and Plosser (1982) :

$$d\,1Y(t) = b + du(t) \qquad (2)$$

where d stands for first differencing.

The choice of an appropriate model requires a test of stationarity of the error process. One popular test is the Dicky-Fuller (DF) test/Augmented Dicky-Fuller (ADF) test.[4] However, as Perron (1989) argued, an error process of a trend-stationary (TS) model may be non-stationary not because of fitting a TS model when a DS model is appropriate but because of a structural shift in intercept and/or slope parameters of the TS model during the period under study. Hence a test of stationarity should be preceded by an examination of the series to determine whether data contain the impact of any 'exogenous events'. If there is any evidence of this impact, Perron (1989) test of stationarity should be applied.

This is what Cuddington-Urzua (1989) did to examine the Prebisch–Singer hypothesis on the basis of the Grilli–Yang series. They examined the Grilli-Yang series over the period 1900-1983 and found some evidence of one time drop in 1921. Therefore they conducted Perron test of stationarity with dummies :

$$d\,1Y(t) = a + c.D + e.D_s + p.1Y(t-1) + \sum_{i=1}^{k} p_i.\, d1Y(t-i) + u(t) \qquad (3)$$

where D = 1 between 1900 and 1920 and zero thereafter, D_s is a spike dummy that equals unity only in 1921 and zero otherwise; k is the order of lags so chosen as to insure that the residuals, u(t), are 'white noise' - free from autocorrelated errors.

Cuddington-Urzua (1989) chose an 8 - lag version of Eq-(3) and observed that the Perron t-ratio (t- statistic of the coefficient p) shrank monotonically from - 2.71 at 5 lags to –2.46 at 8 lags. All these are lower (in absolute value) than the critical value at the 10 per cent level calculated by Cuddington-Urzua (-2.97) indicating non-stationarity in the Grilli-Yang series. However, for four lags or less, the null of non-stationarity was rejected. But Cuddington-Urzua (1989) preferred the 8 lag model as the 8th lag of d 1Y (t) was statistically significant. Hence their conclusion was one of non-stationarity.

In the next step, they fitted a trend-stationary model with intercept dummy (D) and difference-stationary models with and without spike dummy (D_s) :

$$1Y(t) = a + b.t + c.D + u(t) \qquad (4)$$

$$d1Y(t) = b + u(t) \qquad (5)$$

$$d1Y(t) = b + e.D_s + u(t) \qquad (6)$$

In no case was the estimate of b found to be statistically significant. Rather, the coefficient of intercept dummy D was found to be significantly positive and the coefficient of spike dummy, D_s, was found to be significantly negative. Hence their conclusion was that there was no evidence of 'secular deterioration in commodity prices, but only a permanent one-time drop in prices after 1920' (Cuddington-Urzua, 1989, p-438).

However, Powell (1991) tried to show that there exist three negative jumps, one in 1921 as noted by Cuddington-Urzua (1989) and two other negative jumps in 1938 and 1975; apart from these 3 jumps, the Grilli–Yang series is trend–less. To show this, Powell (1991) used an indirect approach – the Engle-Granger cointegration analysis. He observed that the numerator (commodity price index, P_x) and the denominator (manufactured goods price index, P_m) of the Grilli-Yang terms of trade index are non-stationary in log-levels but stationary at first (log) difference. That is to say, these are 'integrated' of order 1. In the next step, it was examined whether the two price series(P_x and P_m) are 'cointegrated'. The cointegrating relationship fitted is:

$$\log P_x(t) = A + g.\log P_m(t) + u(t) \qquad (7)$$

It is then examined whether the residuals, u (t), of Eq.(7) are stationary. If the terms of trade series is stationary (trendless), commodity prices and manufactured goods prices must be cointegrated (i.e. the residuals u must be stationary) with the cointegrating parameter of unity (g= 1).

From this procedure, no evidence of cointegration was found. However, if a time trend is added to Eq.(7), strong evidence of cointegration can be found. But there was no evidence that g = 1.

Powell (1991) argued that instead of trend, one should include a jump variable as there was some sign of negative jumps in the commodity price series (P_x) in 1921, 1938 and 1975. The jump variable was so defined as to assume the value zero in all the years excepting the three years of negative jumps; in these three years, it assumes the values that equal the differences between the actual changes in the terms of trade in these three years and the average change in the series over the whole period of study. Adding this jump variable to Eq. (7) supports cointegration at 5% level in DF and Sargan-Bhargava (1983) Durbin-Watson (D-W) tests, at 1% level in Augmented Dicky-Fuller (ADF) test. In the 'most reliable' Johansen (1988) procedure, the hypothesis of cointegration was accepted only at 10% level. The hypothesis that the cointegrating parameter is unity (g=1) was accepted in the case of ADF test while in the case of Johansen (1988) procedure, it was close to one although the hypothesis could not be tested as the computer package used did not report the standard error of the cointegrating parameter.

Both Cuddington-Urzua (1989) and Powell (1991) can be criticised on various grounds. The criticism of Powell (1991) is obvious. Their evidence is not robust. While DF test does not show cointegration at 1% level, the Johansen procedure finds evidence of cointegration only at 10% level. Moreover, the procedure is indirect. It passes through various steps in which tests adopted have low power (Helg, 1990). Lastly, the statistical inference for cointegration analysis in the presence of structural break is not yet developed (Granger & Engle,1991). So when Cuddington-Urzua (1989), Powell (1991) himself and others pointed out breaks in the series, the cointegration analysis may not be justified.

The analysis of Cuddington-Urzua (1989) provoked many critics. One of the critics is Helg (1990). In analogy with the literature on model specification in presence of outliers (see Tsay, 1986), he distinguished between an additive outlier (AO) model and an innovational outlier (IO) model. In an AO process, the reaction to a shock is immediate while in the IO process, the reaction is more gradually spread over time. From an examination of the graph of the Grilli-Yang series. Helg (1990) concluded that an AO specification is appropriate whereas Cuddington-Urzua (1989) chose an IO model. So an AO model was fitted over the period, 1900-88 through a two step procedure. Firstly, the raw series was detrended to obtain the estimated residuals, u(t), from a regression which allows for a time trend with possible changes in slope and intercept after 1920 :

$$1Y(t) = a + b.t + c.D + f(D.t) + u(t) \qquad (8)$$

where D= 1 and D.t = t for t < 1921 and = 0 otherwise.

Secondly, the following regression was estimated by Ordinery Least Square (OLS) method :

$$du(t) = pu(t-1) + \sum_{i=1}^{k} p_i . du(t-i) + u'(t) \qquad (9)$$

where u (t) is the residual of Eq. (8) and u' (t) is the residual of Eq. (9)

For k=1, the t-value of p of the Eq. (9) was estimated to be

(i) - 4.91 for most general specification with time trend and slope dummy; (ii) - 4.81 with time trend and without slope dummy — the 'crash' hypothesis of Perron (1989) and (iii) – 4.24 with no time trend and slope dummy. That means, the null hypothesis of non-stationarity is rejected in all the cases. This conclusion is robust irrespective of whether a deterministic trend is included or not. Moreover, the time trend was found to be significantly negative and both the intercept and slope dummies were found to be significantly positive. Thus Helg (1990) concluded that the true process in the Grilli - Yang series is a trend-stationary one with a deterministic structural change in 1921. He argued that the main reason for the different conclusion of Cuddington-Urzua (1989) is their different specification of the dynamic structure regarding the effects of the exogenous shock in 1921.

However, the close scrutiny of the Cuddingtion-Urzua (1989) analysis by Sapsford-Sarkar-Singer (1992) shows that choice of the DS model is inconclusive even in their IO specification. It was observed that although the 8th order lagged term in Eq: (3) is significant, most of the lower order lag coefficients are not significant. Replicating Perron's test with the non-significant lag terms deleted, it was shown that their conclusion of non-stationarity does not hold good. Deleting the first, third, fifth, sixth and seventh lag terms on the ground that each has an absolute t-value less than 1.3 in Cuddingtion-Urzua (1989) formulation, the Perron test statistic can be found to be -3.26; if the fourth lag term is also dropped as its t-value is only 1.49, the test statistic becomes -3.46. The Lagrange multiplier test statistic for 12th order autocorrelation indicates a white noise error process in each of the cases (Sapsford-Sarkar-Singer, 1992,pp 329-30).

The analysis of Sapsford-Sarkar-Singer (1992) can be carried further. Eq. (3) has been fitted for alternative values of k. The best fit model has been chosen through Akaike criterion. For all the values of k from 0 to 8, it has been found that the Akaike value (AIC) is the lowest for k = 0. The estimate of Eq. (3) for k = 0 is reported below (hereafter t-ratios are in parentheses) :

$$dlY(t) = \underset{(3.26)}{1.58} + \underset{(2.36)}{0.11}(D) + \underset{(2.26)}{0.26}(D_s) - \underset{(-3.27)}{0.34}[lY(t-1)] + u(t) \tag{10}$$

$$\bar{R}^2 = 0.10,\ D\text{-}W = 2.17,\ F(3,81) = 4.28,\ AIC = -4.41$$

Evidently the absolute value of Perron statistic (-3.27) is much higher than the critical value chosen by Cuddingtion-Urzua (1989).

In the IO specification of Cuddingtion-Urzua (1989) given in Eq. (3), trend term was not included. It seems peculiar since the question they were trying to answer was whether the Grilli-Yang series has a deterministic time trend or not (See also Helg, 1990). Including a time trend in Eq. (3) and applying Akaike criterion (AIC), it has been observed again that the best fit model is one with k = 0 and that the absolute value of Perron statistic is much higher than critical value chosen by Cuddington-Urzua (1989) :

$$dlY(t) = \underset{(3.62)}{1.77} - \underset{(-1.83)}{0.0013}(t) + \underset{(1.22)}{0.07}(D) + \underset{(1.96)}{0.22}(D_s) - \underset{(-3.54)}{0.37}[1Y(t-1)] + u(t) \tag{11}$$

$$\bar{R}^2 = 0.13,\ D\text{-}W = 2.18,\ F(4,80) = 4.14,\ AIC = -4.42$$

In the next step, the two step procedure of estimating AO model followed by Helg (1990) has been replicated. Eqs. (8) and (9) have been estimated for alternative values of k. Akaike criterion suggests 1 year lag model (k=1) irrespective of whether the slope dummy (D.t) is included or not. Perron statistic has been estimated to be -4.88 (-4.73) when slope dummy is included (excluded). Thus irrespective of choosing IO model or AO model, the conclusion of non-stationarity drawn by Cuddington-Urzua (1989) does not follow.

Furthermore, the Cuddingtion-Urzua (1989) conclusion of non-stationarity is very sensitive to the particular values taken by the G–Y index over the First World War period 1914-1921 (Sapsford-Sarkar-Singer, 1992). Some kind of interpolation was made by Grilli-Yang (1989) to obtain data for these years. If these are replaced by the inverse of British terms of trade data available in Schlote (1938), Perron's test decisively rejects the non-stationarity hypothesis. Moreover the reestimation of the trend-stationary model of Cuddington-Urzua (1989) as given in Eq.(4) shows a significant negative trend over the period 1900-1983 analysed by Cuddington-Urzua and also over longer period 1900-1986. Not only that, adding another intercept dummy which takes the value one for 1950 onwards and zero otherwise (as chosen by Sapsford, 1985) does not alter the conclusion of a significant negative trend over the period, 1900-83/86.

The sensitivity analysis of Sapsford-Sarkar-Singer (1992) shows that even relatively modest revisions of the 1921 observation of the Grilli-Yang series (with all other observations unchanged) are sufficient to totally reverse the Cuddington-Urzua conclusion. For instance, if the 1921 observation of the Grilli-Yang series is revised to be 25 rather than 50% below the 1920 value (a decline of twice that shown by Schlote's index) Perron test rejects the non-stationarity hypothesis.

In view of this sensitivity, the first two decades may be excluded from the period of study. Helg (1990) considered the period 1921-1988 and applied not only Dicky-Fuller/ Augmented Dicky-Fuller test but also other tests such as Likelihood Ratio Test (Dicky-Fuller, 1981), Sargan-Bhargava (1983) Durbin-Watson (D-W) test and Lagrange Multiplier test (Schmidt and Phillips, 1992). None of the tests supported existence of non-stationarity; rather his conclusion was in favour of a significantly negative deterministic trend over the period 1921-88. This is also the conclusion of Barros and Amazonas (1993).

In the Sapsford-Sarkar-Singer (1992) study, the starting date of 1930 was used to avoid the implicit inclusion of 1921 observation via the lagged difference terms. As Sapsford (1985) found some evidence of parametric shift after 1950, Perron test was applied. For the period, 1930-83, Perron's procedure yielded a test statistic equal to -5.02 which shows a clear rejection of the null hypothesis of non-stationarity. Hence trend-stationary model was applied over the two periods, 1922-83 and 1922-86. In each case, they found the existence of a negative and significant trend, accompanied by a significant upward intercept movement in 1950. Thus serious doubts were expressed against the conclusion of Cuddington-Urzua regarding both the absence of a downward trend and the absence of an upward intercept displacement in the Grilli-Yang series after 1950 (Sapsford-Sarkar-Singer, 1992).

To sum up, the Grilli - Yang series does not have any 'unit root' — it is not generated through a stochastic process. This is also admitted by Cuddington (1992). So to examine trend in the series, a trend stationary specification is appropriate as was made in the studies before Cuddington-Urzua (1989). Hence, the parameters of the trend -stationary model given in Eq. (1)

have been estimated below on the basis of the Grilli -Yang series over the period, 1900-1986:

$$1Y(t) = 4.97 - 0.0056\,(t) + u(t) \qquad (12)$$
$$\qquad (157.34) \quad (-8.75)$$
$$\bar{R}^2 = 0.47,\ D\text{-}W = 0.58$$

In the next step, structural instability in the series has been examined. The procedure of using dummies is arbitrary. So the CUSUM Squares test of Brown *et al.* (1975) has been used after estimating the trend-stationary model. It shows that the values of the series during 1914 - 21 are outliers. These are by and large interpolated values and as Sapsford - Sarkar - Singer (1992) showed, Cuddington-Urzua (1989) conclusion is sensitive to these observations. Therefore, all these outliers are omitted and the trend-stationary model has been reestimated :

$$1Y(t) = 4.92 - 0.0049\,(t) + u(t) \qquad (13)$$
$$\qquad (161.41) \quad (-8.28)$$
$$\bar{R}^2 = 0.46,\ D\text{-}W = 0.64$$

Box-Jenkins analysis of residuals of Eq.(13) indicates first order autocorrelation. Hence the trend-stationary model has been reestimated through Cochrane - Orcutt (C-O) iterative procedure:

$$1Y(t) = 4.97 - 0.0059\,(t) + u(t) \qquad (14)$$
$$\qquad (62.13) \quad (-4.04)$$
$$\bar{R}^2 = 0.46,\ D\text{-}W = 1.80$$

It shows a clearly significant deteriorating trend in the Grilli-Yang Series. Estimate of t- statistic of the time trend through Newey-West (1987) procedure based on autocorrelation - consistent robust covariance matrix computed for any reasonable lag also confirms this deteriorating trend (for a 4-year lag, the t - statistic has been estimated to be - 5.30).

It is interesting to note that in the Grilli-Yang Series, 1900-86 (excluding 1914-21), there is no support of the hypothesis of Sapsford (1985) and Sapsford-Sarkar-Singer (1992) regarding structural instability after 1949 :

(Estimates are obtained through the C-O procedure)[5]

$$lY(t) = 5.07 - 0.011\,(t) + 0.08\,(D\,50) + 0.003\,(D\,50.t) + u\,(t) \qquad (15)$$
$$\quad (59.56)\ (-4.28) \quad (0.33) \quad (0.66)$$
$$\bar{R}^2 = 0.57,\ F\,(3, 75) = 35.62,\ D-W = 1.81$$

where D50 = 0 during 1900-49 and = 1 during 1950 -86.

Finally, it can be examined whether excessive debt burden of the commodity producers and their consequent 'export desperation' (see Sarkar, 1991) has any influence on the long –term behaviour of terms of trade of primary products. To make this study, intercept and slope dummies are redefined : intercept dummy, D80 = 0 during 1900-79 (excluding 1914-21) and = 1 during 1980-86 ; slope dummy, D80.t varies accordingly. In view of first order autocorrelation in the residuals, the equation is estimated through the C–O procedure:

$$lY\,(t) = 4.94 - 0.0051(t) + 5.35\,(D\,80) - 0.06\,(D\,80.t) + u\,(t) \qquad (16)$$
$$\quad (63.69)\ (-3.28) \quad (2.46) \quad (-2.45)$$
$$\bar{R}^2 = 0.49,\ F(3,75) = 25.85,\ D\text{-}W = 1.81$$

The estimates show that the Grilli - Yang series exhibited a declining trend over the period, 1900-79 (excluding 1914-21) and the trend decline accentuated during 1980-86 perhaps under the influence of debt crisis and export desperation (see also Bleaney and Greenaway, 1993)[6].

III

From our critical review of the recent hi-tech debate, it appears that the Grilli-Yang series exhibited a statistically significant declining trend over the whole period, 1900-86 (excluding the outlying observations 1914-21). The annual average trend rate of decline is around –0.6 per cent (See also Ardeni and Wright, 1992)[7]. Eruption of debt crisis in the 1980s accentuated this trend decline. Since this trend is deterministic, it follows from some fundamental forces inherent in the global system. It is the task of future research to examine these forces (for theoretical issues, see Singer, 1987; Sarkar, 1993; Barros *et al*, 1993). This study may be facilitated by an examination of disaggregated terms of trade series as attempted by Cuddington (1992) and Bleaney and Greenaway (1993). Such an examination is beyond the scope of the present paper.

NOTES

1 The classical belief follows from the two laws — the operation of the law of diminishing returns in primary production and the law of increasing returns in manufactures.

2 The basic concern of both Prebisch and Singer was a secular decline in the factoral terms of trade of less developed countries (LDC s) which was inferred from the secular decline in the terms of trade of primary products vis-a-vis manufactures (see Sarkar-Singer, 1991; Sapsford-Sarkar-Singer, 1992).

3 Initially it was thought that exports of primary products

for imports of manufactures was the root of the trouble. Gradually their emphasis shifted from relations between *commodities* to relations between *countries* (Singer, 1975; Sarkar-Singer, 1991). Whatever commodities are exported by the LDCs to the developed countries, the LDCs will face some decline in their terms of trade (see Singer, 1987; Sarkar-Singer, 1991,1993 and Sarkar, 1993).

4. In Dicky-Fuller (DF) test, the following equation is fitted to examine non-stationarity against the alternative of trend-stationarity :

$$d\,1\,Y(t) = a + b.\,t + df\,.\,1\,Y(t-1) + u(t)$$

It is to be tested whether $df = 0$ against the alternative $df < 0$. If $df = 0$, it implies that the residuals of Eq. (1) are non-stationary (the series is said to have a 'unit root').

In Augmented Dicky- Fuller (ADF) test, the residuals of the above equation are made 'white noise' by adding sufficient number of lagged values of the dependent variable d 1 Y (t) as regressors :

$$d\,1\,Y(t) = a + b.t + df.1\,Y(t-1) + \sum_{i=1}^{k} df_i\,.\,d\,1\,Y(t-i) + u(t)$$

where k is sufficiently large to ensure that u (t) is 'white noise'.

Here, also $df = 0$ is tested against the alternative that $df < 0$.

5. Box-Jenkins procedure confirms first order autocorrelation in the residuals. The OLS and Newey-West (1987) procedures do not tell a different story.

6. Bleaney-Greenaway (1993) updated the G - Y series to

include the years, 1987-91 and found a similar impact of the debt crisis over the period, 1980-91.

7. Ardeni & Wright (1992) applied a different approach – the structural time series approach which does not rest on a prior assumption regarding the stationarity of the underlying data generating process or its first differences. The trend, the cycle and the residual components are explicitly modelled as unobserved components that are independent and statistically uncorrelated. This approach is not yet much known.

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