Impact of the 1997 Crisis on Korea's Growth Trend: Unobserved Component Model Based Analysis

Chan-Guk Huh and Won-Am Park*

We offer a systematic examination of the extent of permanence of the adverse influence of the 1997 crisis and its aftermath on Koreas potential GDP trend by employing unobserved component models that decompose observed output into a stochastic trend and a stationary cycle. We consider models that allow a nonzero covariance between trend and cycle innovations. Results from several unobserved component models indicate that long-term trend growth rate has shifted downward in the post-1997 crisis period as the event seemed to have left some permanent adverse impact. At the same time, output gaps measured using output-Phillips curve models have remain negative in the same period. Results seem to suggest that it is premature to make such pronouncements that the Korea economy has now entered an era of low growth, say, an annual growth rate in the 4% range.

Keywords: Potential output, Output gap, Korean economy, 1997 crisis

JEL Classification: E23

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[**Seoul Journal of Economics** 2003, Vol. 16, No. 2]

I. Introduction

Interest in actually measuring potential GDP grew as Keynesian macroeconomic views of fine tuning aggregate demand to close gaps between potential and actual outputs gained currency in the U.S. and elsewhere until the 1980s (Okun 1962; and Gordon 1984). Even though macroeconomic thought has changed considerably regarding the efficacy of aggregate demand management policies since that time, those gaps are still understood to have a first order positive influence over movements in inflation. As such, interest in potential GDP persists, especially among central banks as well as macroeconomists. For example, measuring GDP gap accurately has become more important with the advent of the practice by central banks of inflation targeting in recent years. Potential GDP has gained importance in the area of fiscal policy as more governments try to extend policy horizons for fiscal plans. For example, a better projection of potential GDP path would help reduce uncertainty in revenue growth assumptions for out years.[1]

A better understanding of potential GDP has more practical relevance in Korea in the post-1997 crisis period, as economists and policy-makers are still unsure of the extent of permanence of the adverse influence of the crisis and its aftermath. Despite the quick rebound in 1999 and 2000, the experience of the 1997 financial crisis created a distinctively pessimistic assessment of growth potential of the economy by policymakers and economists. A key problem with such a view is that there is little coherence between a positive assessment (it might be difficult to grow at the pace seen in the pre-1997 crisis period, 7.2% per year for 1990-7), and normative implications; targeting more than 5% growth per year, for example, will have ruinous consequences for the economy. Perhaps that might be the case. But a more systematic examination appears to be in order before making such a judgment with potentially serious repercussions. We offer such an investigation in this paper. What evidence can one draw from data? How confidently can one make such a claim in this regard, one way or another?

There have been efforts to gauge potential GDP in Korea. Some

[1]Some examples are; OECD (1994), St-Amant and Norden (1997), Laxton, Isard, and Eliasson (1998).

followed a method based on estimating aggregate production function, a popular methodology in the U.S, in the late 1970s.[2] This approach, in spite of its theoretical appeal, however, has to deal with difficult measurement issues such as labor and capital, their utilization rates as well as productivity and growth trends. Thus, it might be more suitable to gauge a long-term trend and less helpful in assessing the impact of a single event, such as the 1997 financial crisis, which did not affect the aggregate supplies of extant labor and capital as well as technology.

A more common approach is to use econometric detrending methods to extract a trend using either univariate or multivariate models involving real GDP series. This category includes variety of techniques. At one end is the agnostic approach of applying the Hodrick-Prescott (HP) filter to the real GDP; at the other end is univariate unobserved component models developed by Watson (1986) and Clark (1989), and some variations thereof, such as bivariate models incorporating Okuns Law (*e.g.*, Clark (1989)) or Phillips curve (*e.g.*, Kuttner (1994)). These models offer a way to decompose real GDP series into a stochastic trend and stationary cyclical components. As our main interest lies in examining the growth trend in Korea, the unobserved components (UC) model seems to offer a suitable framework. Also, a considerable body of research on unobserved components models has expanded our understanding about these models. Hence, we follow this approach.[3]

We start with an univariate UC model of Korean real GDP to examine and compare the growth trend before and after 1997 using quarterly data from 1970 to 2003. Following a common practice, we specify a random walk with a drift for trend component and a stationary AR(2) for cyclical component, or output gap. Estimation results show a lowering of unconditional long run growth rate from 7-8% to 6-7% when the post-1997 crisis observations are added. Actual trend growth, which is the sum of the drift and past and present shocks, has distinctively slowed in recent years. A drop in

[2]See Perry (1977), Clark (1979), Perloff and Wachter (1979) for the U.S., and Kim and Kim (1992), Chang (1996), Kim and Moon (2000), Pyo (2000), Hahn *et al.* (2003) for Korea (written in Korean).

[3]Some examples from the pre-1997 period are Lee (1996), Choi (1996), Chang (1996), all written in Korean, and examples of studies that include the post-1997 period are Chang (1997), Kim and Moon (2000, in Korean), Gerlach and Yiu (2002), Cerra and Saxena (2003).

trend output accounts for most of the fall in output that occurred in 1998; relatively little is accounted for by a widening of output gap. Indeed, the crisis is taken to have been a large, negative permanent shock. Compared to the pre-crisis period, variance of the trend innovation term increases sharply when post-1997 observations are added.

However, we find strong indications of instability in data as trend and cycle properties change noticeably depending on the sample periods. For example, AR(2) becomes more significant and the correlation between the shocks to trend and cycle component becomes statistically significant only in the sample that includes the post-1997 data. Morley, Nelson, and Zivot (2003) showed that different treatment of this correlation in the model setup has serious implications on the output gap estimates by economy. In addition, Proietti (2002) found that negatively correlated shocks imply large revisions to trend-cycle decomposition as new data becomes available. Indeed, the aftermath of the 1997 financial crisis has not only left in its wake uncertain adverse economic effects, but has also made efforts to assess them difficult by making data less amenable to simple models.[4]

We then proceed to examine whether a more stable decomposition might emerge when additional explanatory variables are added. In particular, we examine two variations. First, we add a Phillips curve and jointly estimate a bivariate model as in Kuttner (1994). This is intended to obtain a more stable output gap, and through this trend component, by providing an additional source of identification. Secondly, we add investment to the model. We hope to gauge more directly potential effects of a noticeable slowdown in investment spending on growth and cyclical trend. To some observers, both domestic and foreign, over-investment in the period leading up to 1997 was an important catalyst for the onset of the crisis. An apparent under-investment in the post crisis period could then potentially carry important implications in Korea's growth

[4]This does not necessarily make atheoretical decomposition methods such as HP filter more useful. Changes in cyclical properties such as duration of business cycles before and after the 1997 event would make adjustment to the parameter specifications in the HP filter necessary for it to remain as a useful detrending tool. Furthermore, an increase in output variability would bring about more frequent revisions of estimated trend paths thus worsening the end-point problem for any fixed parameter values in practice.

potential.

Patterns in trend and cycles from the bivariate models are more stable compared to the univariate cases. Results indicate that, indeed, the addition of the Phillips curve helps to achieve a more sensible identification of output gap. The coefficient of the output gap (*e.g.*, measured as actual minus trend output) in the Phillips equation remains significantly positive in all specifications.

In terms of the output equation, unconditional long run growth rate estimates are generally higher in the bivariate models compared to the univariate models. However, estimates fall from around 8% to 7% (both annualized) when the post-1997 crisis observations are added, as seen in the univariate model results. In terms of parceling the output collapse of 1998 into changes in trend *vs.* cyclical components, the bivariate models regard it as more of a cyclical contraction rather than a permanent shift in trend growth, as compared to the univariate results. In bivariate models, widening output gap accounts for about half the output contraction. Trend innovations become more variable when the post-crisis observations were added in the bivariate models, although not to the same extent seen in the univariate cases. We take this as evidence of strengthened stability of bivariate models.

There is potentially interesting additional source of cross equation dynamics in the bivariate model in addition to the covariance between the trend and cycle innovations. It is noteworthy that estimated covariance between inflation innovations and output trend innovations are statistically significantly positive, whereas the same between inflation innovations and cycle innovations are significantly negative.

Considering the trend in investment explicitly appears to help somewhat identifying growth trends and cycle patterns. For example, the lowering of trend growth is most noticeable with this specification. A natural corollary of this finding is that behavior of investment in the near term is likely to have a first order effect on the future output growth trend.

Results from several unobserved component models indicate that long-term trend growth rate has shifted downward in the post-1997 crisis period as the event seemed to have left some permanent adverse impact. For instance, the average growth rate of the trend output for the recent period has been lowered to around 6%. Such a pattern of trend output growth has been accompanied by a

persistent negative output gap since the 1997 crisis. Our results should raise a caution in making hasty assessments regarding Korea's growth potential in the near future. In particular, it is premature to make such pronouncements that the Korean economy has now entered an era of low growth, say, an annual growth rate in the 4% range. The downside risk of such a position leading to a self-fulfilling prophecy is considerable.

Further research would be required to refine focus on the shift in the dynamic pattern of the unconditional long-term growth rate in the UC framework in the post-crisis period. It might be useful to adopt a specification that allows more variation to the long-term growth rate itself.[5] The remainder of the paper is organized as follows: Section II offers a brief overview of output, inflation, investment and employment time series data; Section III presents univariate unobserved component models and estimation results; Section IV presents both an output-Phillips curve model as well as output-Phillips curve-investment models and their estimation results; and Section V presents our conclusions.

II. A Preliminary Look at Data

We first offer a cursory overview of some key macroeconomic data of real GDP, investment, employment, and inflation. Both log level and year-on-year growth rate of real GDP (measured in 1995 Korean won) for 1970Q1-2003Q2 are shown in Figure 1. Two points are noteworthy. The fall in real GDP in 1998 was indeed severe. Except for 1980, when real GDP fell 2% in the wake of the second oil shock, the 1998 contraction in real output was quite remarkable. Second, growth in real GDP appears to have become more volatile since 1998.

Figure 2 shows year-on-year growth rates of the consumer price index and import prices for the same piriod. Inflation in both price indices was quite high until the early 1980s. Their patterns, however, have diverged ever since. Inflation in import prices (measured in Korean won) remained high in the beginning of the 1980s then started to fall and turn negative in early the 1990s.

[5]We thank Chang-Jin Kim for suggestions regarding this and related issues.

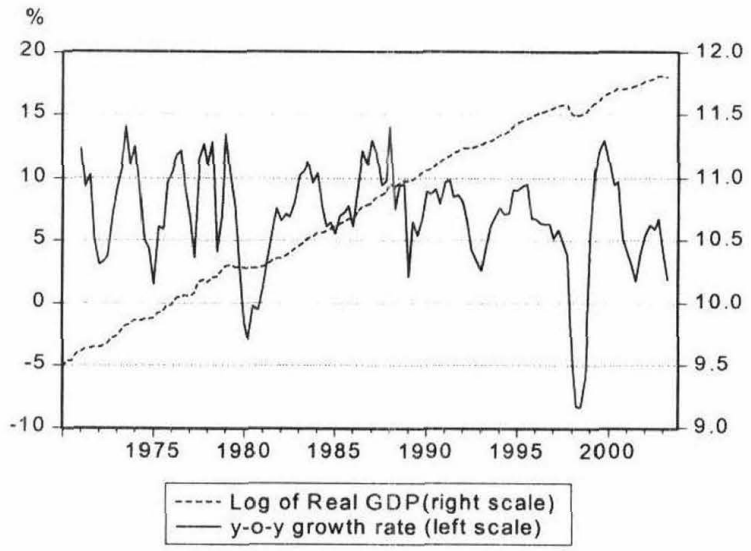

FIGURE 1

REAL GDP AND GROWTH RATE (1970Q1-2003Q2)

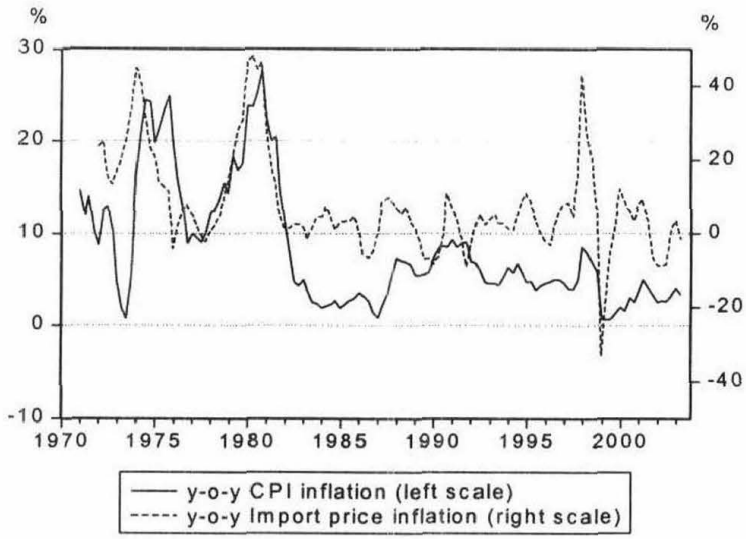

FIGURE 2

INFLATION IN CPI AND IMPORT PRICES (1971Q1-2003Q2)

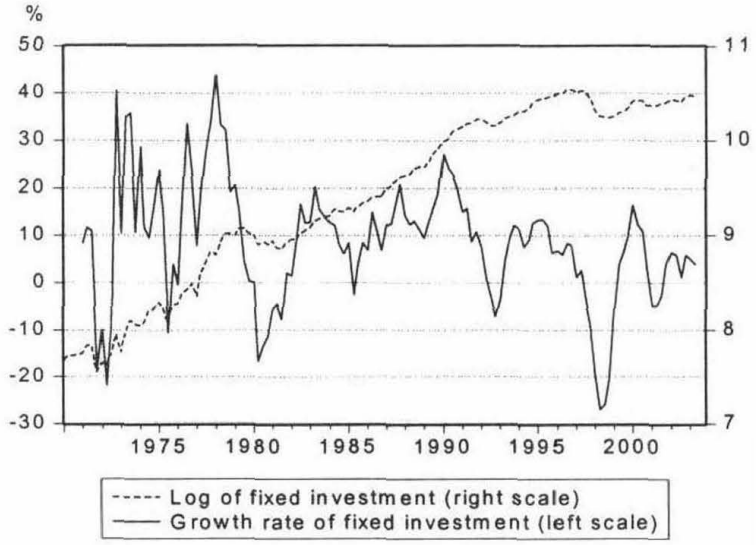

FIGURE 3

FIXED INVESTMENT (1970Q1-2003Q2)

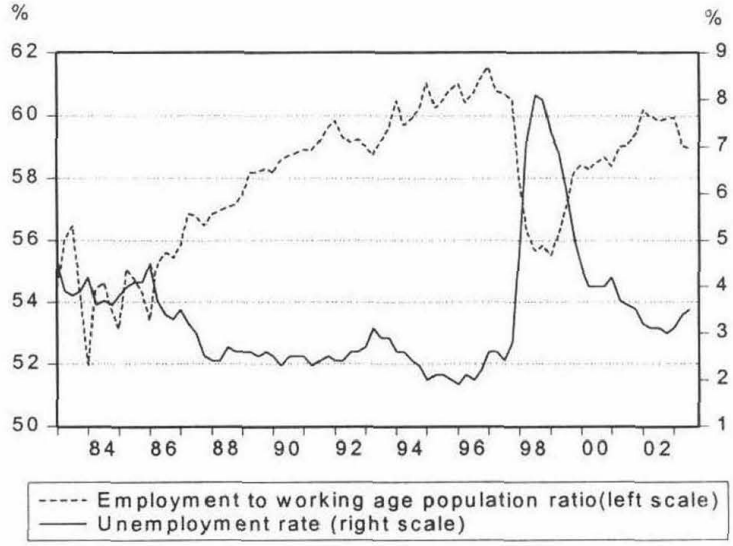

FIGURE 4

EMPLOYMENT AND UNEMPLOYMENT (1981Q1-2003Q2)

CPI inflation remained relatively high throughout most of the 1990s with a distinct peak in 1998. A sharp rise in import price inflation explains this episode that was caused by a pronounced depreciation of the Korean won at that time. However, inflation in both prices has become relatively subdued in recent years.

Figure 3 shows log levels and year-on-year growth rates of fixed investment series (*i.e.*, total investment net of inventory investment, measured in 1995 Korean won) for. On average, capital spending has been strong, howbeit rather erratic. Investment has been particularly variable since the early 1990s.

Figure 4 presents both the unemployment rate and employment-to-working age population ratio for the sample period, as determined by data availability, of 1983 to 2003Q2. Relatively little variation is evident in the unemployment rate with the exception of a sharp rise in 1998. This pattern casts doubt on the unemployment rate as a reliable source of information regarding business cycles in Korea. On the other hand, the employment ratio has been far more variable and it reveals a level shift since 1998. The ratio falls by more than 2%, indicating that approximately seven hundred thousand potential workers, who used to be employed before 1998, are now out of work.

Finally, Figure 5 shows the contemporaneous business cycle index compiled by the National Statistics Office and the official business cycle dating from 1970. One noticeable feature of these cycles is the lack of a close correspondence actual real GDP growth pattern as shown in Figure 1. The average duration of expansions and contractions of business cycles, so defined, until 1998 were 34 and 19 months, respectively (Table 1). Changes in cyclical properties since the 1997 crisis have been of such significance as to delay official business cycle dating.[6]

[6]The post-crisis cyclical changes like these potentially present problems to a widely used detrending method of HP filter. The key smoothing parameter value of 1,600 typically used is chosen to draw out cyclical features from business cycles with a particular range of durations (about 4 years). Thus, one might obtain misleading inferences when dealing with data that do not share such a property. See King and Rebelo (1993) for more discussion about the optimality of the HP filter.

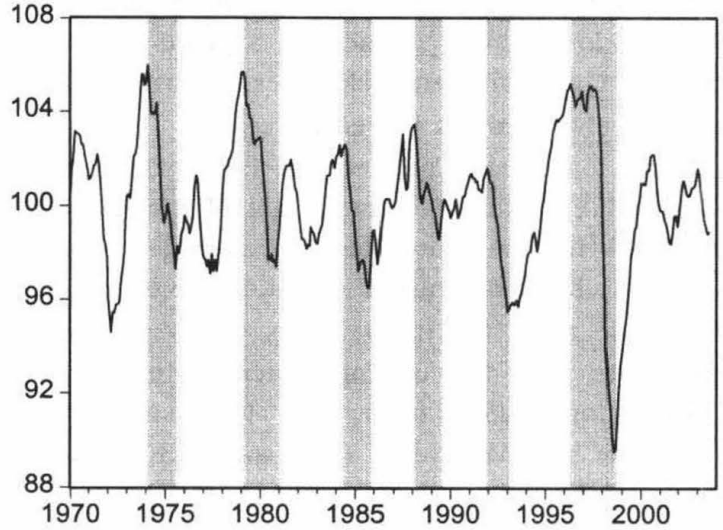

FIGURE 5

NATIONAL STATISTICS OFFICE BUSINESS CYCLES (1970Q1-1998Q4)

TABLE 1

BUSINESS CYCLE DATES BY NATIONAL STATISTICS OFFICE

	Turning point dates		Duration (in months)	
	Trough	Peak	Expansion	Contraction
1st	72.3	74.2	23	16
2nd	75.6	79.2	44	19
3rd	80.9	84.2	41	19
4th	85.9	88.1	28	18
5th	89.7	92.1	30	12
6th	93.1	96.3	38	29
7th	98.8	NA	NA	NA
Average			34	19

III. Univariate Unobserved Component Model of Real GDP

We start with an univariate unobserved components (UC) ARIMA model of Korean real GDP. One important appeal of this approach is that it allows one to decompose an integrated series into trend

and cycle without imposing artificial smoothness a priori, as done in detrending methods such as the HP filter.

The UC model examined in this paper has the following structure. Observed output (y_t), which is integrated with order 1 according to a preliminary examination, is decomposed into two unobserved components of a stochastic trend (y_t^p) and a stationary transitory deviation from trend, or cycle (c_t),

$$y_t = y_t^p + c_t \tag{1}$$

and the trend and cycle components have following specifications.

$$y_t^p = \mu + y_{t-1}^p + \nu_t \tag{2}$$

$$c_t = \phi_1 c_{t-1} + \phi_2 c_{t-2} + \omega_t \tag{3}$$

$$\nu \sim N(0, \sigma_\nu^2), \quad \omega \sim N(0, \sigma_\omega^2), \text{ and cov } (\nu_t, \omega_t) = \sigma_{\nu\omega}.$$

Seasonally adjusted quarterly GDP measured in 1995 prices from 1970Q1 to 2003Q2 was used as the output series. y_t is the log of output multiplied by 100, so output gap can be thought of as a percentage deviation from trend.

A number of papers have applied this method, starting from Harvey (1985) and Watson (1986). Clark (1989) specifies the drift term as time varying to analyze the trend-cycle properties of 6 advanced economies. The UC model is recast in a state space form and estimated by the maximum likelihood method with Kalman filtering.[7]

One important finding by Morley, Nelson, and Zivot (2003) is that a very smooth trend and a highly persistent cycle with large amplitude usually obtained from the UC models, might be due to the restriction of zero correlation between the innovations of trend and cycle. They show that the covariance between ν and ω can be identified in the class of ARIMA(2,1,0) models, and when it was actually estimated using the U.S., the correlation was close to minus one. The cycles implied by this unrestricted model is distinctly different from that of the model with zero covariance

[7]Estimations were done using GAUSS programs. Data as well as a technical appendix that offers more detailed description of the state space representation will be available upon request.

restriction in that it lacks any discernable pattern and noisy, as in the case of the Beverage-Nelson decomposition (Beverage and Nelson 1981).

Setting aside methodological issues, this finding has potentially important implications on how one interprets the events of the 1997 crisis. The 1998 fall-off in real GDP of 6.7% was indeed significant (more so if one accounts for the fact that real GDP grew 5% in the previous year). The model somehow has to determine how much of the large drop in output represents a permanent shock (*i.e.* a trend change) and how much is due to a temporary shock (*i.e.* a cyclical change). Removing the zero restriction between the two types of shocks thus could influence how much of the fall is attributed to each source.

If one views that the contraction in 1998 represents a permanent change, then the Korean economy indeed has experienced a serious structural break. We could then easily discount the growth trend in the pre-crisis period in calculating the output gap for a more recent post-crisis period. On the other hand, if one adopts the view that the drop in output of 1998 reflects a fair number of temporary factors, then it becomes necessary to take the 1998 episode into consideration when calculating output gap for the subsequent period. This is a very difficult question to answer even with enough time series data, which we yet do not have. However, many observers who are making the following inferences are answering the above question implicitly: the Korean economy has now entered an era of low growth since the onset of the 1997 crisis based on real GDP growth trend observed in the wake of the 1997 crisis.

Estimation results are shown in Table 2. Results are grouped into two sample periods, and within each sample, according to the imposition of the zero covariance restriction. The three panels of Figure 6 show trend and cycle decompositions for the pre-1997 sample with the zero covariance restriction in addition to the year-on-year growth rate of the trend output component (the middle panel). Figures 7 and 8 show the same from the full sample (*i.e.*, with post-crisis observations), respectively, with and without the zero covariance restriction. Several points are noteworthy.

First, trend output growth is lowered noticeably when the post-crisis observations are added. In terms of the unconditional long-run growth rate (μ), it falls from 7.5% (1.888×4) to 6.9% (1.730×4) in the full sample. Actual trend output appears to have

TABLE 2

RESULTS FROM UNIVARIATE UNOBSERVED COMPONENT MODEL

Sample period	1970Q1-1997Q4		1970Q1-2003Q2	
Model	$cov(\nu, \omega)=0$	$cov(\nu, \omega)\neq 0$	$cov(\nu, \omega)=0$	$cov(\nu, \omega)\neq 0$
μ	1.888**	1.888**	1.736**	1.730**
ϕ_1	0.849**	0.849**	1.647**	0.042**
ϕ_2	0.048	0.048	-0.774**	-0.772**
σ_ν	0.0003	0.0077	1.797**	2.083**
σ_ω	1.746**	1.745**	0.425**	0.193**
$cov(\nu, \omega)$		0.003		-0.402**
$\rho(\nu, \omega)$		0.223		-0.999**
Likelihood	-220.36	-220.36	-274.52	-273.52

Note: ** denotes cases that are significant at 1%.

slowed distinctly in recent years. According to Figure 7, the year-on-year growth rate has fallen to as low as 1.5% in the most recent period (2003Q2) from a recent peak of 6.8% seen in 2002Q4. However, Figure 7 is almost indistinguishable from Figure 1, which shows the actual growth rate of real GDP. The trough of 1998Q3 (when real GDP fell 8.1%), which is more or less fully reflected by a similar drop in the trend output (y^p fell 7.6%) in the same quarter. Indeed, the crisis is taken to have been a large, negative, and permanent shock.

Second, variance of the trend innovation term increases sharply when post-crisis observations are added. In terms of the ratio of the estimated standard errors of trend and cycle innovations (that is, σ_ν/σ_ω), it shifts from the pre-crisis period of 0.0003/1.746 to 1.797/0.425 in the post-crisis sample for the restricted UC. A similarly drastic shift is observed in the unrestricted UC model.[8] At the same time, a sizable increase in the standard error of the trend innovation makes the trend component more volatile and leaves relatively little for the cycle component and its innovation (σ_ω). The financial crisis event and the consequent fall-off in output have increased uncertainty about the trend.

[8]This is an illustration of how extreme the contraction resulting from the 1997 crisis was. Intuitively speaking, the Kalman filter updating method was forced to drastically change estimates of the standard error of the trend term (σ_ν) when faced with a large fall-off of the level of real GDP, say, in 1998Q1 to improve the fit of the model.

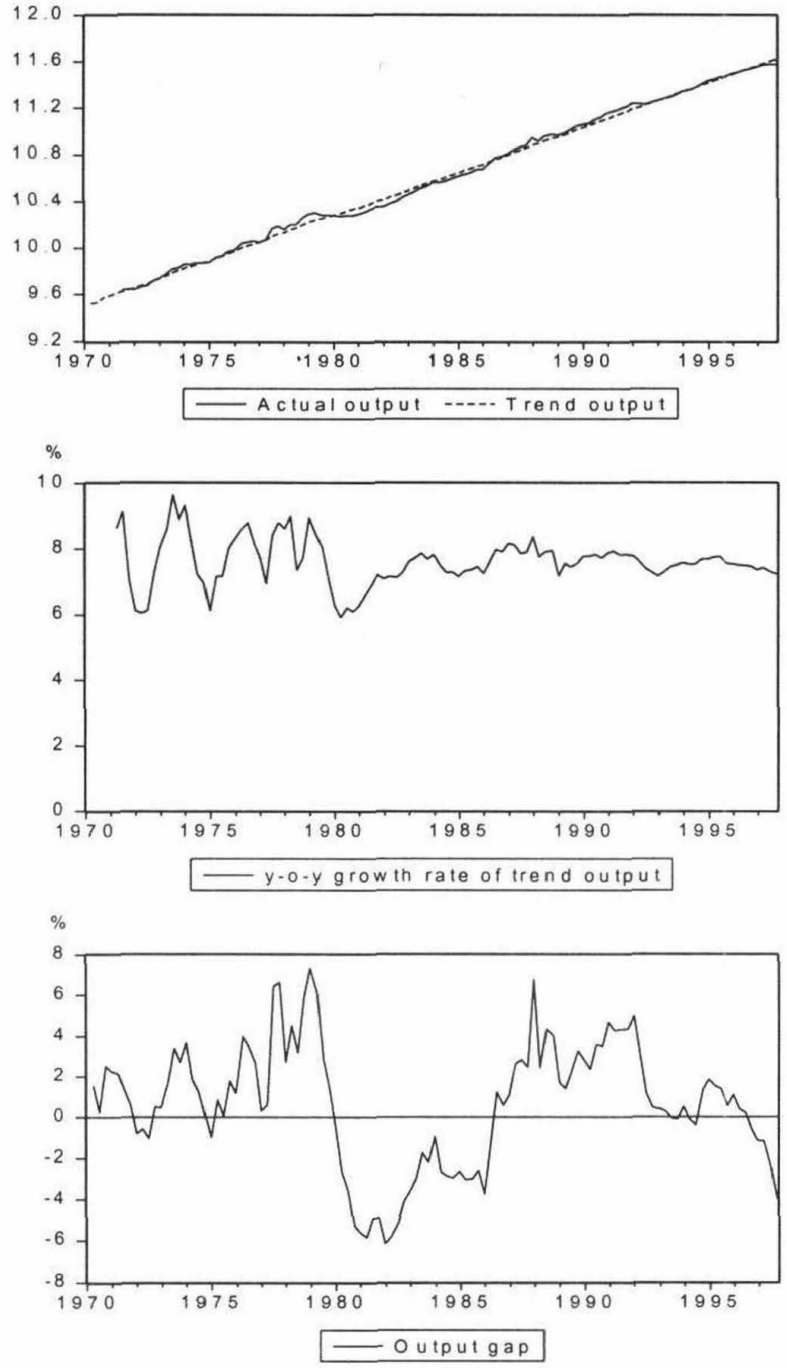

FIGURE 6

TREND OUTPUT AND GAP FROM RESTRICTED UNIVARIATE UC MODEL (1970Q1-1997Q4)

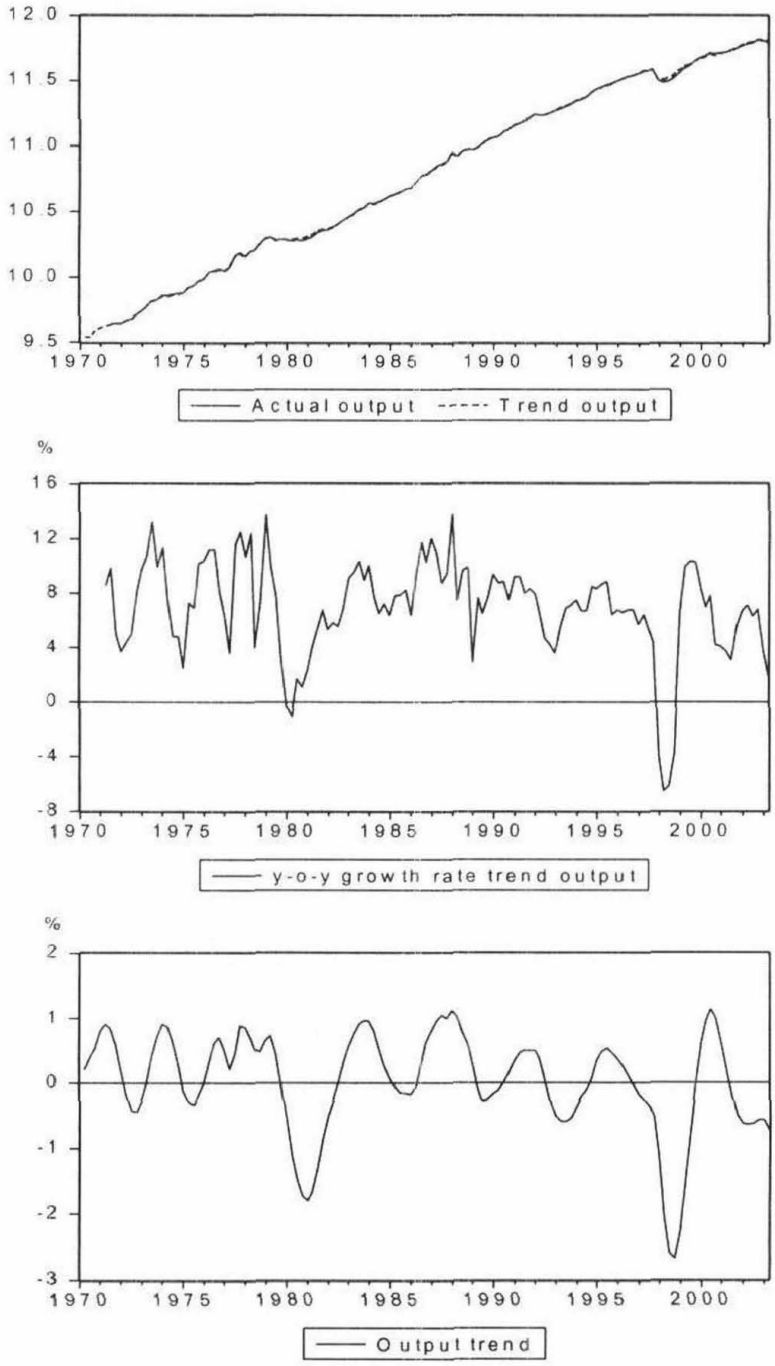

FIGURE 7

TREND OUTPUT AND GAP FROM RESTRICTED UNIVARIATE UC MODEL (1970Q1-2003Q2)

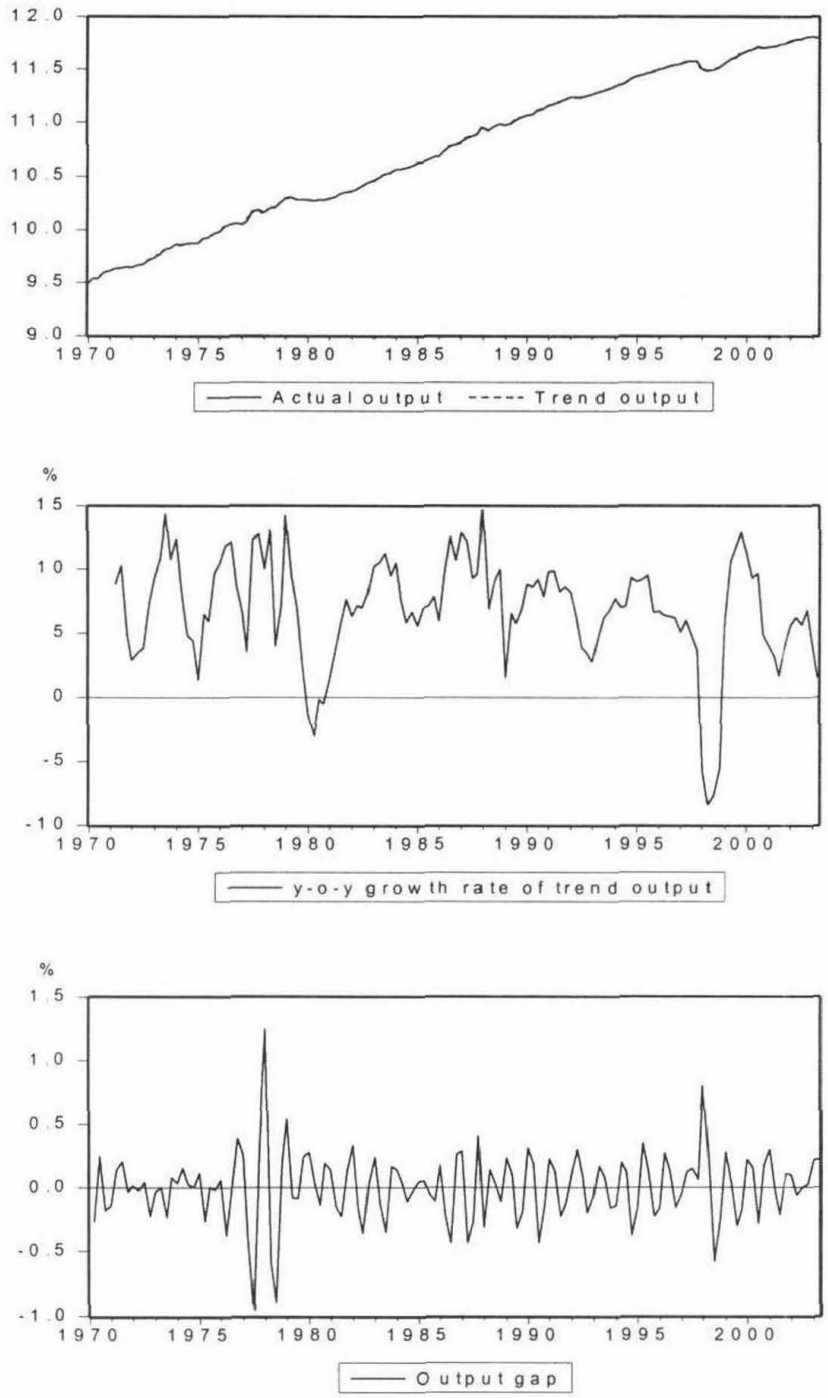

FIGURE 8

TREND OUTPUT AND GAP FROM UNRESTRICTED UNIVARIATE UC MODEL (1970Q1-2003Q2)

Third, the cyclical properties are expected to differ across the two samples considering that only in the full sample are the roots of the characteristic equations of AR(2) coefficients complex. Thus, a more periodic behavior is expected. Also the zero covariance restriction has serious consequences only in the full sample. Cyclical patterns of the two cases for the pre-crisis sample are identical (results from the unrestricted model are not shown). Cyclical patterns of models with and without the zero covariance restriction are distinctly different when Figures 7 and 8 are compared. The cyclical pattern in the model with zero restriction has a larger amplitude than that from the one without the restriction and less noisy.[9] It is noteworthy that the cycle component from the unrestricted UC model appears quite noisy as in the BN decomposition and unrestricted UC AR(2) model for the U.S. shown in Morley, Nelson, and Zivot (2003). Differences between the conventionally defined cycle, such as the one shown in Figure 6, and this are very different in terms of duration and persistence. Perhaps one factor that could limit the usefulness of the unrestricted univariate UC decomposition is the apparently large revision of the past caused by newly available observations. This result seems to confirm a finding by Proietti (2002) that a negative correlation implies that future observations carry most of the information needed to assess cyclical stance in UC models with correlated trend-cycle components.

In general, univariate decomposition results seem unstable in the full sample. Trend components appear excessively variable and cyclical components shift quite drastically depending on sample periods as well as different treatments of the covariance between the trend and cycle innovation. One interpretation of these results is that the behavior of output in the post-crisis period was quite extraordinary to introduce instability in results obtained from a well-developed univariate trend-cycle decomposition model. To examine whether this observation can be mitigated once we move away from the univariate framework, we turn to adding more variables to this univariate setup.

[9]However, the cycles from the restricted model appears rather smooth and artificial. These characteristics are also found in Gerlach and Yiu (2002)s estimates using Korean data for 1973Q1-2001Q1 sample. However, it is interesting to note that Gerlach and Yiu add a dummy variable for 1998Q1 to account for the aftermath of the financial crisis in their estimation.

IV. Bivariate Models

A. Output-Phillips Curve Model

We expand the UC model by first adding a second equation of inflation, or so called Phillips curve, as in Kuttner (1994). We hope to achieve a better identification of cycle component in the decomposition by exploiting an apparent explanatory power of the short-run Phillips curve. We proceed by finding the Korean CPI series to be I(1). Then we found an AR(4) to be most suitable autoregressive structure for the differenced logged CPI for the sample period of 1971Q1-2003Q2.[10] Based on that, we specify the Phillips curve as follows:

$$\pi_t = \sum \Psi_i \pi_{t-1} + \alpha X + \eta_t \tag{4}$$

where, π represents demeaned quarterly inflation rate in CPI, η_t, a white noise term, and X stands for other explanatory variables, α is vector of coefficients for those variables. Variables included as part of X are: lagged real GDP growth rate (Δy_{t-1}), lagged output gap (or, cyclical component, c_{t-1}), and lagged inflation in import prices ($\Delta \mathrm{imp}_{t-1}$). The reason for the inclusion of the lagged output is to determine whether the output gap would remain significant when the lagged real GDP is added. The inflation in import prices is considered here to account for a surge in CPI inflation seen in 1980, caused by a steep increase in oil prices (the second oil shock), and another uptick in inflation in 1998, which was a result of a sharp drop in Korean won's exchange value in early 1998 while output was rapidly falling. In addition, it is to account for any secular trend in CPI.

Once a bivariate specification is determined, it is converted into state space specification and estimated by applying the Kalman filter through a maximum likelihood method, as in the univariate case. Table 3 shows estimation results for the pre- and post-crisis samples under the different covariance structure of innovations

[10]The logged CPI series was found to be I(1) using the Augmented Dicky-Fuller and Phillips-Perron tests. Then the Baysian Information Criteria was used to determine the ARMA structure of the first differenced and demeaned log CPI series (π) for the 1971Q1−2003Q2 sample period. The minimum BIC was obtained for the ARMA(4, 0) specification.

TABLE 3

ESTIMATION RESULTS FROM THE OUTPUT-PHILLIPS CURVE BIVARIATE MODEL [EQUATIONS (2), (3), AND (4)]

	1971Q3-1997Q4		1971Q3-2003Q2	
Model	$\text{cov}(\nu, \omega)=0$	$\text{cov}(\nu, \omega)\neq 0$	$\text{cov}(\nu, \omega)=0$	$\text{cov}(\nu, \omega)\neq 0$
Output equation				
μ	1.911**	1.889**	1.787**	1.827**
ϕ_1	0.883**	0.911**	1.146**	1.629**
ϕ_2	0.008	-0.034	-0.277**	-0.696**
σ_ν	0.036**	2.443**	1.526**	1.694**
σ_ω	1.723**	1.537**	1.074**	0.362**
$\text{cov}(\nu, \omega)$		-2.682**		0.266**
$\rho(\nu, \omega)$		-0.713**		0.433**
Phillips curve equation				
Ψ_1	0.248**	0.373**	0.307**	0.332**
Ψ_2	0.031	0.077	0.073	0.075
Ψ_3	0.084	0.120	0.120	0.144*
Ψ_4	0.229**	0.220**	0.288**	0.310**
α_{imp}	0.124**	0.118**	0.052*	0.048*
α_y	-0.110	-0.132**	-0.062	-0.079*
α_c	0.120**	-0.136**	0.322**	0.314**
σ_η	1.560**	1.582**	1.377**	1.516**
$\text{cov}(\nu, \eta)$	-0.056**	-2.761**	1.389**	0.162**
$\rho(\nu, \eta)$	-0.997**	-0.713**	0.661**	0.063**
$\text{cov}(\omega, \eta)$	-0.024**	2.434**	-1.481**	-0.479**
$\rho(\omega, \eta)$	-0.009**	0.999**	-0.999**	-0.872**
Log likelihood	-303.57	-305.16	-376.46	-374.79

Note: * and ** denote cases that are significant at 5% and 1%, respectively.

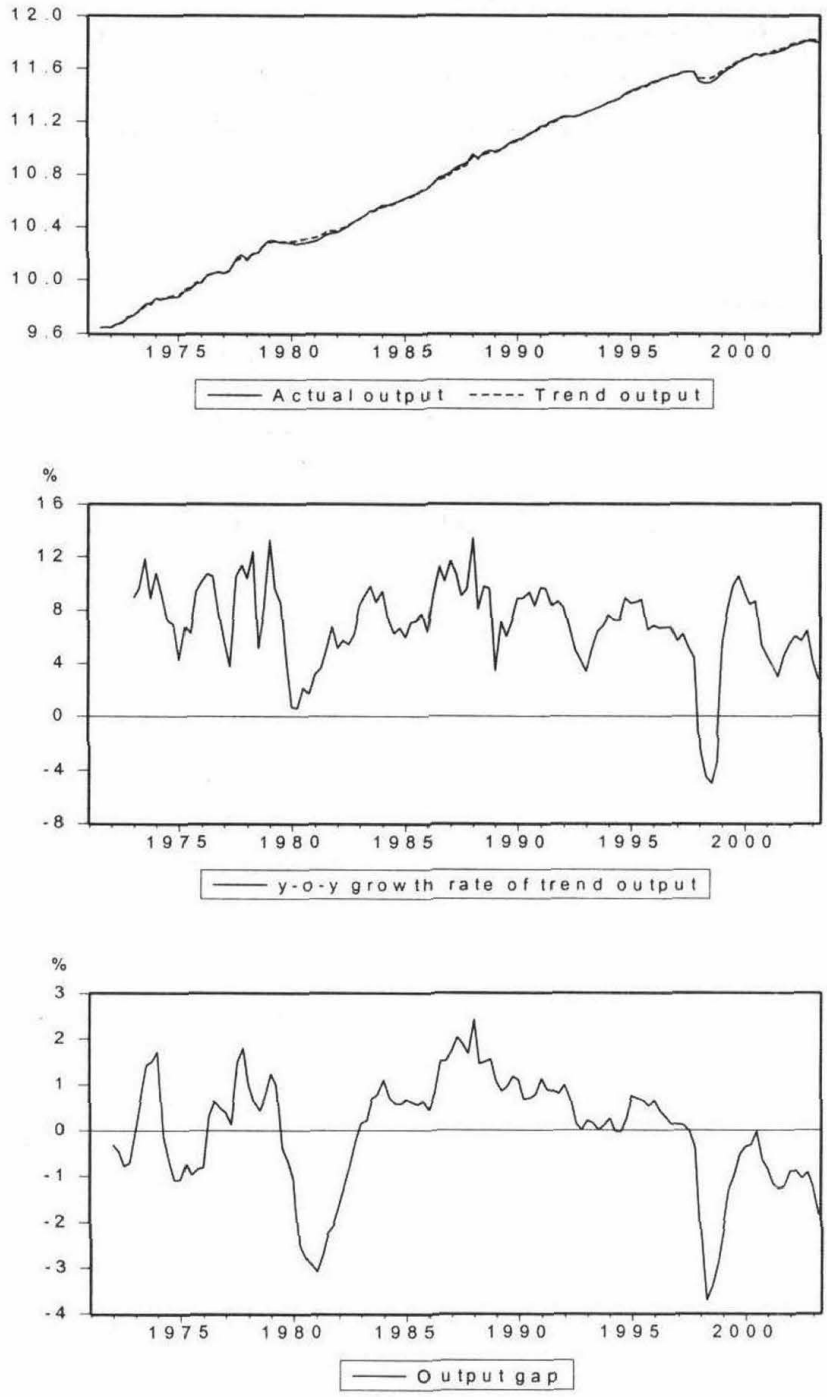

FIGURE 9

TREND OUTPUT AND GAP FROM RESTRICTED OUTPUT-PHILLIPS CURVE UC MODEL (1972Q3-2003Q2)

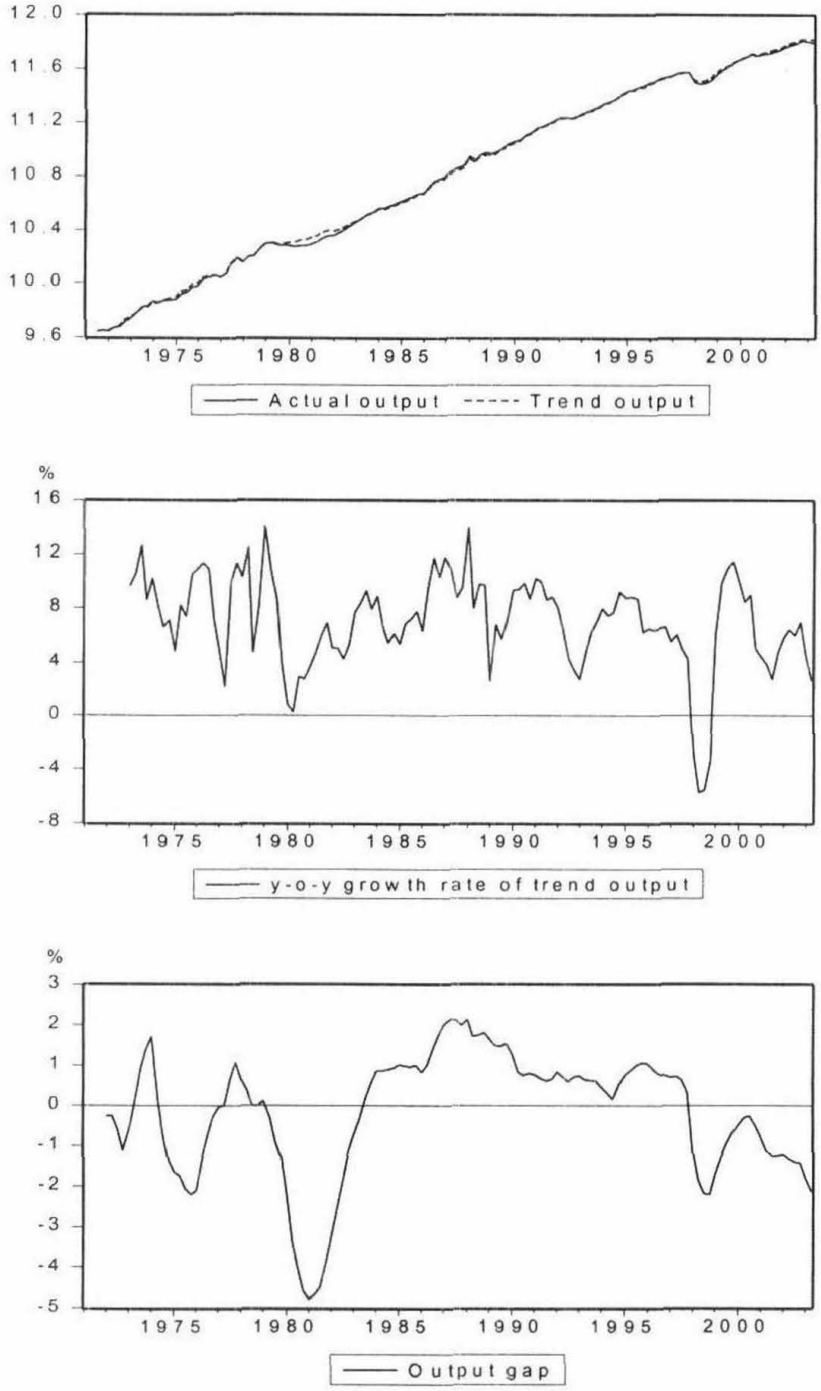

FIGURE 10

TREND OUTPUT AND GAP FROM UNRESTRICTED OUTPUT-PHILLIPS CURVE UC MODEL (1972Q3-2003Q2)

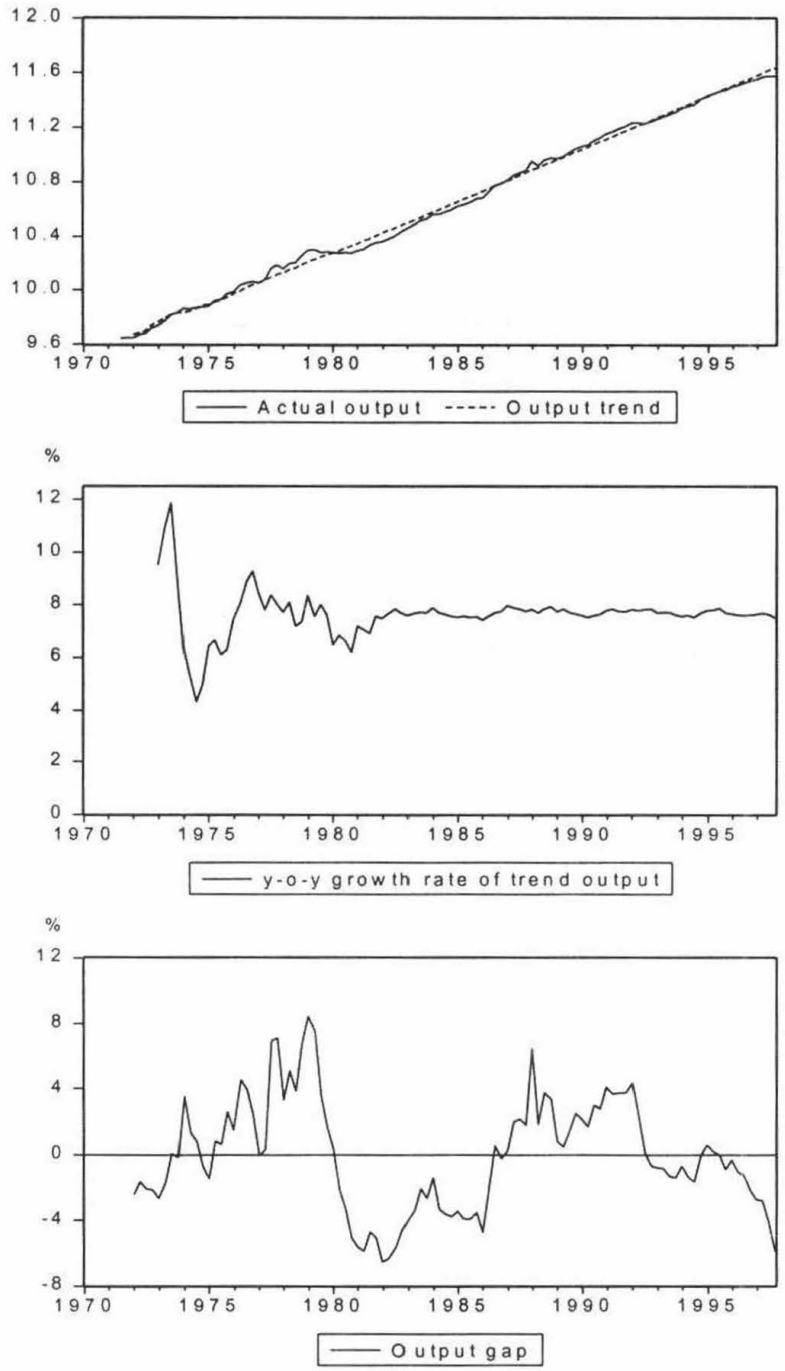

FIGURE 11

TREND OUTPUT AND GAP FROM RESTRICTED OUTPUT-PHILLIPS CURVE UC MODEL (1972Q3-1997Q4)

observations. The bivariate results shown in Table 3 do not seem to vary noticeably with respect to the inclusion of the zero covariance restriction. However, results do vary when a different covariance structure is assumed and we will discuss this later in the section. We will mainly focus on the full sample results and note any interesting discrepancies across different specifications and samples.[11]

We first turn to the Phillips curve estimation results. The output gap (α_c) is consistently significant in the Phillips equation across all four cases, whereas lagged output growth (α_y) is mostly insignificant. This seems to confirm the veracity of the decomposed output gap measure according to conventional definition of the short run Phillips curve.[12] The inflation rate in import prices (α_{imp})also remains significant in all four cases. AR coefficients of lagged inflation are significant for AR(1) and AR(4) (Ψ_1,Ψ_4)terms, the latter probably suggesting a seasonality in the CPI data.

Second, we turn to issues related to output decompositions. Figures 9 and 10 present the trend output, its year-on-year growth rate, and the output gap, respectively, estimated from the full sample model of (1)-(4) with the zero covariance restriction (Figure 9) and with no restriction (Figure 10). This particular case was most representative of four sets of different model estimation results.[13] In terms of trend output, the implied long-term growth rate (μ) is significantly lower in the full sample; it falls from 7.64% per year in the pre-crisis sample, to 7.14% in the longer sample with post crisis observations. Although the year-on-year growth rate of the trend output for this model appears quite variable, it is much less

[11]Even though not shown here, results from the model where all covariance terms were restricted to zero were more or less identical to this case.

[12]Kuttner (1994) found both lagged output growth and output gap significant for the US data for sample period from 1954 to 1992.

[13]In the full sample, we estimated two more models in addition to ones shown, that differ in terms of covariance structure: one with zero restriction on all covariance (*i.e.*, $\mathrm{cov}(\nu, \omega)=0$, $\mathrm{cov}(\nu, \eta)=0$, $\mathrm{cov}(\omega, \eta)=0$), the other, with unrestricted trend-cycle covariance and zero restrictions on cross-equation covariance (*i.e.*, $\mathrm{cov}(\nu, \omega)\neq 0$, $\mathrm{cov}(\nu, \eta)=0$, $\mathrm{cov}(\omega, \eta)=0$). Estimation results do not appear materially different from those shown here, but dynamic patterns from models not shown in Table 3 are more similar to those from the model with the zero trend-cycle covariance restriction, shown in Figure 9.

so compared to Figure 8, which showed the same from the univariate model. For example, the drop-off in the trend growth rate is not as severe. Thus, about half of the actual contraction in real GDP in 1998 of 6.7% is attributed to a widening of output gap. Figure 11 shows the same for the model estimated using only pre-crisis data for comparison purposes.[14]

With regard to the cyclical component, the AR(2) specification is generally more significant in these models. Unlike the comparable univariate case, the characteristic roots are real, thus we can expect less of a periodic pattern in the output gap series. The output gap shown in the third panel of Figure 9 appears very different from those of univariate cases (shown in Figures 7 and 8). Two troughs (1980, 1998) and a peak (late 1980s) are most visible. Since the 1997 crisis, output gap fell close to 4% in 1998 and has remained below zero. In comparison, the bivariate output gap estimated from the pre-crisis sample (Figure 11) appears nearly identical to those from univariate models (Figure 6).[15] This suggests that in the pre-crisis sample period, moving from the univariate output framework to the bivariate framework makes little difference in identifying trend-cycle decompositions.

It is interesting to note that the covariance between trend and cycle innovations turns positive in this bivariate model compared to the earlier univariate models. The fourth column of Table 3 shows that the covariance is positive and the correlation is 0.433. The same correlation was −0.999 for the univariate case shown in Table 2. This appears to be a demonstration of the lack of robustness of the negative correlation result of the univariate UC model.

The bivariate specification offers 6 potential covariance restrictions. In addition to cases shown in Table 3, 4 additional cross-equation covariance restrictions are possible as explained in the footnote 13. Different restrictions do not cause substantive changes in estimation results in general. In most cases, cycle innovation variances tend to be larger than those of trend innovations regardless of samples and the covariance restriction with the

[14]Dynamic patterns from various models appear very similar to those shown here.

[15]Estimates of output gap in Korean papers using similar UC models published before 1997 also show similar patterns (Lee 1996; and Choi 1996).

exception of the cov(ν, ω)=0 for the pre-crisis sample period (the first column of Table 3). One interesting covariance pattern reveals that innovations of the Phillips curve equation is significantly and positively correlated (*i.e.*, $\rho(\nu, \eta)>0$) with trend innovation in three out of four cases, while the correlation is negative with respect to cycle innovations (*i.e.*, $\rho(\omega, \eta)<0$). Inclusion of output gap in the Phillips curve could be the source of the latter covariance pattern. However, the former pattern is somewhat of a curious result.

There are a couple of instances where results do seem to differ when different covariance restrictions are imposed. The first case is the unrestricted model in the pre-crisis period (shown in Table 3). Variances of both trend and cycle innovations are substantially large. The output gap (not shown) behavior is also distinct from the other three cases shown in Table 3. For one, the amplitude of the output gap becomes much smaller. Basically, cycles are less visible. Next, we observe similar changes when only cross-covariance restrictions are imposed in the full sample, *i.e.*, cov(ν, η)=0, cov(ω, η)=0, with cov(ν, ω)$\neq$0. Variances of both trend and cycle innovations are about 3 to 4 times larger than the full sample unrestricted case. The output gap (not shown) gets quite noisy and the amplitude gets much smaller. Basically, no clear cyclical pattern emerges. In turn, trend output becomes much more variable. Those patterns bear some resemblance to those of the unrestricted univariate case.

B. Output-Phillips Curve-Investment Model

Throughout the past several decades, investment grew close to 25% per year in Korea (see Figure 3). Compared to such a longstanding trend, a remarkable sluggishness in investment growth is perhaps one of the most distinctive changes observed in the post-crisis period. An increase in investment has both short-run, as a part of GDP that year, and long-run, by boosting productive capacity of the economy over time, positive effects on output. Thus, taking dynamic investment behavior into account could help explain patterns in both short run cyclical as well as long-run trend components of output.

To incorporate investment in our current framework, we add an exogenous investment variable as part of the trend and cycle components alternatively while retaining the bivariate model used in

the previous section.[16] [17] To be more specific, in one case the lagged growth rate in investment (inv_{t-1}) is added to the trend equation (2), which is rewritten as (2)′

$$y_t^p = \mu + y_{t-1}^p + \phi_{inv} \Delta \, inv_{t-1} + \nu_t \qquad (2)'$$

Alternatively, it is added to the cycle equation (3), which is rewritten as (3)′

$$c_t = \phi_1 c_{t-1} + \phi_2 c_{t-2} + \phi_{inv} \Delta \, inv_{t-1} + \omega_t \qquad (3)'$$

Both fixed investment and facility investment are alternatively used. Full sample estimation results of four different cases are shown in Table 4.

A few points are noteworthy. First, investment variables are generally significant with the exception of the fixed investment in trend.

Second, the long-run trend growth rate (μ) estimates are perceptibly lower, compared to the output-Phillips curve bivariate models of Table 3. The implied annual growth rate is generally in the mid-6% range, instead of 7% as seen earlier. This lowering of the long-run trend growth estimates has some interesting implications for decomposition outcomes. For example, output gap estimates in recent year (since 2000) have been positive according to models that include one of the two investment variables as a part of the cycles. However, the gap remains negative when investment is included as a part of trend. Figure 12 shows trend output, its year-on-year growth rate, and output gap estimated from the model with fixed investment in trend specification.[18]

[16]This perhaps is not a common way of modeling investment. However, a somewhat similar approach introducing an exogenous variable in a bivariate UC model can be found in Gerlach and Smet (2002). They add a real interest rate variable in the output gap equation in their analysis of European monetary policy.

[17]We considered two investment variables: fixed as well as facility investment. The latter is a component of the fixed investment but exclude construction as well as inventory investments.

[18]In addition, there are a couple of distinct shifts in patterns that we cannot explain. Cyclical innovation variances tend to be larger than those of trend innovations. The opposite was true in estimation results of Table 3 for output-Phillips curve specifications. The sign of covariance between the

TABLE 4

ESTIMATION RESULTS FROM OUTPUT-PHILLIPS CURVE-INVESTMENT MODEL (1971Q3-2003Q2)

	Investment in trend		Investment in cycle	
Models	Facility	Fixed	Facility	Fixed
Output equation				
μ	1.664**	1.660**	1.704**	1.625**
ϕ_1	1.698**	1.682**	1.595**	1.508**
ϕ_2	-0.755**	-0.742**	-0.686**	-0.654**
ϕ_{inv}	0.033	0.041	0.040**	0.061**
σ_ν	1.724**	1.726**	1.776**	1.919**
σ_ω	0.284**	0.281**	0.295**	0.307**
cov(ν, ω)	0.127**	0.161	-0.053**	-0.255**
ρ(ν, ω)	0.260**	0.332**	-0.102**	-0.434**
Phillips curve equation				
Ψ_1	0.414**	0.413**	0.391**	0.344**
Ψ_2	0.093	0.092	0.075	0.067
Ψ_3	0.115	0.117	0.094	0.087
Ψ_4	0.322**	0.320**	0.287**	0.288**
α_y	-0.044**	-0.046**	-0.151**	-0.125**
α_c	0.331	0.358	0.241**	0.381**
σ_η	1.552**	1.549**	1.552**	1.508**
cov(ν, η)	0.179**	0.093**	0.283**	-0.110**
ρ(ν, η)	0.067**	0.034**	0.102**	-0.038**
cov(ω, η)	-0.418**	-0.406**	-0.458**	-0.215**
ρ(ω, η)	-0.945**	-0.931**	-0.999**	-0.646**
Log likelihood	-388.35	-389.02	-383.08	-385.41

trend and cycle innovations switches from positive (investment in trend specification) to negative (investment in cycle specification).

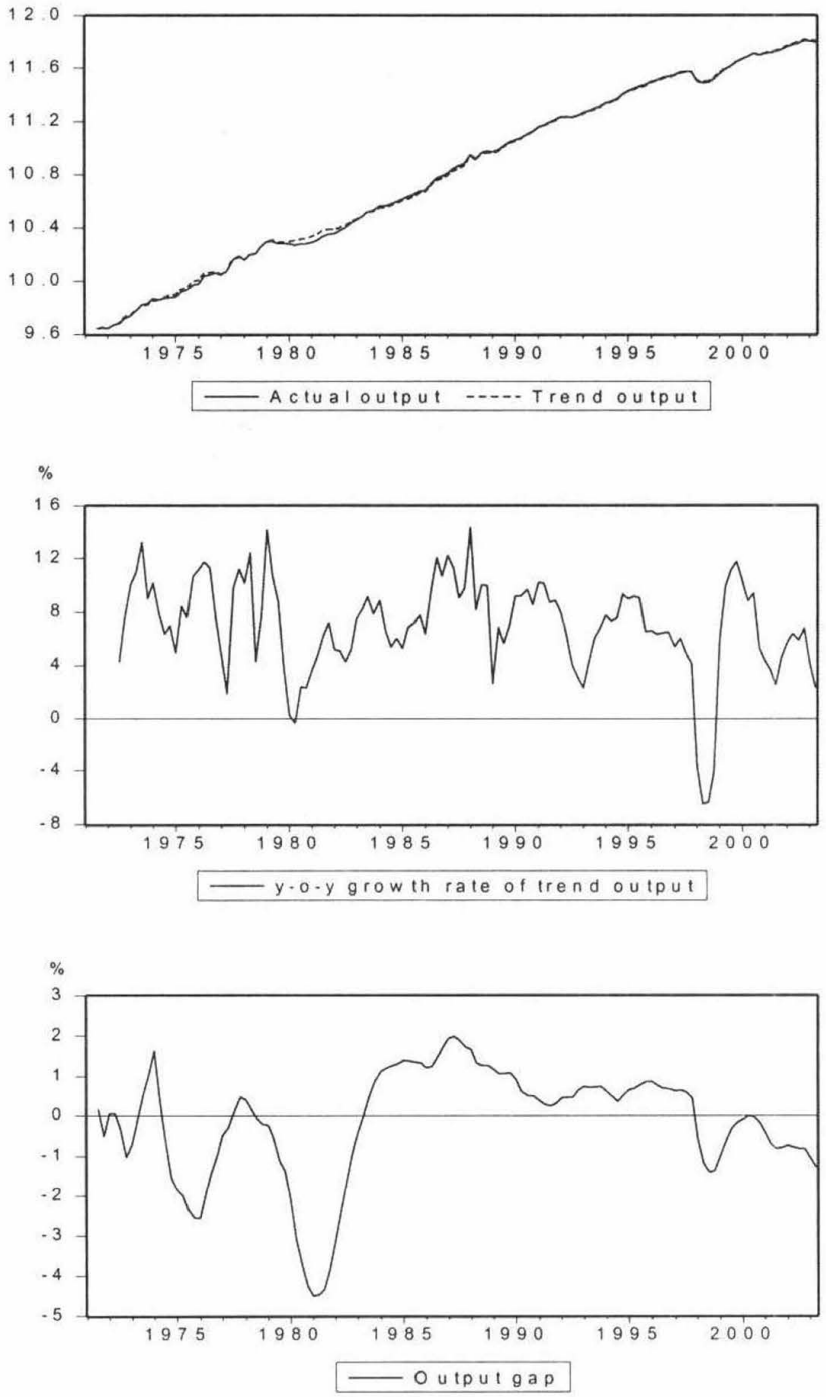

FIGURE 12

TREND OUTPUT AND GAP FROM UNRESTRICTED OUTPUT-PHILLIPS CURVE-INVESTMENT UC MODEL (INVESTMENT IN CYCLE SPECIFICATION 1971Q3-2003Q2)

To summarize, taking long-term trend in investment into account helps to identify growth trends and cycle patterns over time. Those patterns remain relatively stable with respect to the use of different types of investment variables (fixed investment, or facility investment), or how it is incorporated in the model (in trend component, or in cycle component). A notable exception appears when investment is included as a component of cycle. That specification gives rise to a positive output gap in recent years.

V. Conclusion

Results from several unobserved components models, which decompose observed output series into a random walk trend and stationary cycle components, indicate that long-term trend growth rate has shifted downward in the post-1997 crisis period as the event seemed to have inflicted some permanent adverse impact. For instance, according to the bivariate UC models, the growth rate of the trend output (sum of the long-run growth rate and past and present shocks) for the recent period has been lowered to around 5%. Such a pattern of trend output growth is accompanied by a persistent negative output gap since the 1997 crisis. These results are most sanguine when investment is explicitly considered as an explanatory variable. However, uncertainties associated with the results are quite large stemming from the following considerations.

First, according to our estimation results, the 1997 financial crisis appears to have had perceptible impact on output data properties, which, in turn, can affect inferences based on time series models including a univariate unobserved component model framework. Particularly, it appears that the data generating system of output has yet to settle after being subject to a very large shock that was the 1997 crisis. In such a situation, a few new observations can cause a large revision to the model. This can be seriously problematic because inferences about the economy – *e.g.*, long-run growth estimates, cyclical characteristics – might need to be changed frequently as new data become available. Results from the univariate model of this paper illustrate this point. The AR coefficients change drastically with the inclusion of post-crisis data and the trend output growth has been lowered to below 4% for recent period. These findings tend to exacerbate the so-called

end-point problem of the HP filter that is casually used as a convenient detrending tool. However, moving to a multivariate framework seems to mitigate such problems.

In particular, more stable trend-cycle decompositions were obtained when we added the Phillips curve to the univariate framework. At the same, adding a lagged investment variable as an exogenous explanatory variable to the trend and cycle components did not change results materially.

The second point is not drawn from this paper's exercises but remains relevant to discussions about Korean economy's growth potential. Namely, there has been a perceptible reduction in labor inputs in the form of a lower utilization of the existing working age population in the post-crisis period. The drastic decline and subsequent slow recovery of investment is a definite source of concern for future potential growth. However, the rate of capital formation as well as labor utilization could rise without putting undue strain on resources, at least in the medium term. Persistent robust growth in exports, for example, could provide a boost to both investment and employment.

We interpret results gained to date as raising a cautionary flag to the practice of making hasty assessments of Korea's growth potential in the near future. In particular, it is premature to issue pronouncements that the Korean economy has now entered an era of low growth, *i.e.* an annual growth rate in the 4% range. More data would surely help. In the meantime, the downside risk of such a position leading to a self-fulfilling prophecy is considerable.

(*Received 29 October 2003; Revised 21 December 2003*)

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