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The macroeconomic effects of government debt in Korea

JANG C. JIN* and W. DOUGLAS McMILLIN‡

* *Department of Economics, Kyungnam University, Masan, South Korea* and ‡ *Department of Economics, Louisiana State University, Baton Rouge, LA 70803-6306, USA*

The effects of government debt on output, the price level, the interest rate, and the real exchange rate in Korea is examined. The framework of analysis is a nine variable vector autoregressive model. The effects of debt are evaluated by computing variance decompositions, impulse response functions, and cumulative impulse response functions. The variance decompositions indicate significant effects on output, the price level, the interest rate, and the exchange rate. The impulse response functions indicate significant, negative short-run effects of a shock to debt on these variables, but no long-run effects. The cumulative impulse response functions yield similar results. These results are at odds with conventional models in which government debt is wealth, but appear consistent with some models in which debt is not wealth.

I. INTRODUCTION

One of the more intensively studied topics in recent years is the macroeconomic effects of government debt. The impetus for much of this research is the Ricardian equivalence hypothesis which suggests that government debt is not wealth and that a switch from (lump-sum) tax finance to debt finance of a given level of government purchases has no first-order macroeconomic effects.[1] This contrasts with the Keynesian view that government debt is wealth and that a switch from tax-to-debt finance has substantive macro effects. No consensus has emerged on the theoretical appropriateness of the Ricardian position. The theoretical foundations of the Ricardian equivalence hypothesis are discussed in Barro (1974, 1989), and criticisms of this view are found in Tobin and Buiter (1980), Brunner (1986), Bernheim (1989) and Haliassos and Tobin (1990).

Theoretical disagreement on the role of government debt is matched by mixed empirical evidence. Surveys of the empirical evidence are found in Bernheim (1989) and Barth *et al.* (1991). Most studies of the macroeconomic role of government debt have focused upon developed economies. However, it seems desirable to consider the effects of government debt in developing economies as well. One such economy well-suited to the study of the macroeconomic effects of government debt is the Korean economy which has grown rapidly over the last several decades and has simultaneously run persistent government budget deficits that led to a rapid expansion of government debt. Although the Korean economy has been characterized by rapid growth of economic activity and government debt, only Evans (1988, 1990) has conducted even a limited study of the macro effects of Korean government debt. Evans estimated the effects of deficits on output using a single equation framework and found a negative, but statistically insignificant effect of deficits on output.

The aim of this paper is to examine empiricially the effect of government debt on output, the price level, the interest rate and the exchange rate within a nine variable vector autoregressive (VAR) model of the Korean economy. Monthly data for the period 1973:5–1989:11 are used in the analysis. In the traditional view, changes in government debt have implications for a variety of macro variables, and it is, thus, important to simultaneously examine the effects of debt on these variables. The variables included in the model are consistent with the reduced form of an aggregate demand-aggregate supply framework, where the IS–LM–BP model underlies the aggregate demand side. Output, the

[1] In this view, the discounted value of future taxes required to service and retire government debt issued when current period taxes are cut is equal to the market value of the debt. Thus, government debt is not private sector wealth and a switch from (lump-sum) tax finance to debt finance of government expenditures has no effects on output, the price level, the interest rate or the exchange rate.

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price level, the interest rate, the money supply and two fiscal policy variables, government spending and the stock of government debt, are included as are the real exchange rate and two external shock variables. The latter two variables measure output and price shocks emanating from Korea's major trading partners. The impact of government debt is examined by computing variance decompositions (VDCs), standard impulse response functions (IRFs), and cumulative impulse response functions (CIRFs) for which standard errors are calculated using a Monte Carlo simulation approach.

The VAR modelling approach is employed since there is little agreement on the appropriate structural model and since few restrictions are placed on the way in which the system's variables interact in the estimation of the system. In the specification and estimation of the model, all variables are treated as jointly determined; no *a priori* assumptions are made about the exogeneity of any of the variables in the system at this stage of analysis. However, in the computation of the VDCs, IRFs, and CIRFs, some decisions about structure must be made. These decisions are discussed in Section III, but the results are not sensitive to the decisions made about structure.

Section II discusses the data and the specification of the model while the empirical results are presented and analysed in Section III. Our conclusions are summarized in Section IV.

II. DATA DESCRIPTION AND MODEL SPECIFICATION

As noted earlier, the macroeconomic effects of government debt are examined within the context of a nine-variable VAR model. The model is specified and estimated using monthly data for 1973:5–1989:11. The period 1973:5–1975:8 is used as presample data to generate the lags in the VAR, and the model is estimated over the period 1975:9–1989:11. Monthly data are used for two reasons. One reason is that the size of our system requires monthly data in order to have enough degrees of freedom for estimation. The second reason is based on a desire to minimize any problems with temporal aggregation (see Christiano and Eichenbaum, 1987) that might arise with the use of quarterly or annual data. Our sample begins in 1973:5 since this is the earliest date for which we can obtain our interest rate series. Furthermore, the beginning of our sample roughly coincides with the period in which the Korean government placed increased reliance on the sale of bonds to the private sector and with the demise of the Bretton Woods system. The end of our sample coincides with the latest data available to us at the inception of the study.

The model variables include: (1) industrial production (y) which is employed as a proxy for output; (2) the consumer price index (P) which is our measure of the domestic price level; (3) the yield on national housing bonds (r) which is used as a proxy for the long-term interest rate; (4) the narrowly defined money supply ($M1$) which is the monetary policy variable; (5) real government expenditures (g); (6) the par value of privately held government debt (D); (7) the real exchange rate (e); (8) a foreign output shock measure (y^*); and (9) a foreign price shock measure (P^*). All data with the exception of that discussed below were taken from the March 1990 *International Finance Statistics* (IFS) tape produced by the International Monetary Fund. All data except for e and r are seasonally adjusted. The use of seasonally adjusted data was required since industrial production data were available only in seasonally adjusted form.

The yield on national housing bonds was selected as the proxy for the long-term rate since no other consistent series for a long-term rate is available over our sample. Our focus is upon the long-term rate rather than a short-term rate, since it is generally thought that investment decisions depend, more closely, upon the long-term rate than the short-term rate. Real government expenditures are measured as the total expenditures of the consolidated central government and are deflated by the consumer price index. This measure includes transfer payments; ideally, a series that excludes transfer payments would be preferred, but no series of this type is available monthly. It is important to include government expenditures in our model since government expenditures can affect economic activity even if Ricardian equivalence holds. Since government expenditures and debt are correlated, macro effects due to changes in government spending might be incorrectly attributed to government debt if government spending were omitted from the model.[2]

The series for the par value of privately held government debt is constructed by subtracting the central bank's holdings of debt from total government debt outstanding. Total outstanding debt is from the IFS tape while central bank holdings of debt are calculated from information in the *Monthly Statistical Bulletin*, Bank of Korea. The par value of government debt, rather than the market value, is employed for the following reason. Changes in the nominal interest rate automatically lead to changes in the market value of debt. Links between the market value of debt and economic variables may thus reflect a relationship between nominal interest rates (which reflect, in part, changes in expected inflation as well as other expectational effects) and these economic variables rather than a link between government debt and these variables. Hafer and Hein (1988) demonstrate that these concerns are of empirical significance in a bivariate study of the relationship between the par value of US

[2] It would also be desirable to include a measure(s) of the marginal tax rate(s) since changes in distortionary taxes can affect output even if Ricardian equivalence holds. This was not done since no reliable marginal tax rate series are available.

federal debt and inflation and the market value of federal debt and inflation. They find no evidence of Granger-causality from the par value of debt to inflation, and they show that Granger-causality from the market value of debt to inflation essentially disappears when they control for nominal interest rates. In order to reduce problems of interpretation, the focus of this study is on the par value of debt.

The real exchange rate is a multilateral trade-weighted measure constructed in the manner outlined by Rhomberg (1976). The countries used in the construction of the exchange rate are the US and Japan, Korea's two major trading partners.[3] Since the nominal exchange rates used in the construction of this measure are the number of Korean won per unit of foreign currency and the weighted average of these nominal rates is multiplied by the ratio of the foreign price level to the Korean price level, an increase (decrease) in the real exchange rate represents depreciation (appreciation). We note that the estimation period includes the 1975–79 period in which the nominal exchange rate was fixed and the subsequent period in which a managed float applied. Data are combined from these periods for several reasons. One is based on Stockman's (1986) argument that, if debt is wealth, changes in debt will alter real exchange rates regardless of the exchange rate regime. A second reason is that not enough data are available to estimate the model over just the fixed rate subperiod of our sample. Furthermore, a formal stability test yielded no evidence of a structural shift after 1979.[4]

The foreign output shock variable is measured as a weighted average of industrial production in the US and Japan, and the foreign price variable is a weighted average of the wholesale price indices in the US and Japan.[5] Because of the openness of the Korean economy, it is important to include variables like the real exchange rate and the foreign price and output variables.

Prior to specification and estimation of the VAR, augmented Dickey–Fuller tests were employed to check for first-order unit roots. These tests suggested that first differences of the logs of y, P, $M1$, g, D, e and P^* and the first difference of the level of r should be used in specifying and estimating the model. The evidence was more ambiguous for y^*; a unit root could be rejected at the 5% level but not at the 1% level. Based upon the arguments of Engle and Granger (1987), cointegration tests were also performed for the eight variables that required differencing to achieve

[3]The real exchange rate is defined as

$$e=[s_m EXCH_m + s_x(1/EXCH_x)](P^*/P)*100,$$

where $s_m = M^T/(M^T+X^T)$, $s_x = X^T/(M^T+X^T)$, $EXCH_m = \Sigma(M_i/\Sigma M_i)(E_{it}/E_{i0})$, $EXCH_x = \Sigma(X_i/\Sigma X_i)(E_{i0}/E_{it})$, M^T denotes Korea's total imports from the world market as an annual average over the period 1973–87, measured in US dollars, X^T denotes Korea's total exports to the world market as an annual average, measured in US dollars, M_i denotes Korea's imports from country i (the US and Japan) as an annual average, measured in US dollars, X_i denotes Korea's exports to country i (the US and Japan) as an annual average, measured in US dollars, E_{it} denotes the nominal exchange rate as the ratio of the Korean won to currency i at time t, E_{i0} denotes the nominal exchange rate as the ratio of the Korean won to currency i in the base period, P is the price index in Korea, and P^* is the foreign price level.

The first term of the equation is an import weighted index ($EXCH_m$), while the second term is the reciprocal of an export weighted index ($EXCH_x$). These two indices are weighted, respectively, by import shares (s_m) and export shares (s_x). A more detailed discussion of this measure is provided in Jin (1991).

We note that although only data from Korea's two major trading partners are used in constructing e, our measure moves in a similar fashion over the common sample of 1973–82 to the exchange rate of Edwards (1989), who focuses upon Korea's ten largest trade partners in 1975.

[4]The test is a multivariate extension of the procedure suggested by Dufour (1980, 1982). The system was first estimated with 12 lags on each variable over the period 1975:9–1989:11. 0–1 dummy variables for each observation in the fixed nominal rate period 1975:9–1979:12 were then added to each equation in the system, and this system was estimated over the sample 1975:9–1989:11. The joint significance of the coefficients on the dummy variables was tested by computing the likelihood ratio statistic

$$(T-C)(\log|DR|-\log|DUR|)$$

where $|DR|$ is the determinant of the variance–covariance matrix of the restricted system, $|DUR|$ is the determinant of the variance–covariance matrix of the unrestricted system (the system with the dummy variables), T is the number of observations from 1975:9–1989:11 and C is the number of parameters in each unrestricted equation (161). ($T-C$) is the small sample correction suggested by Sims (1980). The statistic is distributed as χ^2 with degrees of freedom equal to the number of restrictions (number of dummy variables in the system, 468 in this study). The marginal significance of the calculated χ^2 was 0.99. The hypothesis that the coefficients on the dummy variables jointly equalled zero could not be rejected and, hence, no instability was indicated.

[5]The measures are defined as

$$\begin{aligned} y^* &= w_{US} y_{US} + w_{JP} y_{JP} \\ P^* &= w_{US} P_{US} + w_{JP} P_{JP} \\ w_{US} &= s_x(X_{US}/\Sigma X_i) + s_m(M_{US}/\Sigma M_i) \\ w_{JP} &= s_x(X_{JP}/\Sigma X_i) + s_m(M_{JP}/\Sigma M_i) \text{ and} \\ & i = \text{US and Japan.} \end{aligned}$$

w_{US} and w_{JP} are, respectively, Korea's trade weights with the US and Japan; y_{US} and y_{JP} are, respectively, industrial production indices for the US and Japan; and P_{US} and P_{JP} are, respectively, wholesale price indices for the US and Japan.

stationarity and then for all nine variables.[6] Since no evidence of cointegration was found, the system was estimated with the differences of y, P, $M1$, g, D, e, r and P^* and the log level of y^*. However, since a unit root in the log level of y^* could not be rejected at the 1% level, the robustness of the results was checked by estimating a system that included the first difference of the log of y^*. The results were robust to the treatment of y^*.

Akaike's AIC criterion was used to determine the lag length of the VAR model. Use of this criterion suggested an optimal lag of 12 months. Q-statistics indicated the absence of any serial correlation in the residuals of the model.[7]

III. EMPIRICAL RESULTS

As indicated in Section I, the effects of government debt are analysed through computation of VDCs, IRFs and CIRFs which, in turn, are based on the moving-average representation of the VAR model and reflect both direct and indirect effects. The VDCs show the percent of the forecast error variance for each variable that may be attributed to its own innovations and to fluctuations in the other variables in the system. Therefore, the VDC for y indicates the percent of the forecast error variance in y accounted for by D and the other variables in the system. This suggests that if D is an important determinant of movements in y, it should explain a significant portion of the forecast error variance in y. Moreover, Sims (1982) has suggested that VDCs give an indication of the strength of Granger-causal relations that may exist between variables. Therefore, if D explains a large and significant portion of the forecast error variance of y, this could be interpreted as a strong Granger-causal relation.

The IRFs indicate the size and direction of effect of a one standard deviation shock to one of the system's variables on the other variables in the system. By computing IRFs, the direction of effect of a shock can be examined to D on e, r, y and P. Since D is converted to log differences as are e, y, and P and since r is converted to differences prior to estimation, the IRFs reported here indicate the effect of a shock to the growth rate of D on the growth rates of e, y and P and the change in r. It is also of interest to know whether shocks to D have significant effects on the levels of e, r, y and P. This is accomplished by computing cumulative IRFs. The value of the CIRF in any period is the sum of the IRF value in the current and prior periods.

Since Runkle (1987) has argued that reporting VDCs and IRFs without standard errors is similar to reporting regression coefficients without t-statistics, a Monte Carlo integration procedure like that described in Doan (1990) is employed to estimate standard errors for the VDCs, IRFs and CIRFs. One thousand draws are employed in the Monte Carlo procedure. For the VDCs, the estimates of the proportion of forecast error variance explained by each variable are judged to be significant if the estimate is at least twice the estimated standard error. For the IRFs and CIRFs, a two standard deviation band is constructed around the point estimates. If this band includes zero, the effect is considered insignificant.

Since the equations of the VAR contain only lagged values of the system's variables, it is assumed that the residuals of the VAR model are purged of the effects of past economic activity. Any contemporaneous relations among the variables are reflected in the correlation of residuals across equations. In this paper, the Choleski decomposition is used to orthogonalize the variance–covariance matrix. In this approach, the variables are ordered in a particular fashion, and, in this way, some structure is imposed in the computation of the VDCs, IRFs and CIRFs. When a variable higher in the order changes, variables lower in the order are assumed to change. The extent of the change depends upon the covariance of the variables higher in the order with those lower in the order.[8]

The orderings chosen for study are the following: (1) y^*, P^*, $M1$, g, D, e, r, y, P; and (2) y^*, P^*, y, P, e, r, $M1$, g, D. Ordering (1) is chosen for the following reasons. Placement of y^* and P^* first is based on the assumption that any contemporaneous effects flow from the large economies of the US and Japan to the Korean economy. Placement of the policy variables next is consistent with the familar textbook (IS–LM–BP) treatment of aggregate demand in which current period shocks to the policy variables contemporaneously affect e, r, y and P. Assumed in this ordering are

[6] The lag length for the unit root and cointegration tests was determined using the criterion suggested by Schwert (1987). Cointegration tests of the sort suggested by Engle and Yoo (1987) were performed. However, since Hansen (1990) pointed out that the power of this test, as well as the test proposed by Johansen (1988), falls substantially as the size of the system increases, Hansen's two-stage test was also employed. The power of Hansen's test is unaffected by the size of the system. Neither the Engle and Yoo nor the Hansen tests yielded any evidence of cointegration.

[7] Because the optimal lag chosen was also the maximum considered, the robustness of the results was checked by estimating a 13-lag model. The results were essentially the same as for the 12-lag model. A desire to conserve degrees of freedom prevented us from checking longer lags.

[8] One alternative to the Choleski decomposition has been suggested by Bernanke (1986). This is a two-stage procedure in which the residuals from the VAR are used in the estimation of a structural model. The residuals from this structural model are then treated as 'fundamental' shocks. However, unless the structural model is just identified, in general, there will be correlation across equations in the residuals of the structural model, and the issue of appropriate ordering arises again.

A second alternative is recommended by Blanchard and Quah (1989). In this procedure, long-run constraints that are, in principle, consistent with alternative structural models are imposed in order to obtain estimates of the fundamental shocks. Implementation of this procedure in a model the size of ours is a formidable task and is beyond the scope of this paper.

that current period shocks to e, r, y and P have no contemporaneous effect on the policy variables; this is consistent with the typical policy reaction functions in which the current values of the policy variables depend only on the lagged values of domestic macro variables. Placement of e and r next is consistent with the transmission mechanism embedded in the typical formulation of aggregate demand. Finally, placement of y and P last allows these variables to respond directly and indirectly to contemporaneous shocks to all other model variables. Thus, ordering (1) is designed to be consistent with a model in which the IS–LM–BP model underlies aggregate demand and where the exchange rate, the interest rate, output and the price level respond to current innovations in domestic policy variables as well as foreign shocks.

Ordering (2) places the policy variables last in the ordering, with D occupying the very last position. This allows contemporaneous effects of all other model variables on D. Furthermore, in this ordering, the effects of D on the other variables do not depend on the order in which these variables precede D. Ordering D last is consistent with the set of structural models in which the other model variables have both direct and perhaps indirect contemporaneous effects on D. This purges the shocks to D of any effects of current economic activity on D. Thus, the shocks to D when D is ordered last represent variation in D that is independent of current and past economic activity.

Table 1. *Variance decompositions*[a]

Variable	Horizon	Explained by shocks to debt	
		Ordering 1	Ordering 2
e	6	6.1 (2.9)*	6.5
	12	6.3 (2.8)*	6.5
	24	6.3 (2.7)*	6.3
	36	6.9 (2.8)*	6.8
	48	7.1 (3.0)*	7.1
r	6	3.1 (1.9)	3.0
	12	7.3 (2.7)*	6.0
	24	6.9 (2.6)*	6.4
	36	9.1 (3.2)*	8.2
	48	9.4 (3.5)*	8.3
y	6	4.4 (2.6)	4.3
	12	6.1 (2.6)*	5.5
	24	5.5 (2.5)*	5.3
	36	7.1 (3.1)*	6.9
	48	7.3 (3.4)*	7.2
P	6	12.7 (4.2)*	12.5
	12	18.0 (5.1)*	17.9
	24	19.4 (5.4)*	18.1
	36	19.3 (5.3)*	18.0
	48	19.0 (5.3)*	17.6

[a]Standard errors are in parentheses next to the point estimates. *Indicates the point estimate is at least twice the standard error. All entries in column (2) are within one standard deviation of those in column (1).

The VDCs for both orderings are reported in Table 1. Column (1) reports results for ordering (1) while column (2) reports results for ordering (2). The estimated standard errors are in parentheses beside the point estimates. A* indicates the point estimate is at least twice the standard error. VDCs at horizons of 6, 12, 24, 36 and 48 months are shown in order to convey a sense of the dynamics of the system. Only the effects of D on e, r, y, and P are shown in order to focus upon the variables of central interest to the paper and to conserve space. It is observed that, for ordering (1), with only a few exceptions, the effects of D on the macro variables in Table 1 are significant.[9] The results are essentially unchanged when D is ordered last. The point estimates from ordering (2) are all within one standard deviation of those in column (1).[10]

The VDCs, thus, indicate a significant effect of D on the macroeconomy. However, the VDCs give no indication of the direction of effect; this information is obtained from the IRFs and CIRFs which are presented in Figs 1 and 2. In these figures, point estimates of the IRFs are plotted with a dotted line while the solid lines represent a two standard deviation band around the point estimates. Figure 1 presents the results of standard IRFs. It is observed that the significant effects in Fig. 1 are all negative and that the effects are not significantly different from zero over longer horizons. In the case of e, the effect of a shock to D is initially positive (indicating depreciation) but not significant and that the effect quickly becomes negative. The negative effects are significant at horizons of three and four months. For r, the initial effects are negative, and significant negative effects are observed at horizons of eight and twelve months, although the 12-month effect is of marginal significance. For y, the initial effects are positive but not significant. Significant negative effects are observed at horizons of four and eight months. More significant negative effects are observed for P than for the other variables. Significant negative effects are observed at horizons of two and four months and then again at horizons of 10–13 months.

Figure 1 indicates the response of the growth rates of the exchange rate, output and the price level and the change in the interest rate to shocks to the growth rate of D. It is of interest to determine whether there are any lasting effects on the levels of e, r, y and P; this information is provided in

[9]We note that the effects of D on $M1$ are also significant. The IRFs indicate the initial effects of D on $M1$ are negative, which does not support arguments that government debt is monetized. We also note that D does not, as expected, have any significant effects on y^* and P^*.

[10]The robustness of the results is checked in several ways. First, as noted in footnote 7, the lag was extended to 13 months. The results were all within one standard deviation of those in column (1) of Table 1. Second, the model was estimated with all variables in first differences. The results in the text are for a model in which y^* is in levels since the hypothesis of a first-order unit root for this variable could be rejected at the 5% level. However, the hypothesis could not be rejected at the 1% level, so, it was decided to estimate the system with the first difference of the log of y^* in place of the log level of y^*. Again, the results were all within one standard deviation of those in column (1) of Table 1.

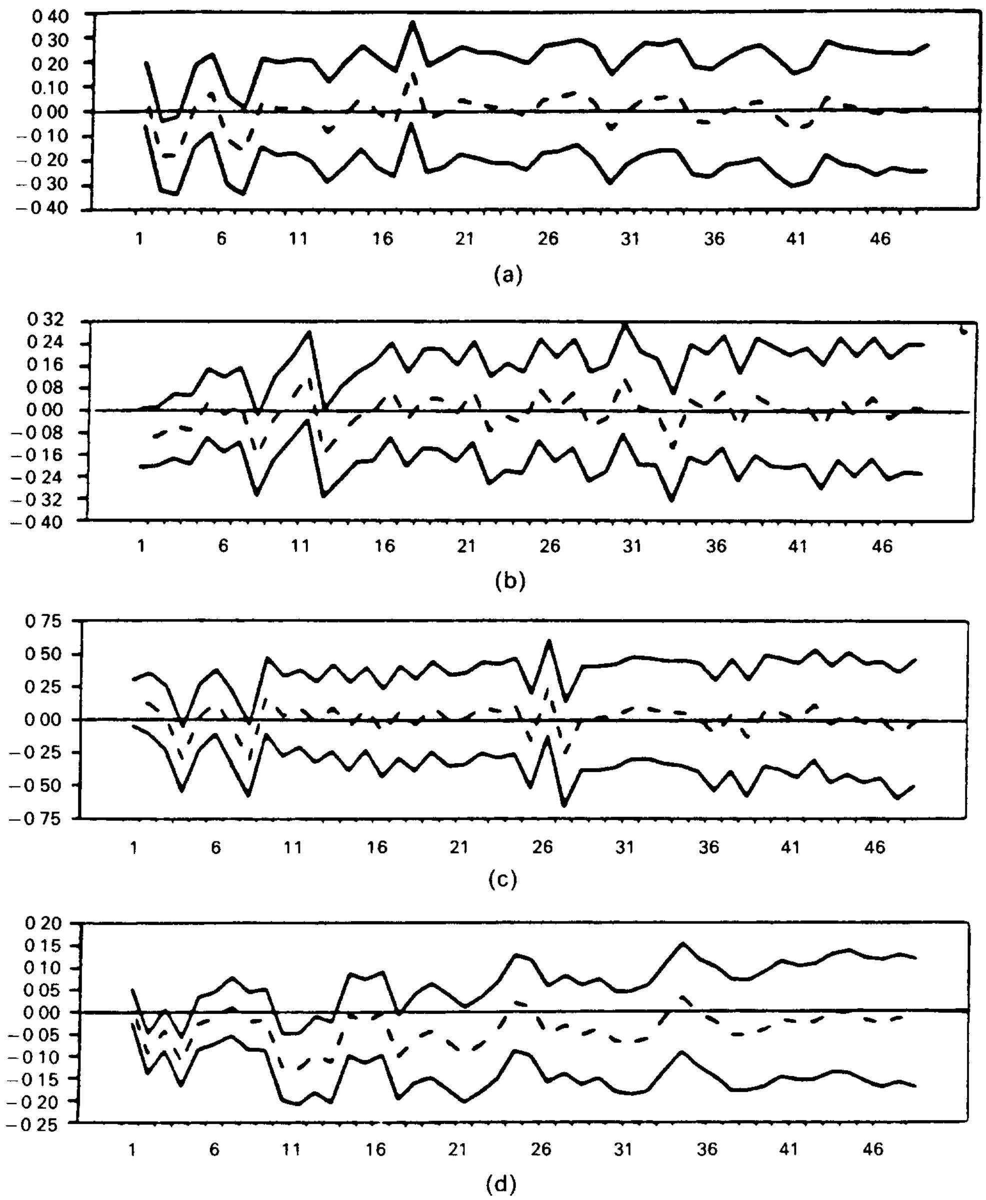

Fig. 1. *Impulse response functions (a) response of real exchange rates to D innovation (b) response of interest rates to D innovation (c) response of output to D innovation (d) response of prices to D innovation*

Fig. 2 which reports the CIRFs. In the case of e, one essentially observes no significant effects of D on the level of the exchange rate. Significant negative effects of D on the level of r is observed, but these effects are not significantly different from zero after 18 months. In the case of y, there is a marginally significant positive effect at a horizon of three months. It appears that the initially positive effect observed in period 1 in Fig. 1 slightly outweighs the subsequent negative effects. However, after this slight positive effect at month three, the remaining effects are not significantly different from zero. Significant negative effects are observed for P. Although not significant in every month, negative effects are observed for about 27–28 months. After this, the effects are not significantly different from zero.

The IRF and CIRF results do not appear to support the conventional view that government debt is wealth. The significant effects in Figs 1 and 2 are, with one minor exception, all negative. The negative effects also appear to be at odds with the Ricardian view in which debt is not wealth and debt shocks have no effect on the macroeconomy. These negative effects are generally consistent with the findings of Evans (1988, 1990) for the Korean economy even though his methodology differs from that used here.

One explanation for some of the negative effects found here has been suggested by Kormendi (1983) based on *a priori* argument of Barro (1974). The argument is that because of uncertainty about the individual's share of future taxes and the timing of taxes, individuals may perceive the

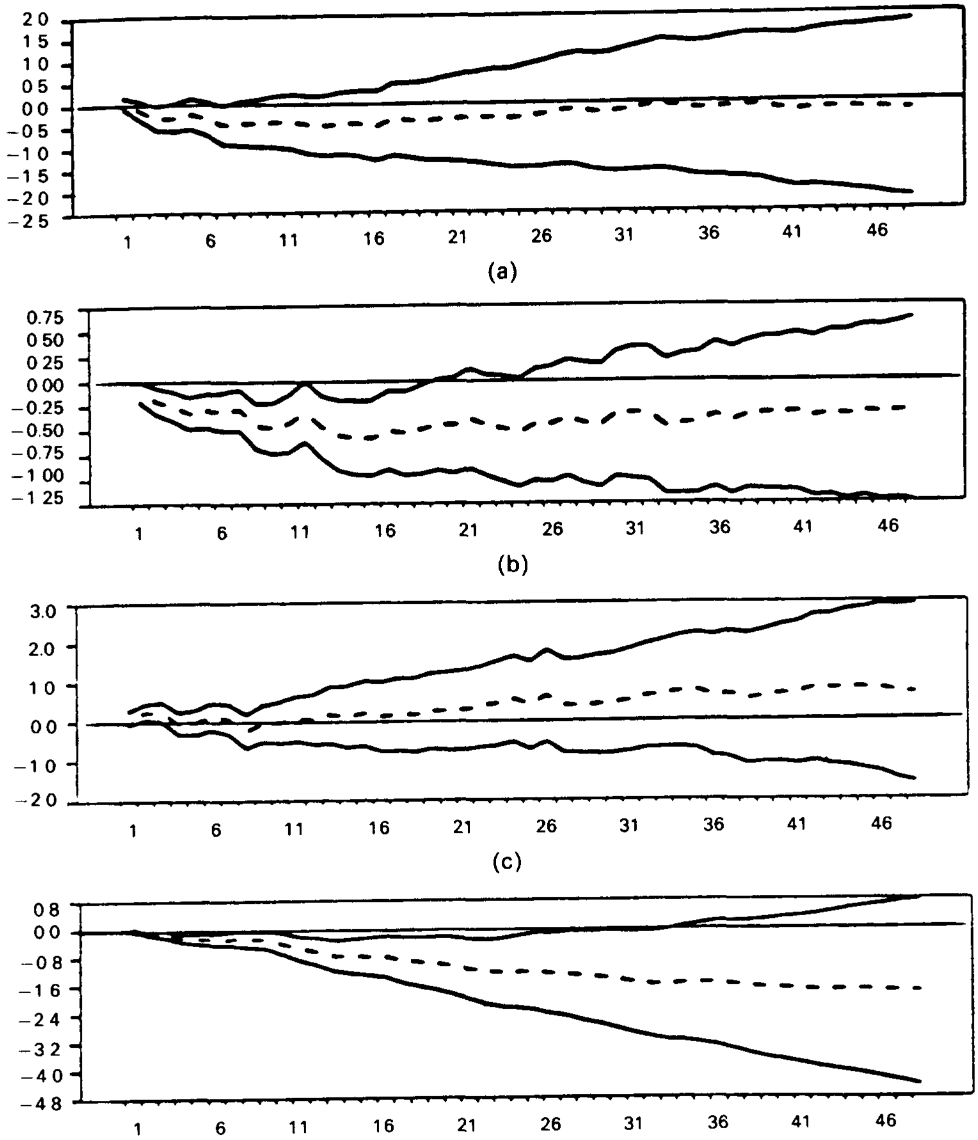

Fig. 2 *Cumulative impulse response functions* (*a*) *CIRF of real exchange rates to D innovation* (*b*) *CIRF of interest rates to D innovation* (*c*) *CIRF of output to D innovation and* (*d*) *CIRF of prices to D innovation*

present value of the implied future taxes associated with debt finance to be greater than the present value of the income stream associated with the government bonds. In this case, private sector wealth falls as does aggregate demand. This argument can explain the negative effects on r, y and P, but the negative effect on e remains a puzzle. If debt is wealth, the IS–LM–BP model would suggest that a substitution of debt for tax finance of a given level of expenditures would lead to an appreciation of the exchange rate (decline in e). However, if wealth actually falls, a depreciation (rise in e) would be expected. We note that in Fig. 1, e does rise initially, but this effect is not significant. The only significant effect in Fig. 1 on e is negative. It is noted in Fig. 2, which focuses on the level of e, that there are no significant effects of D on the exchange rate.

IV. CONCLUSION

This paper has examined the effects of government debt on output, the price level, the interest rate, and the real exchange rate in Korea. Unlike most studies of the effects of government debt on the macroeconomy which concentrate on developed economies, this study focuses upon a rapidly growing developing economy, one in which rapid growth has been accompanied by persistent expansion of government debt held by the private sector. The framework of analysis is a nine-variable vector autoregressive model which includes measures of output, the price level, the interest rate, the real exchange rate, the money supply, government expenditures, government debt and foreign output and price shock variables.

The effects of government debt are evaluated through the computation of variance decompositions, impulse response functions and cumulative impulse response functions. The variance decompositions indicate significant effects of government debt on output, the price level, the interest rate and the exchange rate. However, the impulse response functions indicate that the only significant effects of a shock to debt on the growth rates of the exchange rate, output and the price level and the change in the interest rate are negative. The longer-run effects are, however, not significantly different from zero. The cumulative impulse response functions indicate debt shocks have, at best, marginally significant short-run effects on the levels of the exchange rate and output. Significant negative effects on the levels of the interest rate and the price level are observed in the short run. Again, in all cases, there are no long-run effects of debt shocks on the levels of these variables.

In the conventional Keynesian model, a switch from tax-to-debt finance of a given level of government expenditures raises aggregate demand. In the short run, output is affected although there are no long-run effects. Both short- and long-run effects are expected for the real exchange rate, the interest rate and the price level. With the exception of output, the long-run results appear at odds with the conventional view of debt, and the short-run results lead to a similar conclusion since the short-run effects are in the opposite direction of those predicted by the conventional view.

The significant negative effects of debt on the interest rate, output and the price level appear consistent with the argument of Kormendi (1983) and Barro (1974) that uncertainty about the individual's share of future taxes and the timing of taxes may cause individuals to perceive the present value of the implied future taxes associated with debt finance as more than the present value of the income stream associated with the government bonds. In this case, wealth and aggregate demand fall.

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