Journal of International Economics 32 (1992) 241–263. North-Holland

Empirical evidence on the insulation properties of fixed and flexible exchange rates

The Japanese experience

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Received July 1989, revised version received March 1991

This paper investigates whether Japanese output stability since the mid-1970s is attributable to the change in exchange rate regime or to the changing nature of underlying disturbances. We decompose foreign and domestic influences on output movements using restrictions on the long-run dynamics to identify the disturbances. We find that the flexible exchange rate regime is more effective in insulating the economy from foreign disturbances than is the fixed rate regime. However, in the case of Japan, it is primarily the changing nature of the fundamental disturbances to the economy that is responsible for greater output stability since the mid-1970s.

1. Introduction

The introduction of generalized floating exchange rates in the early 1970s was widely expected to increase the degree to which national economies would be insulated from foreign disturbances, particularly foreign monetary disturbances. The high covariance of national output fluctuations during the recent experience with flexible exchange rates, however, has led many to question the insulation properties of flexible rates. This has been an important motivation for the development of a large body of theoretical

*The authors thank the Ford Foundation and the Graduate Division, UCSC, for financial support, and Sven Arndt, Koichi Hamada, Jürg Niehans, Steve Sheffrin, two anonymous referees and seminar participants at the Ford Foundation seminar in Tokyo, the University of California, Davis, the Pacific Rim Conference, and the University of California, Santa Cruz, for helpful comments. Michael Curtis and Cindy Beltz provided excellent research assistance.

work that re-examines the insulation properties of flexible exchange rates and the properties of optimal exchange rate regimes. Central to most of this literature is the notion that a flexible exchange rate regime may not insulate an economy to the extent previously believed. Dornbusch (1983), for example, concludes that '...flexible rates leave us with as much interdependence, or even more, as there is under a fixed rate regime' (p. 4).

On theoretical grounds, foreign disturbances will affect real domestic output under flexible exchange rates through channels such as the terms of trade, the real interest rate and the aggregate price level [Artus and Young (1979), Dornbusch (1983), Marston (1985)]. Marston (1985), for example, argues that insulation applies only in special cases, and that in general flexible rates do not even insulate the economy from foreign monetary disturbances.[1]

In contrast to these theoretical developments, relatively little systematic empirical work has been done comparing the transmission of business cycles across fixed and flexible exchange rate regimes. Moreover, most of the existing empirical literature has focused on measuring the covariances between national output movements under fixed and flexible exchange rate arrangements, but not on isolating the underlying causes of the covariation.[2] In particular, it is not clear whether the strong covariation observed in national outputs under flexible rates has been attributable to the lack of isulation from business cycle disturbances or to the changing nature of the underlying country-specific and world-wide shocks.[3]

This paper investigates the insulation issue from a different perspective. Rather than address the question of why many industrial countries appear so tightly linked under flexible exchange rates, our empirical focus is on the regime shift experience of one country (Japan) where it appears that the stylized facts lend support to the superior insulation and monetary indepen-

[1]The increased emphasis on the interdependence of national economies under flexible rates spans a broad literature and a variety of theoretical frameworks. In the context of non-stochastic extended Mundell–Flemming models [Mussa (1979), Tobin and DeMacedo (1980) and Sweeney (1985)], flexible rates will in general not insulate the economy perfectly from foreign monetary disturbances due to the effects of exchange rate changes on desired expenditure [the Laursen–Metzler (1950) effect], the aggregate price level (working through changes in real money balances and aggregate supply), and revaluation of asset stocks. Similar results are found in the context of dynamic perfect foresight models, e.g. Dornbusch (1983) and Niehans (1984), as well as in the context of stochastic rational expectations models of the open economy, e.g. Saidi (1980), Duck (1984), Frisch (1985), Marston (1985) and Glick and Wihlborg (1990).

[2]For example, Swoboda (1983) uses correlation and principle-component analysis and finds that the cross-country correlation of price level, output, and interest rate changes were generally higher during the 1970s and 1980s than during the 1960s. Similarly, Gerlach (1988) uses cross-spectral methods and finds that output covariances between the United States and most other industrial nations have significantly increased following the move to managed floating.

[3]Dornbusch and Frankel (1988) argue on the basis of descriptive evidence that floating exchange rates have allowed the major industrial nations to follow largely independent policies, and that common shocks are largely responsible for positive cyclical output correlations.

Table 1
Sample statistics.

	Mean	Variance	Autocorrelations 1	2	3	4	5	6
			Japanese real GNP growth rate					
1956:2–1972:4	9.07	43.88	−0.15	0.09	−0.01	0.08	0.02	−0.17
1973:2–1986:4	3.50	10.18	0.16	−0.01	0.39	0.19	−0.06	0.09
			United States real GNP growth rate					
1956:2–1972:4	3.38	15.04	0.22	0.15	−0.06	−0.08	−0.14	0.06
1973:2–1986:4	2.27	19.43	0.34	0.23	0.02	−0.00	−0.01	−0.00

dence properties of flexible exchange rates. Among the major industrial economies, Japan offers the greatest prima facie evidence that the shift in exchange rate regime provided greater insulation from foreign shocks and, as a consequence, was of fundamental importance in reducing domestic output variability. Taylor (1988), for example, has termed Japan an 'outlier' because she experienced the most notable reduction in output variance among the major industrial countries with the move from fixed to flexible exchange rates. Table 1 reports summary statistics for quarterly growth rates (at annual rates) for real GNP in Japan and the United States. While Japanese output variance declined by more than 75 percent between 1956–1972 and 1973–1986, output variance in the United States increased by almost 30 percent over the same period.[4]

No existing empirical evidence has identified whether the move to a flexible exchange rate regime, or the changing nature of underlying disturbances to the economy, is primarily responsible for greater output stability in Japan. In particular, what remains unresolved is (i) whether the disturbances generating output fluctuations in the fixed rate period were primarily of domestic or foreign origin, and (ii) whether it was the move to flexible rates, or the reduction in domestic or foreign disturbances, which was primarily responsible for dampened output fluctuations in Japan. To address these questions we examine the nature of the disturbances that have accounted for post-war fluctuations in Japanese real output. We distinguish between foreign and domestic shocks in the context of a dynamic simultaneous equations model of Japan and the United States. The foreign shocks (from the perspective of Japan) are decomposed into real and monetary components. This empirical approach allows us to test the standard theoretical prediction that a given foreign shock has larger domestic output effects under fixed exchange rates than under flexible exchange rates, as well as to measure the

[4]The greater stability of real output in Japan is attributed by Taylor (1989) to its synchronized process of wage determination.

extent to which shocks of foreign origin contributed to the observed output variability in Japan during each exchange rate regime.

To achieve such a decomposition we must impose some restrictions on the dynamic system we examine. In the context of VAR (vector autoregression) systems, identifying restrictions are usually imposed on the contemporaneous interactions among the variables in the system, either by assuming a recursive structure, as in the approach popularized by Sims (1972), or by the simultaneous equations approach used by Bernanke (1986), Blanchard and Watson (1986) and Walsh (1987). We follow Blanchard and Quah (1988), Jun (1988) and Shapiro and Watson (1988), and identify the model parameters by imposing restrictions on the long-run dynamics of the model. In particular, imposing the restriction that monetary disturbances do not have permanent effects on real variables (long-run neutrality) aids us in achieving model identification while limiting the restrictions imposed on the contemporaneous interactions among the variables in the system.

The rest of the paper is organized as follows. Section 2 discusses the empirical support for the view that the reduction in Japanese output variance is primarily attributable to the superior insulation properties of the flexible exchange rate regime. We suggest a plausible alternative explanation, also consistent with the stylized facts, which focuses on the reduction in domestic shocks as the Japanese economy was transformed from a semi-industrialized economy to a mature industrial state. Section 3 presents the dynamic model and the methodology employed in analyzing the insulation properties of fixed and floating exchange rates in Japan. Section 4 presents the preliminary data analysis investigating the stationarity properties of the data and the results from the multivariate model estimation. Section 5 concludes the paper.

2. Flexible exchange rates and output stability in Japan

One explanation for the greater output stability in Japan since the mid-1970s has focused on a monetary regime change that was made possible in large part by the move to floating exchange rates [e.g. Hamada and Hayashi (1985), Fischer (1988)].[5] Under fixed exchange rates the growth rate of Japanese output, shown in fig. 1, exhibited the largest fluctuations in output among the major industrial countries, while under flexible rates Japan has had the smallest output fluctuations. There is also prima facie evidence that greater output stability in Japan is associated with the greater monetary independence provided by the move to flexible exchange rates: the marked reduction in Japanese money variability since the mid-1970s, shown in fig. 2,

[5] Hamada and Hayashi (1985), for example, state that: 'After 1973 Japan adopted flexible exchange rates, along with other industrial countries. Thus, monetary policy gained more autonomy and the Bank of Japan began to focus on the control of money' (p. 97).

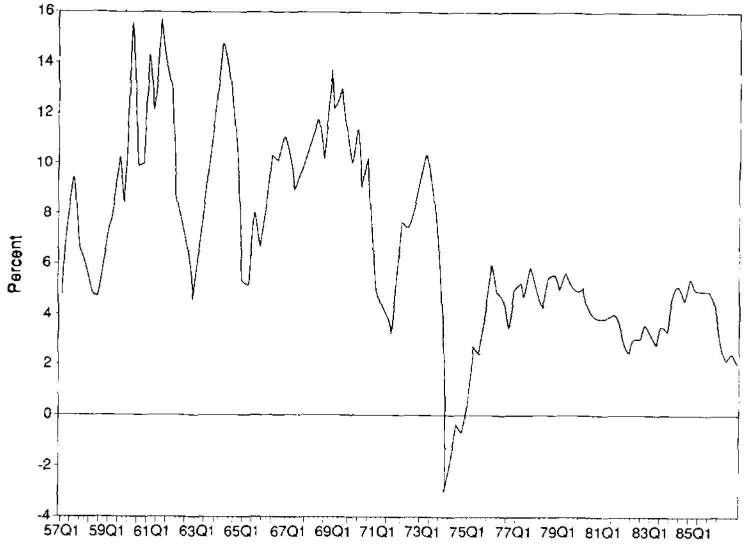


Fig. 1. Japanese real GNP growth.

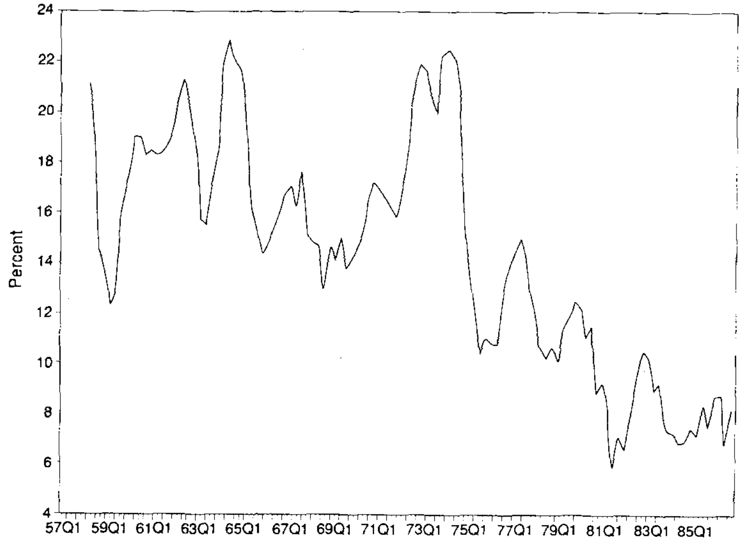


Fig. 2. Japanese money growth.

represents the predominant explanation for the dampened nature of the Japanese business cycle during the 1980s.[6]

Evidence linking monetary and output instability in Japan under the fixed rate regime to external imbalances (current account imbalances) is noted by Ackley and Ishi (1976), Hamada and Hayashi (1985), Fischer (1988) and others. However, the existing empirical evidence has not clearly identified the nature of the fundamental shocks disturbing the economy during the fixed-rate period nor has it distinguished causal linkages.[7]

An alternative explanation for the increased output stability in Japan focuses on the reduction in fundamental domestic disturbances, rather than the switch in exchange rate regime – a view that is consistent with the recent work emphasizing the degree of linkage between economies under flexible exchange rates, rather than their insulation properties. Specifically, the abrupt end of the Japanese 'high growth period' [Suzuki (1985)] in the early seventies – a period of double digit real growth rates and fundamental sectoral demand and supply shifts as well as technological innovations in the Japanese economy – may be primarily responsible for more stable output growth.[8] Under these circumstances, Japanese real output and monetary disturbances during the fixed-rate period may have been predominantly of domestic and not of foreign origin.

During the first three decades of the post-war period, Japan was a semi-industrialized country in the process of catching up with the developed countries of Europe and North America. During most of this period labor was an abundant resource, but there were shortages of capital equipment, raw materials and technical expertise. Post-war economic policy was therefore designed strategically to give priority to plant and equipment investment and to the export sector (used to secure imported raw materials). Elements of this policy included tight trade and exchange rate management, protection of industry, tax treatment favoring saving and investment, and credit subsidies to the priority sectors [Ackley and Ishi (1976)].

The rapid growth in the investment and export sectors of the Japanese economy during this period, and the consequent double-digit growth rates in overall output, is well documented. A major component of this growth, however, was the absorption of latent unemployment in agriculture and low-productivity traditional services into the rapidly growing investment and

[6]Taylor (1988) states unambiguously that 'The reduction in output and inflation variability in Japan is clearly associated with the reduction in monetary variability' (p. 26).

[7]Taylor (1988), however, suggests that U.S. monetary disturbances largely were responsible for Japanese (and German) monetary variability and output instability under fixed exchange rates.

[8]Suzuki (1986) emphasizes the distinction in development status: 'From Japan's viewpoint this was a historical watershed, symbolizing completion of its metamorphosis in status from semi-industrial nation to full industrial nation. The last spurt of postwar high growth was completed, and with it the process of catching up with the West in industrialization and modernization (p. 9).'

export sectors. Sectoral output and resource shifts, as well as rapid technological developments, characterize the Japanese economy during the 'high-growth' period. Denison and Chung (1976), for example, found that five factors contributed more to growth in Japan than other industrial countries: increased labor inputs, increased capital stock, advances in knowledge, reallocation of resources away from agriculture, and economies of scale.[9] Yoshikawa and Ohtake (1987) trace business cycles during this period to demand-induced fluctuations in the construction-related industries caused by population shifts. In contrast, the financial system and monetary policy played a fairly passive role during this period. The financial system during the high-growth period has been termed 'underdeveloped' [Suzuki (1985)]. In particular, the financial system was highly concentrated and extensively regulated to ensure resources flowed to priority sectors. Similarly, the stance of monetary policy during the period was generally expansionary, with the primary objective of supporting the rapid growth of the economy and to support high priority sectors with subsidized credit [Cargill and Hutchison (1988)].

The 'low-growth' period in Japan is usually identified as beginning in the early 1970s. By this time the greater part of the sectoral transformation of the Japanese economy was completed. The sectoral reallocation of resources stabilized, sectoral shifts dampened in magnitude and frequency, and average economic growth dropped to less than half its previous rate. Average real GNP growth in Japan was 9.1 percent per annum during the high-growth period (1956:2–1972:4), and 3.5 percent during the low-growth period (1973:2–1986:4).

Less dynamic sectoral transformation, a lower rate of productivity growth reflecting a larger per capita capital stock and the slowdown in technical progress as Japan 'caught up' with the West, are the primary reasons for the reduction in Japanese output growth and may be the primary reason for the reduction in output variability, rather than the change in exchange rate regime. Economic maturity may thus be responsible for the more stable, though less dynamic, Japanese economic environment and the reduction in domestic sectoral disturbances that predominated in the high-growth period [Lincoln (1988)].

3. Methodology

Different interpretations of the impact on Japan of the shift from a fixed to a flexible exchange rate system hinge crucially on whether the nature of the disturbances has changed or whether the exchange rate system has affected

[9]For example, agricultural employment as a percent of total nonresidential business declined from 37 percent in 1953 to 16 percent in 1971.

the manner in which Japan has adjusted to these disturbances. In particular, one potential explanation is that the decline in Japanese output variance since the mid-1970s is primarily attributable to the greater insulation properties of flexible exchange rates. An alternative explanation posited above suggests that the reduction in domestic disturbances, rather than either the exchange rate regime or dampened foreign disturbances, may be primarily responsible for greater output stability in Japan since the mid-1970s.

To shed light on these hypotheses, we examine a multivariate system that includes real GNP for both Japan (*JGNP*) and the United States (*USGNP*), the U.S. nominal money supply (*USM1*), and real oil prices (*OIL*) for both the fixed and flexible exchange rate periods. The objective is to examine the sources of the economic disturbances that have affected Japanese output by distinguishing between foreign and domestic shocks. Residual movements in Japanese real GNP not attributable to U.S. real GNP shocks, U.S. money shocks or oil shocks are interpreted as arising from domestic disturbances.

The greater difficulty in identifying exogenous domestic disturbances accounts for our asymmetric treatment of foreign and domestic shocks. This method of identifying residual movements in Japanese GNP with disturbances of domestic origin can potentially overstate the importance of domestic shocks. If there are foreign disturbances that affect output in Japan but are uncorrelated with oil prices, U.S. GNP, and U.S. M1, they will show up as domestic Japanese shocks in our framework. We speculate that this problem is potentially more serious during the flexible exchange rate period when foreign shocks unrelated to our control variables might affect the Japanese economy through exchange rate movements. Thus, our results may be most likely to overstate the role of domestic shocks during the flexible rate period. To provide some indication of the robustness of the particular decomposition we derive, we also estimated a system that included West German industrial production. Since our basic conclusions were unaffected by the introduction of this additional proxy for foreign disturbances, the results for the expanded system are discussed only in the footnotes.

In order to decompose the disturbances into their various components we must impose some restrictions on the multivariate dynamic system. Such identifying restrictions have taken a variety of forms in the recent literature. One approach achieves identification by imposing a priori restrictions on the contemporaneous interactions among the variables in the system. These restrictions normally take the form of exclusion restrictions, and in the context of VAR systems include the recursive structure popularized by Sims (1972) and the simultaneous equations approach used by Bernanke (1986), Blanchard and Watson (1986) and Walsh (1987).

An alternative approach to identification relies on restrictions on long-run effects implied by an underlying theoretical model. For example, in a system containing a set of real variables and the nominal money supply, long-run neutrality implies that variations in the level of the money supply should not have permanent effects on the levels of the real variables.[10] This can be translated into a restriction on the dynamic system that may aid in the identification of model parameters. This technique has been employed by Blanchard and Quah (1988), Jun (1988) and Shapiro and Watson (1988).

We borrow from both these approaches to estimate foreign and domestic sources of fluctuations in Japanese real output. The specific model can be represented by a $k \times 1$ vector of endogenous variables y_t (in this case a 4×1 vector comprising *OIL*, *USGNP*, *USM1* and *JGNP*) with Wold representation given by

$$y_t = B(L)\varepsilon_t, \tag{1}$$

where $B(L) = B_0 + B_1 L + B_2 L^2 + \dots$ is a $k \times k$ matrix of polynominals in the lag operator L and ε_t is a $k \times 1$ vector of while noise disturbance terms. We assume that B_0 has 1's along its diagonal and that $E\varepsilon\varepsilon' = \Sigma_\varepsilon$ is a diagonal matrix. It will be convenient to define the diagonal matrix P such that $PP' = \Sigma_\varepsilon$; the diagonal elements of P are the standard errors of the elements of ε. The variables in y_t may be in first difference form if necessary to ensure stationarity. The ε's are viewed as the fundamental structural disturbances, and we are interested in estimating the response of the elements of y to innovations in the elements of ε. For example, one element of ε represents the foreign (U.S.) monetary disturbance, and we are interested in the response of Japanese output to a foreign monetary shock and in the contribution of such shocks to Japanese output fluctuations under alternative exchange rate systems.

One way to summarize the sample information contained in our observations is to estimate the VAR representation of y_t:

$$H(L)y_t = u_t, \quad \text{where } H(0) = I.^{11} \tag{2}$$

Inverting the VAR representation yields $y_t = D(L)u_t$, where $D(L) = H(L)^{-1}$ and $D(0) = I$. In terms of (1), $D(L) = B(L)B(0)^{-1}$ and $u_t = B(0)\varepsilon_t$. Thus, in order to recover estimates of the structural disturbances, ε_t, from the estimated VAR residuals, u_t, it is necessary to estimate $B(0)$.

The covariance matrix of the VAR residuals, Σ_u, is related to $B(0)$ and Σ_ε by

[10] Variations in the stock of money generated via open market operations may be non-neutral due to the fiscal effects of the resulting changes in the government's interest payments [Sargent (1987)]. We assume any such effects are of secondary importance empirically and view our assumption of long-run neutrality as a good approximation to the predictions of most macro theoretical frameworks.

[11] That is, $H(L) = I + H_1 L + \dots$.

$$\Sigma_u = B(0)\Sigma_\varepsilon B(0)' = B(0)PP'B(0)'. \tag{3}$$

We have $k(k+1)/2$ bits of sample information in Σ_u to estimate the k^2 unknown elements in $B(0)$ and P; in general, $k^2 - k(k+1)/2 = k(k-1)/2$ additional restrictions are required for identification.[12]

The approach to identification pioneered by Sims assumes that $B(0)$ is lower triangular; economically, this is equivalent to assuming a recursive structure. Bernanke, Blanchard and Watson and Walsh also impose zero restrictions on $B(0)$, but do not necessarily require that a recursive structure be assumed. An alternative approach exploits the following relationship between $B(0)$, $B(L)$ and the lag polynomial $D(L)$:

$$D(L)B(0) = B(L). \tag{4}$$

Since $D(L)$ is estimable (it is just the lag polynominal obtained by estimating a standard VAR representation), a priori restrictions on $B(L)$ might allow $B(0)$ to be estimated. If $B(0)$ can be estimated, then estimates of the structural disturbances, ε_t, are given by $B(0)^{-1}u_t$, where u_t is the vector of VAR residuals.

As an example of the type of restrictions one might impose on $B(L)$, assume that economic theory implies that certain structural disturbances have no long-run impact on some elements of y. This imposes zero restrictions on the elements of $B(1)$.[13] In this case, the restrictions imposed by

$$D(1)B(0) = B(1), \tag{5}$$

together with the restrictions implied by (3), may allow $B(0)$ to be estimated.[14]

The model we estimate restricts the underlying joint moving average process y in two ways. First, we assume a structure of block exogeneity in which *OIL* is exogenous to the remaining variables, and the U.S. variables (*USGNP* and *USM1*) are exogenous to *JGNP*. In terms of the moving average representation (1) the first row of the matrix polynomial $B(L)$ takes the form $(b_{11}(L) \quad 0 \quad 0 \quad 0)$, while the second and third rows both contain zeros as their fourth element. Two additional restrictions arise from the

[12]This is clearly an order condition and is only necessary, but not sufficient, for identification.

[13]If $B(L) = B_0 + B_1L + B_2L^2 + \ldots$, then $B(1) = B_0 + B_1 + B_2 + \ldots$ equals the sum of the lag coefficients.

[14]For the models we consider, all matrices in (3) are 4×4. Thus, $B(0)$ contains 12 unknown off-diagonal elements while Σ_ε contains 4 unknown variances. The estimated VAR provides 10 bits of sample information (the number of unique elements in the 4×4 symmetric covariance matrix of the VAR residuals Σ_u). Six additional restrictions on $B(0)$ or $B(1)$ would be necessary to identify the remaining unknown elements of $B(0)$.

assumption of long-run monetary neutrality which implies that $b_{23}(1)=b_{43}(1)=0$, where $b_{ij}(L)$ is the ijth element of $B(L)$. Thus, evaluated at $L=1$, $B(L)$ has the form:

$$\begin{bmatrix} b_{11}(1) & 0 & 0 & 0 \\ b_{21}(1) & b_{22}(1) & 0 & 0 \\ b_{31}(1) & b_{32}(1) & b_{33}(1) & 0 \\ b_{41}(1) & b_{42}(1) & 0 & b_{44}(1) \end{bmatrix}. \tag{6}$$

This gives us seven restrictions and implies that the system is overidentified.

We carry out our actual estimation using a two-step procedure, following Jun (1988). Eq. (5) implies that $B(0)$ can be estimated as $D(1)^{-1}B(1)$ if $B(1)$ is known. $D(L)$ and Σ_u can be consistently estimated from the VAR system (2). Noting that $D(1)u_t=B(1)\varepsilon_t$, where $D(1)$ is the matrix of estimated long-run multipliers from the VAR,

$$\begin{aligned} S \equiv D(1)\Sigma_u D(1)' &= B(1)\Sigma_\varepsilon B(1)' \\ &= B(1)PP'B(1)' = K(1)K(1)', \end{aligned} \tag{7}$$

where $K(1)=B(1)P$. Our strategy is to use (7) to estimate $K(1)$. Letting $\boldsymbol{S}$ be the $k(k+1)/2\times 1$ vector of the stacked elements of the estimated matrix S defined by $D(1)\Sigma_u D(1)'$ and $\boldsymbol{K}$ be the vector of stacked elements of $K(1)$, we estimate the unknown elements of $K(1)$ by minimizing $(\boldsymbol{S}-\boldsymbol{K})'(\boldsymbol{S}-\boldsymbol{K})$. We then esitmate $B(0)P$ from $D(1)$ and $K(1)$.[15] Since the diagonal elements of $B(0)$ equal 1, we can recover estimates of $B(0)$ and P from an estimate of $B(0)P$.

Once we have obtained consistent estimates of $B(0)$, and therefore ε_t, we report the properties of the moving average representation in terms of impulse response functions and variance decompositions. Since we obtain consistent estimates of ε_t, we are also able to estimate the variances of the underlying structural disturbances.

4. Data and results

4.1. Preliminaries

Before estimating the model represented by eq. (1), we need to examine some of the characteristics of the joint process of *JGNP*, *USGNP*, *OIL* and *USM1* (all expressed in log form). In particular, as a preliminary analysis, we test for the existence of nonstationarity in the log levels of the four variables

[15]Recall that $B(0)=D(1)^{-1}B(1)$, or $B(0)P=D(1)^{-1}B(1)P=D(1)^{-1}K(1)$.

Table 2

Unit root tests.

	1956:1–1986:4	1956:1–1972:4	1973:1–1986:4
		Levels	
JGNP	$t_{\alpha^*} = -3.15^*$ $Z(t_{\alpha^*}) = -3.34^*$	$t_{\tilde{\alpha}} = -2.52$ $Z(t_{\tilde{\alpha}}) = -3.04$	$t_{\tilde{\alpha}} = -3.80^*$ $Z(t_{\tilde{\alpha}}) = -3.80^*$
OIL	$t_{\tilde{\alpha}} = -2.06$ $Z(t_{\tilde{\alpha}}) = -2.05$	$t_{\alpha^*} = -1.87$ $Z(t_{\alpha^*}) = -1.92$	$t_{\alpha^*} = -3.52^*$ $Z(t_{\alpha^*}) = -4.16^{**}$
USGNP	$t_{\tilde{\alpha}} = -1.88$ $Z(t_{\tilde{\alpha}}) = -1.92$	$t_{\tilde{\alpha}} = -2.40$ $Z(t_{\tilde{\alpha}}) = -2.34$	$t_{\tilde{\alpha}} = -2.58$ $Z(t_{\tilde{\alpha}}) = -2.34$
USM1	$t_{\alpha^*} = 3.72$ $Z(t_{\alpha^*}) = 7.16$	$t_{\alpha^*} = 3.66$ $Z(t_{\tilde{\alpha}}) = 5.56$	$t_{\alpha^*} = 3.23$ $Z(t_{\tilde{\alpha}}) = 4.69$
		First differences	
JGNP	$t_{\tilde{\alpha}} = -3.99^*$ $Z(t_{\tilde{\alpha}}) = -11.64^{**}$	$t_{\alpha^*} = -3.64^{**}$ $Z(t_{\alpha^*}) = -9.35^{**}$	$t_{\alpha^*} = -4.69^{**}$ $Z(t_{\alpha^*}) = -6.41^{**}$
OIL	$t_{\hat{\alpha}} = -4.39^{**}$ $Z(t_{\hat{\alpha}}) = -6.66^{**}$	$t_{\tilde{\alpha}} = -3.59$ $Z(t_{\tilde{\alpha}}) = -4.11^*$	$t_{\tilde{\alpha}} = -10.15^{**}$ $Z(t_{\tilde{\alpha}}) = -4.70^{**}$
USGNP	$t_{\alpha^*} = -4.86^{**}$ $Z(t_{\alpha^*}) = -8.16^{**}$	$t_{\alpha^*} = -3.81^{**}$ $Z(t_{\alpha^*}) = -6.33^{**}$	$t_{\alpha^*} = -3.12^*$ $Z(t_{\alpha^*}) = -5.10^{**}$
USM1	$t_{\tilde{\alpha}} = -3.49^*$ $Z(t_{\tilde{\alpha}}) = -8.89^{**}$	$t_{\tilde{\alpha}} = -4.77^{**}$ $Z(t_{\tilde{\alpha}}) = -5.95^{**}$	$t_{\tilde{\alpha}} = -2.18$ $Z(t_{\tilde{\alpha}}) = -6.03^{**}$

*Significant at the 5 percent level.

**Significant at the 1 percent level.

Notes: Critical values are from Fuller (1976).

Lag lengths were determined by Schwert's l_4 formula (4 for whole sample, 3 for subsamples).

of interest. Based on our finding that nonstationarity of the univariate processes cannot be rejected, we proceed to test for cointegration in the multivariate system consisting of all four variables.

Table 2 reports various statistics for testing the null hypothesis that each variable contains a unit root. $t_{\hat{\alpha}}$, t_{α^*}, and $t_{\tilde{\alpha}}$ are the augmented Dickey–Fuller statistics for the cases without a constant or trend, with a constant but no trend, and with a linear deterministic time trend, respectively. These statistics are equal to the reported t statistics for the coefficient on the lagged level in a regression of the first difference of each variable on its lagged level, lagged first differences, and, for t_{α^*} a constant and, for $t_{\tilde{\alpha}}$, a constant and a linear trend. Only the relevant statistic is reported. Corresponding to $t_{\hat{\alpha}}$, t_{α^*}, and $t_{\tilde{\alpha}}$, $Z(t_{\hat{\alpha}})$, $Z(t_{\alpha^*})$, and $Z(t_{\tilde{\alpha}})$ are nonparametric test statistics developed by Phillips (1987) and Phillips and Perron (1986). These statistics are based on transformations of $t_{\hat{\alpha}}$, t_{α^*}, and $t_{\tilde{\alpha}}$ and are discussed in Perron (1988). Critical values for the test statistics are given in Fuller (1976). The lag lengths were determined by applying Schwert's l_4 formula [Schwert (1987)].

For the entire sample period (1956:1–1986:4), the evidence is consistent with the hypothesis that *OIL*, *USGNP*, and *USM1* are difference stationary processes (around deterministic trends in the case of *OIL* and *USGNP*); that is, they are characterized by a unit root in levels but first differencing is sufficient to induce stationarity. We can reject the hypothesis that *JGNP* is nonstationary at the 5 percent level based on t_{α^*} and $Z(t_{\alpha^*})$. For Japanese real GNP, however, the results from the entire sample period may be spurious since there was an apparent shift in the underlying trend process associated with the transition from the high-growth to low-growth periods. When the sample is split at 1972:4, the end of the fixed exchange rate period, deterministic trends in the regressions are now significant and, for the earlier period, both the $t_{\tilde{\alpha}}$ and $Z(t_{\tilde{\alpha}})$ statistics fail to reject nonstationarity around this trend. For the 1973:1–1986:4 period, the test statistic just equals the 5 percent critical value. However, when the test statistic is recalculated for other lag lengths, the failure to reject a unit root is more convincing. For example, when the first difference at lag 4 is added, the ratio of its coefficient to its estimated standard error exceeds 2 and the test statistic $t_{\tilde{\alpha}}$ becomes -2.06. For both subsamples, the hypothesis of a unit root in the first difference of *JGNP* is easily rejected.

For *OIL*, *USGNP*, and *USM1* there is less a priori reason for splitting the sample period. For completeness, however, we report in table 2 the results for these variables for the fixed and flexible exchange rate periods. For the earlier period, 1956:1–1972:4, *USGNP* and *USM1* exhibit behavior consistent with difference stationary processes. For the 1973:1–1986:4 period, U.S. real GNP and *OIL* continue to exhibit behavior consistent with a difference stationary process. However, unit roots in the first differences of *OIL* during the 1956:1–1972:4 period and *USM1* during the 1973:1–1986:4 period cannot be rejected using $t_{\tilde{\alpha}}$ but are rejected using $Z(t_{\tilde{\alpha}})$. We place the most weight on the results from the entire sample period for *OIL*, *USGNP*, and *USM1* and proceed under the maintained hypothesis that all the variables are difference stationary. In terms of eq. (8), y_t will denote the joint process $((1-L)OIL_t\ (1-L)USGNP_t\ (1-L)USM1_t\ (1-L)JGNP_t)$.

Before specifying a joint process in the first differences of the four variables of interest, it is necessary to determine whether the nonstationarity in levels the four variables exhibit might be due to a smaller number of common stochastic trends. If economic fluctuations are the result of persistent technological shocks that are transferred across borders, for example, the unit roots in U.S. real GNP and Japanese real GNP would both reflect this common stochastic trend. In this case, the four-variable system might actually be characterized by only three underlying stochastic trends, and y_t would not have a VAR representation. If this is the case, there exists a cointegrating vector such that a linear combination of the levels of the four variables is stationary.

Table 3 presents three tests for the existence of a cointegrating vector among the four variables in the system. Row 1 reports augmented Dickey–Fuller statistics (i.e. $t_{\tilde{\alpha}}$) for the residual from a regression of *JGNP* on the

Table 3

Test for common trends.

Variables: *JGNP, OIL, USGNP, USM1*			
Test	1956:1–1986:4	1956:1–1972:4	1973:1–1986:4
(1) ADF	−3.35	−2.24	−2.29
(2) $\hat{Z}_\alpha$	−17.38	−13.84	−17.34
(3) Stock–Watson: $\lambda^{\mathrm{c}}_{*}(4,3)$	−21.84	−19.98	−17.96

*Significant at the 5 percent level.

**Significant at the 10 percent level.

[a]Critical values are −4.11 (5 percent) and −3.83 (10 percent). (Phillips and Ouliaris, 1990, table IIb).

[b]Critical values are −32.06 (5 percent) and −27.58 (10 percent). (Phillips and Ouliaris, 1990, table Ib.)

[c]Critical values are −47.00 (5 percent) and −42.00 (10 percent). (Stock and Watson, 1988).

other three variables.[16] If the variables are cointegrated, such a regression should pick out their stationary linear combination [see Engle and Granger (1987)]. Based on the critical values in Phillips and Ouliaris (1990), the test fails to reject the null hypothesis of no cointegration.

Row 2 presents the $\hat{Z}_\alpha$ test of Phillips (1987). Like the augmented Dickey–Fuller statistic, this is a residual-based test that Phillips and Ouliaris show has the advantage that it diverges faster with the samle size than $t_{\hat{\alpha}}$ under the alternative of cointegration. Critical values are reported in Phillips and Ouliaris. The test statistics for the whole sample period and each of the subsamples are consistent with the null of no cointegration.

Row 3 of table 3 is based on the work of Stock and Watson (1988) who have proposed a procedure for testing the hypothesis that a $k \times 1$ vector has $j<k$ distinct unit roots versus the alternative that it has only $m<j$ unit roots. The Stock–Watson test involves transforming the vector so that the first $k-j$ components correspond to the stationary components and the last j correspond to the integrated components. Under the hypothesis that there are j stochastic trends, the eigenvalues of the first-order autoregression of the j integrated components should be equal to 1. Under the alternative, only $m<j$ of the eigenvalues should equal 1. Consequently, under the alternative, the $(m+1)$st largest eigenvalue, λ_{m+1}, should be less than 1. The test statistic is given by $T^*(\lambda_{m+1}-1)$, where T is the sample size; critical values are reported in Stock and Watson. The test results are for the null hypothesis that the four variables in our model have four common stochastic trends versus the alternative that there are only three. The value of the test statistic for all periods is well below the reported critical values; we cannot reject the

[16]The test conclusions were unaffected when the other variables were made the dependent variable.

Table 4

Japanese real GNP variance decomposition.

	Percentage of error due to:							
	OIL		*USGNP*		*USM1*		*JPNGNP*	
Quarters ahead	Fixed	Flex.	Fixed	Flex.	Fixed	Flex.	Fixed	Flex.
1	2	2	9	44	0	31	89	23
5	15	3	8	44	1	24	76	29
10	19	2	10	49	0	17	71	32
15	20	2	11	52	0	13	69	33
20	20	2	11	54	0	10	69	34
30	21	2	11	57	0	7	68	35

null. Based on the results in table 3, we proceed under the assumption that there exists a VAR representation in the first differences of the four variables.

4.2. Estimation results

Table 4 reports the forecast error variance decompositions of Japanese real GNP. This allows a decomposition of the variance into that attributable to foreign disturbances (*OIL*, *USGNP*, and *USM1*) as opposed to domestic disturbances under the two exchange regimes. The first column under each variable refers to the fixed exchange rate period ('fixed') and the second column refers to the flexible exchange rate period ('flex'). Estimation periods were 1957:4–1972:4 for the fixed regime and 1974:4–1986:4 for the flexible regime.

The forecast error variance decomposition results indicate a sharp contrast in the importance of different types of disturbances during the fixed and flexible exchange rate regimes. Under flexible exchange rates, foreign shocks contributed much more significantly to the forecast error variance of Japanese real GNP than they did under fixed rates at both short- and longer-term horizons. Domestic shocks clearly dominate output variance in the earlier period, but were of secondary importance in the period of flexible rates since 1973.

Specifically, under fixed rates the model suggests that the proportion of Japanese real GNP variance associated with foreign shocks ranged from only 11 percent in the short term (2 percent attributable to *OIL*, 9 percent to *USGNP* and less than 1 percent to *USM1*) to 32 percent (21 percent *OIL*, 11 percent *USGNP* and less than 1 percent to *USM1*) in the longer term, where the longer-term horizon is defined here as 30 quarters.[17] In the flexible rate period, foreign shocks accounted for 77 percent of *JGNP* variance in the

[17]Since long-run monetary neutrality is imposed in the estimation procedure, the contribution of *USM1* is constrained to go eventually to zero.

Table 5
Estimated variances of the structural disturbances.

Shock	Fixed (1)	Flexible (2)	Ratio (1)/(2)
OIL	0.0167	0.0362	0.46
USGNP	0.0084	0.0059	1.42
USM1	0.0043	0.0073	0.59
JGNP	0.0142	0.0047	3.02

short term (2, 44 and 31 percent, respectively, for *OIL*, *USGNP* and *USM1*) and 65 percent in the longer-run (2, 57 and 7 percent, respectively, for *OIL*, *USGNP* and *USM1*).[18]

In contrast to the view expressed by Taylor (1988), it appears that shocks of foreign origin played a much smaller role in generating Japanese output instability during the period of fixed exchange rates than during the period of flexible rates. Under fixed rates, domestic shocks (unexplained variance) account for between 68 percent (at the 30 quarter forecast horizon) and 89 percent (at a 1 quarter ahead forecast horizon) of the variance in Japanese real GNP. This result is consistent with the hypothesis that major sectoral shifts and the rapid industrial transformation of the Japaneses economy during the greater part of the fixed exchange rate period were primarily responsible for the large observed output fluctuations, and that foreign factors played a secondary role.

This result does not lead necessarily to the conclusion that the fixed exchange rate regime helped insulate the Japanese economy from foreign shocks, however. Domestic shocks may simply have dominated foreign shocks in Japan during the fixed rate period because of the nature and pattern of the shocks facing the economy, rather than due to the form of the exchange rate regime. To examine this issue we calculated the estimated variances of the disturbances under the two exchange rate regimes. These are presented in table 5. The last column gives the ratio of the variance of each shock under fixed rates to its variance under flexible rates.

Not surprisingly, the variance of real oil shocks in the pre-1973 period of fixed rates is estimated to have been less than half of that in the period since 1973 during the flexible rate regime. The variance of U.S. monetary shocks also was significantly smaller in the pre-1973 period. Real U.S. GNP shocks appear somewhat larger in the fixed rate period, however. Nonetheless, by far the largest change occurred in the variance of domestic Japanese shocks.

[18]Similar results were obtained in the system that included West German industrial production as an additional source of foreign disturbances; the proportion of Japanese GNP variance associated with foreign shocks in the long run was 39 percent during the fixed rate period and 74 percent during the flexible rate period.

During the period to 1973, the variance of such disturbances was three times its value in the period since the shift to flexible rates. Similar conclusions were implied by the expanded system that included industrial production for West Germany as an additional proxy for foreign disturbances.

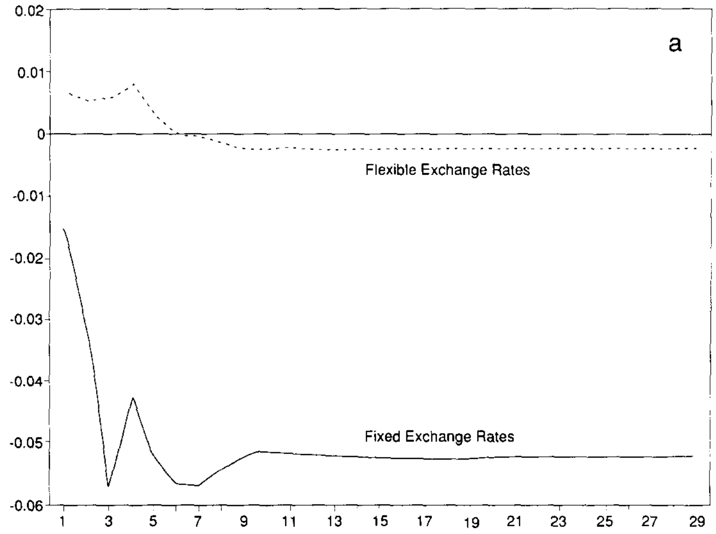
These results indicate that the nature and magnitude of the underlying disturbances facing the Japanese economy under the two exchange rate regimes were quite different. They also suggest that our finding that relatively more of the forecast error variance of Japanese real output under flexible exchange rates is attributable to foreign shocks may not reflect on the relative insulation properties of the two exchange rate regimes, but simply be attributable to the greatly reduced magnitude of the underlying domestic disturbances. This finding is consistent with our general hypothesis concerning the dynamic transition of the Japanese economy in the immediate post-war period.

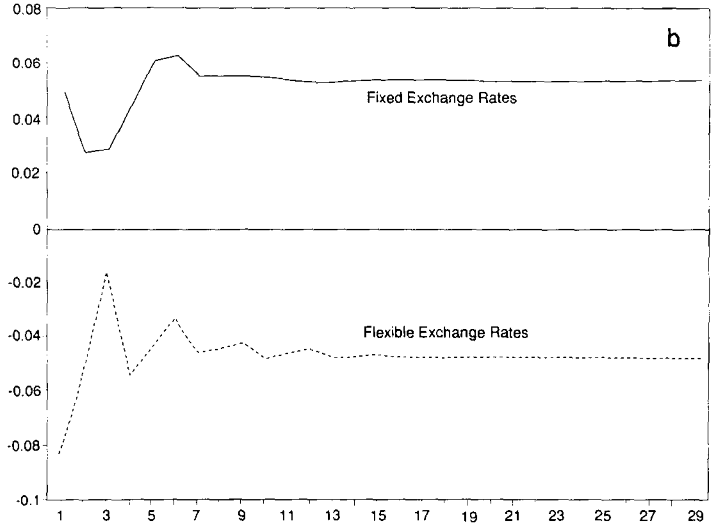
To investigate explicitly the insulation properties of exchange rate regimes, we calculated the responses of the level of Japanese real GNP under the fixed and flexible exchange rate regimes to one unit shocks in *OIL* [Fig. 3(a)], *USGNP* [Fig. 3(b)], and *USM1* [Fig. 3(c)]. We also report the impulse response function results for a unit domestic shock to *JGNP* [Fig. 3(d)]. This allows us to examine the extent to which disturbances have been dampened or exacerbated by the nominal exchange rate regime. Investigating the response of Japanese GNP to a one unit shock to each of these variables, rather than the more typical experiment with a one standard deviation shock, allows us to compare the insulation properties of the two exchange rate regimes when facing disturbances of equal magnitude.

Figs. 3(a)–(d) clearly indicate that the exchange rate regime plays an important role in the profile of the real output response to various shocks, whether foreign or domestic in origin. In particular, the impulse responses under the two regimes indicate that flexible rates do insulate the domestic economy from foreign shocks to a much larger extent than does a fixed rate regime. This finding is consistent with the theoretical prediction that a flexible rate regime has superior insulation features. It is at odds, however, with the recent work of Baxter and Stockman (1989) who argue that the form of the nominal exchange rate regime has not played a role in distinguishing the profile of post-war business cycles for most industrial countries.[19]

Specifically, fig. 3(a) indicates that after several quarters a real oil price shock causes a permanent decline in the level of Japanese real output. The effect under the fixed rate regime is significantly greater, however. The Japanese output response to a real U.S. GNP shock (not induced by

[19]Baxter and Stockman (1989) suggest that the changing nature of the shocks the world economy faces is the fundamental determinant of the evolution of output, a point consistent with our general findings for Japan.


a
0.02
0.01
0
-0.01
-0.02
-0.03
-0.04
-0.05
-0.06
Flexible Exchange Rates
Fixed Exchange Rates
1 3 5 7 9 11 13 15 17 19 21 23 25 27 29


b
0.08
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-0.08
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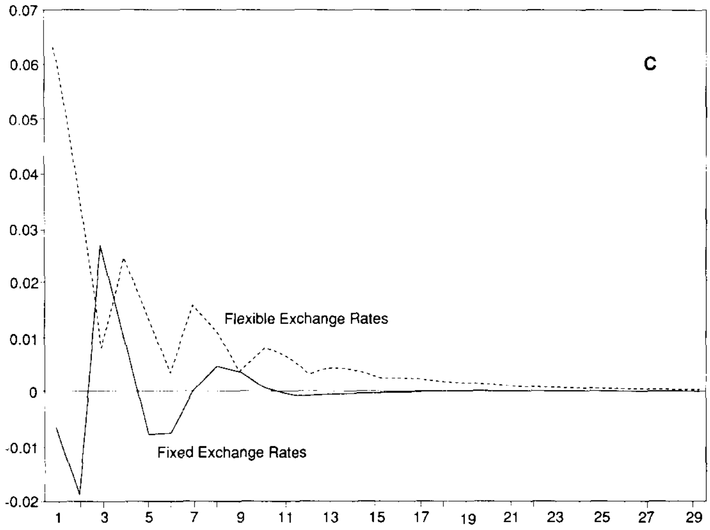
Fixed Exchange Rates
Flexible Exchange Rates
1 3 5 7 9 11 13 15 17 19 21 23 25 27 29



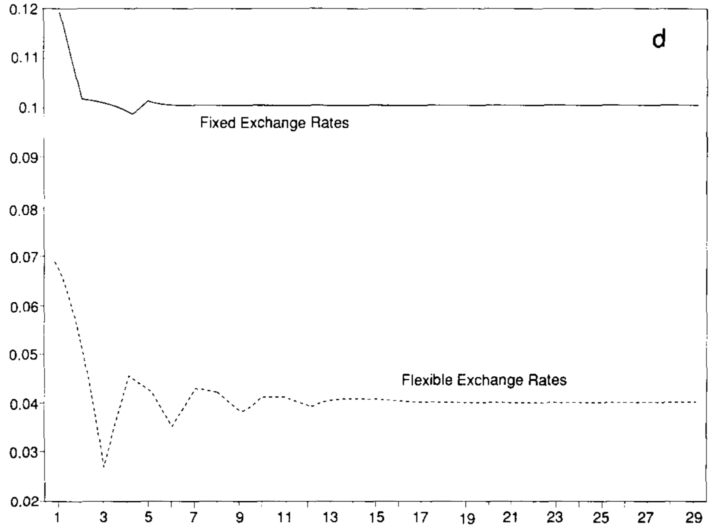


Fig. 3. (a) GNP response to oil shock. (b) GNP reponse to U.S. GNP shock. (c) GNP response to U.S. M1 shock. (d) GNP response to domestic shock.

monetary factors), shown in fig. 3(b), is long-lasting under both exchange regimes, but again the transmission under fixed rates is larger in absolute magnitude. Only in the case of a U.S. monetary shock, shown in fig. 3(c), is the absolute magnitude of the responses roughly similar in both regimes. Finally, consistent with the earlier findings, fig. 3(d) suggests that flexible exchange rate regime serves to dampen domestic shocks in Japan relative to similar disturbances under the fixed rate regime.

Figs. 3(a)–(c) provide some insights into the transmission mechanism by which foreign shocks affect Japanese real economic activity.[20] While fig. 3(a) shows the expected negative effects of an oil price shock under fixed exchange rates, there is almost no impact under flexible rates. In contrast, the impact of a U.S. GNP shock on Japan's GNP under flexible rates is virtually a mirror image of its impact under fixed rates. One possible interpretation of the negative effect of such a shock when exchange rates are flexible is that the resulting appreciation of the dollar relative to the yen reduces Japanese economic activity by raising domestic wages or the prices of imported inputs. Fig. 3(c) shows that, under flexible exchange rates, positive U.S. monetary shocks tend to raise GNP in Japan temporarily, perhaps indicating that the effect of raising U.S. income on Japan's exports may dominate the impact of the accompanying dollar depreciation. While the figures are suggestive, a more detailed analysis of the possible channels of transmission would be needed in order to draw firm conclusions.[21]

5. Conclusion

Empirical work to date has not adequately discriminated between recent views questioning the ability of a flexible exchange rate regime to effectively insulate the domestic economy from foreign disturbances and reasonable alternative explanations for the observed business cycle linkages between nations. It is commonplace to note that the nature of the disturbances the world economy has faced in the 1970s and 1980s, in terms of relative frequency, magnitude and duration, has departed from the norm established in the 1950s and 1960s. It is quite possible that this difference has played a more important role in shaping economic fluctuations than has the change in exchange regimes.

Our empirical approach in addressing these issues is to examine the Japanese experience with fixed and floating exchange rates. The focus on Japan is motivated in part because it appears to fit the traditional theory

[20] We thank one of the referees for suggesting the usefulness of discussing the implications of figs. 3(a)–(c) in terms of the signs of the effects.

[21] See Bryant, Helliwell and Hooper (1989) and Frankel (1988) for detailed discussions of empirical estimates of the channels through which U.S. macroeconomic policies are transmitted abroad.

regarding the superior insulation properties of flexible exchange rates; since the move to flexible exchange rates in February 1973 output variance has fallen markedly, with one potential explanation attributing this to the insulation provided the economy (particularly the monetary authority) from economic disturbances of foreign origin. We argue, however, that a more apt characterization is simply that domestic disturbances were more frequent and their magnitude much greater during the period of Japan's transition to a mature industrial economy.

Supporting our basic hypothesis, the estimation results indicate that shocks of foreign origin generated approximately 10–30 percent (depending on the forecast horizon) of the observed Japanese output instability during the period of fixed exchange rates. Foreign shocks played a much larger role during the period of flexible exchange rates, however, accounting for approximately 65–75 percent of output volatility. Nonetheless, this result appears to be attributable to the nature of the shocks facing the Japanese economy rather than a failure of flexible exchange rates to insulate. In particular, the variance of domestic shocks was more than three times as large under fixed rates, while the variance of foreign shocks was significantly less. Consistent with traditional theoretical predictions, however, our results suggest that in the face of a one unit standardized foreign shock, flexible exchange rates in Japan are generally more effective in insulating Japanese real output.

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