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### Money, Wages and Inflation in Middle-Income Developing Countries

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#### Abstract

This paper examines the short-run links between money growth, exchange rate depreciation, nominal wage growth, the output gap, and inflation in Chile, Korea, Mexico, and Turkey, using a generalized vector autoregression analysis. Nominal historical wage shocks are shown to have an important effect on movements in inflation only in Mexico. Generalized impulse response functions show that a positive historical shock to nominal wage growth generates a transitory but significant reduction in output. Inflation increases in all countries, particularly Mexico. A positive shock to nominal money growth raises real cash balances on impact and exerts an expansionary effect on output, despite an increase in real wages.

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### SUMMARY

This paper examines the short-run links between money growth, exchange rate depreciation, nominal wage growth, cyclical movements in output, and inflation in four middle-income developing countries: Chile, Korea, Mexico, and Turkey, using generalized vector autoregression techniques. Variance decompositions show that wage movements are greatly influenced by their own shocks at short horizons, with inflation shocks playing a more important role at longer horizons in all countries except Korea. They also indicate that, at short forecast horizons, historical shocks associated with inflation largely explain movements in inflation. Nominal exchange rate depreciation also plays a substantial role in all countries except Korea. Shocks associated with wages have little impact except in Mexico.

Impulse responses show that an innovation in the rate of growth of nominal wages increases inflation in all countries—an effect that displays considerable persistence over time in Mexico and Korea. Real wages also increase in all countries. Output, however, falls in some cases and rises in others, reflecting conflicting demand- and supply-side effects, associated with the rise in wages. A fall in the rate of depreciation of the exchange rate leads to a reduction in inflation on impact in all countries. In Chile and Turkey, the shock is associated with a transitory increase in output (despite a persistent increase in real wages), whereas in Korea and Mexico output falls on impact. Finally, a temporary increase in the rate of growth of the nominal money stock has an expansionary effect on output, and only a short-lived impact on inflation in Chile, Korea, and Turkey. In Mexico the effect on output is expansionary but short-lived, whereas the effect on inflation is positive on impact and persistent. Nominal wage growth increases significantly in Chile, Mexico, and Turkey, but not in Korea. Real money balances in Korea, and to a lesser extent Mexico and Turkey, increase initially but decline over time.

## I. INTRODUCTION

Research on the sources of inflation in developing countries has often led economists to distinguish between the “fiscal” view and the “balance-of-payments” view of the inflationary process.<sup>1</sup> Advocates of the fiscal view, while supporting the notion that inflation and excessive monetary growth are closely related, emphasize that the root cause of monetary growth in developing economies is often found in large fiscal imbalances. In many of these countries, inefficient revenue collection procedures and limited development of financial markets tends to increase reliance on seignorage as a source of financing of budget deficits. Factors such as the adverse effect of inflation itself on the real value of the fiscal deficit (the Olivera-Tanzi effect) also tend to reinforce the link between fiscal deficits, money growth, and inflation. By contrast, proponents of the balance-of-payments view have emphasized the induced effect of exchange rate movements (in the presence of an accommodative monetary policy stance) on domestic prices. Such movements—an exchange rate depreciation, for instance—are often the result of adverse developments in the balance of payments. In turn, such developments may reflect excessive government spending or exogenous shocks, such as a sharp deterioration in the terms of trade.<sup>2</sup>

The role of wages in the transmission of fiscal and exchange rate impulses to prices, and of wage formation mechanisms in inflation persistence, have also received much attention in the literature. Liviatan and Piterman (1986), for instance, emphasized the role of wage indexation mechanisms in the transmission of exchange rate movements to domestic prices.<sup>3</sup> Theoretical and empirical research on the experience of countries like Chile in the early 1980s and more recently Brazil has shown how backward-looking wage indexation can contribute to inflation inertia (Agénor and Montiel, 1996). More importantly, the “wage” or “cost push” view suggests that exogenous shocks in the

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<sup>1</sup> See Montiel (1989) for a more thorough discussion of these views, and the difficulty of drawing analytical distinctions between them.

<sup>2</sup> A sizable empirical literature has also emphasized the possibility (again in the presence of an accommodative monetary policy stance) of vicious circles between inflation and public sector deficits, on the one hand, and between the exchange rate and prices, on the other. An inflation-budget deficit spiral may emerge if inflation itself has an adverse effect on the fiscal deficit, by raising expenditure faster than revenue—as emphasized by the Olivera-Tanzi effect. Likewise, a devaluation-inflation spiral may occur if, following an initial exchange rate adjustment and an increase in domestic prices, policymakers attempt to maintain the real exchange rate constant to prevent adverse effects on competitiveness and external accounts. See, for instance, the evidence for Turkey discussed by Önis and Oznuçur (1990).

<sup>3</sup> Liviatan and Piterman’s argument is that the frequency at which nominal wages are adjusted tends to increase with the inflationary pressures generated by exchange rate movements—thereby raising inflation itself.

wage bargaining process, for instance, could exert independent impulse effects on inflation, which could persist over time in the presence of an accommodative monetary policy. In the New Structuralist tradition, an increase in the real wage demanded by workers, for instance, may trigger an unstable spiral between nominal wages and prices (Taylor, 1983). Money, in this view, plays a completely passive role.

Although the role of wages in the inflationary process is well recognized, few empirical studies focusing on developing countries have attempted to capture these effects. With the exception of Montiel (1989), none of the most recent studies—such as those of Dornbusch et al. (1990), Haan and Zelhorst (1990), Moser (1995), or Loungani and Swagel (1996)—based on either standard regression models or vector autoregression (VAR) techniques, tends to account for the impact of nominal wage growth on inflation.<sup>4</sup> Similarly, studies based on structural VARs, such as Hoffmaister and Roldós (1997), while capable of capturing indirectly the effect of wage shocks on prices through their impact on aggregate nominal shocks, do not identify the specific contribution of wages to nominal fluctuations. Understandably, the absence of a wage variable in some empirical studies may result from the lack of relevant data (particularly in those studies focusing on a large group of countries), but it may ignore important information for characterizing the inflationary process in developing countries.

This paper provides an analysis of the short-run links between nominal wages and inflation in a group of middle-income developing countries, using an empirical model that also accounts for money growth, cyclical fluctuations in output (as emphasized recently by Coe and McDermott (1997)), and exchange rate depreciation.<sup>5</sup> The econometric methodology, which relies on recently-developed VAR techniques, is presented in Section II. A key advantage of the technique used here is that it does not require specifying a particular ordering of the variables—a requirement that often leads

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<sup>4</sup> Loungani and Swagel (1996), in particular, use a VAR model (which includes money growth, exchange rate depreciation, deviations of output from trend, oil prices, and inflation) to study the sources of inflation in a cross-country group of 51 developing countries during the period 1964-92. Their analysis indicates that money growth is the most important influence on inflation in countries with high average inflation rates. However, in countries with low average rates of inflation, past values of inflation appear to be the most significant element in explaining the rate of growth of prices. They attribute this inertia to factors such as inflationary expectations, or labor market rigidities and indexation schemes, but do not identify more specifically the role of wages.

<sup>5</sup> Note, however, that our empirical model does not account for other potential determinants of inflation, such as the degree of central bank independence (Alesina and Summers, 1993), and the degree of openness (Lane, 1997). These other factors typically change slowly over time, and thus cross-sectional or panel data analysis would be needed to ascertain their effects.

to problems in interpreting variance decompositions and impulse response functions in standard VAR models. The empirical results, based on quarterly data covering the period 1979-95, are presented in Sections III, IV and V. Estimates of a VAR model linking money growth, consumer price inflation, the rate of depreciation of the nominal effective exchange rate, cyclical movements in output, and nominal wage growth are presented in Section III for four countries: Chile, Korea, Mexico, and Turkey. These countries cover a fairly large range of average inflation rates in the sample period, which varies from a low level of 7.7 percent in Korea, a moderate level of 18.9 percent in Chile, to the moderately high levels of 42.8 percent in Mexico and 55.9 percent in Turkey. Section IV focuses on variance decompositions and Section V on impulse response functions associated with three types of shocks: a wage shock, an exchange rate shock, and a money shock. The effects of an exchange rate shock (a reduction in the rate of depreciation of the nominal exchange rate) are illustrated in a simple analytical model with backward-looking wage contracts in Appendix II. Section VI summarizes the main results of the investigation, compares our results with some existing studies, and discusses some potentially fruitful extensions.

## II. METHODOLOGY

Our analysis of the links between money, wages, inflation, exchange rates and cyclical fluctuations in output uses the so-called generalized VAR analysis proposed by Koop, Pesaran and Potter (1996). This approach has the distinct advantage of identifying unique (order invariant) impulse responses and variance decompositions. To describe this methodology, we begin with a brief description of standard VAR analysis.

Consider the following VAR model:

$$C(L)y_t = \mu_t, \quad (1)$$

where  $y_t$  is a vector containing the  $k$  variables in the model,  $C(L)$  is a lag polynomial matrix with the VAR coefficients, and the vector  $\mu_t$  contains the VAR innovations with  $E[\mu_t] = 0$  and  $E[\mu_t \mu_t'] = \Omega$ . Direct examination of the VAR innovations can be misleading because these innovations are contemporaneously correlated. In other words, examining the effects of a single component of  $\mu_t$  while keeping the other components of  $\mu_t$  constant is not consistent with the historical information summarized in  $\Omega$  and thus would be misleading. That is, the VAR innovations do not embody the historical correlations of the shocks summarized by the covariance matrix,  $\Omega$ .

Sims (1980) proposed examining instead a transformed model that is obtained by

pre-multiplying equation (1) by a square matrix  $R^{-1}$ :

$$R^{-1}C(L)y_t = R^{-1}\mu_t, \quad (2)$$

which is such that the transformed innovations on the right hand side are orthogonal; that is, letting  $\varepsilon_t = R^{-1}\mu_t$ , the transformation is such that  $E[\varepsilon_t\varepsilon_t'] = I$ .<sup>6</sup> Direct examination of  $\varepsilon_t$  would be, at least in principle, meaningful. Note that the orthogonalization of  $\mu_t$  requires that  $R^{-1}\Omega R^{-1'} = I$  (or simply  $\Omega = RR'$ ), which led Sims to suggest that the matrix  $R$  be the (lower) Choleski decomposition of  $\Omega$ . Since the Choleski decomposition yields a (lower) triangular matrix, it is clear that VAR analysis stemming from the Choleski decomposition of  $\Omega$  describes a recursive system. Part of the uneasiness in the profession regarding this orthogonalization is associated with the fact that few economic models naturally lend themselves to a recursive form specification. Moreover, since the Choleski decomposition varies with the ordering chosen, VAR analysis along these lines does not lead to a unique set of results.<sup>7</sup>

Assuming that the system described by equation (2) is invertible, the impulse response function for the orthogonal shock  $\varepsilon_t$  is obtained by solving this system for  $y_t$ :

$$y_t = C(L)^{-1}R\varepsilon_t. \quad (3)$$

Typically standard VAR analysis presents the impulse responses for an explicit ordering chosen by recourse to economic judgement, and some effort is made to address the reader's natural concern regarding the robustness of the results, mostly by discussing results for alternative orderings. However, analysis of all possible alternative orderings is only feasible for VAR systems involving a relatively small number of variables. Moreover, even with small systems, results associated with alternative orderings are not always devoid of ambiguities.

The variance decomposition of  $y_t$  is obtained by splitting the mean square forecasting error (MSFE) into the portions attributed to each shock. Consider the expression for  $y_{t+s}$ , from equation (3):

$$\begin{aligned} y_{t+s} &= C(L)^{-1}R\varepsilon_{t+s} \\ &= A_0\varepsilon_{t+s} + A_1\varepsilon_{t+s-1} + A_2\varepsilon_{t+s-2} + \dots + A_s\varepsilon_t + A_{s+1}\varepsilon_{t-1} + \dots, \end{aligned} \quad (4)$$

that implicitly defines  $A(L) = C(L)^{-1}R$ . The (mean) forecast of  $y_{t+s}$  in period  $t$  is then:

$$y_{t+s|t} = A_s\varepsilon_t + A_{s+1}\varepsilon_{t-1} + \dots \quad (5)$$

<sup>6</sup> Note that for simplicity of notation, we have implicitly assumed unit variances for the transformed innovations. Since the structural innovations are contemporaneously uncorrelated, this amounts to choosing the scale of the innovations and has no bearing on the subsequent discussion.

<sup>7</sup> For a comprehensive treatment of standard VAR analysis, see Hamilton (1994).

where the expected values for  $\varepsilon_{t+s}$  for  $s > 0$  are equal to zero, that is, equal to their expected value in period  $t$ . From equations (4) and (5) it is clear that the forecasting error in period  $t$  is then:

$$y_{t+s} - y_{t+s|t} = A_0\varepsilon_{t+s} + A_1\varepsilon_{t+s-1} + A_2\varepsilon_{t+s-2} + \dots + A_{s-1}\varepsilon_{t+1}. \quad (6)$$

Note that equation (6) expresses the forecasting error of  $y_{t+s}$  using the first  $s$  terms in  $A(L)$ .

The MSFE is obtained by taking the expectation of  $[(y_{t+s} - y_{t+s|t})(y_{t+s} - y_{t+s|t})']$ .<sup>8</sup> To split the MSFE into the portion associated with each of the  $k$  shocks in  $\varepsilon_n$ , with  $n = t + 1, \dots, t + s$ , it is convenient to partition the square matrices  $A_i$  into  $k$  columns; let  $A_{ij}$  denote the  $j$ th column of the matrix  $A_i$ , and  $e_j$  the  $j$ th shock in  $\varepsilon_n$ , so that

$$A_i\varepsilon = A_{i1}e_1 + A_{i2}e_2 + \dots + A_{ik}e_k,$$

where, for notational convenience, the time subscript to  $\varepsilon$  has been omitted.

The MSFE can be expressed as the sum of  $k$  components, each associated with individual elements of  $\varepsilon$ , as follows:<sup>9</sup>

$$\text{MSFE}(y_{t+s}) = E \left[ \sum_{j=1}^k \{e_j e_j' (A_{0j}A'_{0j} + A_{1j}A'_{1j} + \dots + A_{s-1j}A'_{s-1j})\} \right]. \quad (7)$$

Note that because  $E[\varepsilon\varepsilon'] = I$ , that is, the variance of each shock  $e_j$  is unity and these shocks are orthogonal, the expectation term on the left hand side of equation (7) contains no covariance terms. Thus, the contribution of shock  $j$  to the forecasting error of  $y_{t+s}$  will equal  $(A_{0j}A'_{0j} + A_{1j}A'_{1j} + \dots + A_{s-1j}A'_{s-1j})$ , suitably scaled by  $\text{MSFE}(y_{t+s})$ . It should be clear that since  $A(L)$  depends on the ordering used to calculate the Choleski decomposition, the variance decomposition will vary with the ordering, as indicated earlier.

Cooley and LeRoy (1985), in the context of a bivariate system, and Bernanke and Blinder (1992), in a more general setting, argue that the identification strategy described earlier will recover the pure structural innovations only when the variable of interest is predetermined and is placed first in the ordering.<sup>10</sup> More recently Keating

<sup>8</sup> Note that the MSFE for each of the  $k$  variables in  $y_t$  are the diagonal elements of the expectation of the outer product.

<sup>9</sup> Since  $\varepsilon_t$  is assumed to be i.i.d., the time dimension has been dropped in equation (7).

<sup>10</sup> These issues contributed to the development of alternative identification strategies of structural innovations. These alternative strategies impose more restrictions based on economic theory (or economic structure) on the identification process. The so-called "structural" VAR approaches, beginning with Bernanke (1986), Blanchard and Quah (1989), and Shapiro and Watson (1988), have also come under attack by Faust and Leeper (1994) and Lippi and Reichlin (1993).



(1996) has discussed the general conditions needed for Choleski decompositions to be informative of the underlying structural innovations. Specifically, he has shown that the “appropriate” Choleski ordering will recover the structural innovations provided that the system is block recursive.<sup>11</sup> As noted before, our study follows an entirely different strategy based on the so-called generalized VAR analysis proposed by Koop, Pesaran, and Potter (1996).

Generalized VAR analysis is based on reconsidering what impulse response functions and variance decompositions are meant to uncover. Standard VAR techniques require that the analyst take a stand on how to recover and/or identify structural innovations, through the choice of the square matrix  $R$ . Koop, Pesaran and Potter (1996) argue that the whole notion should be re-examined, and proposed to change the focus from the “pure” structural shocks that are identified by orthogonalizing VAR innovations, to an understanding of what a typical *historical* innovation tells us regarding the dynamics of the model. These historical innovations are not necessarily orthogonal, but contrary to VAR innovations, embody the information regarding the contemporaneous correlation of these innovations.

Consider again equation (1) in moving average form:

$$y_t = C(L)^{-1}\mu_t, \quad (8)$$

and assume further that  $\mu_t$  is distributed multivariate normal, that is,  $N(0, \Omega)$ . Note that this implies that  $y_t$  is also multivariate normal with zero mean and covariance matrix  $C(L)^{-1}\Omega C(L)^{-1'}$ .<sup>12</sup> Rather than orthogonalizing the VAR innovations, generalized impulse responses (GIR) analysis considers the conditional expectation of  $y_t$  given a specific shock to  $\mu_t$ .

The “average” effect on  $y_t$  of a historical shock of say  $\mu_{it}$ , the  $i$ th component of  $\mu_t$ , can be obtained by taking the expectation of equation (8) conditional on the shock  $\mu_{it} = v$ ,

$$GIR(Y_t, \mu_{it} = v) = E[Y_t | \mu_{it} = v, \Omega] = C(L)^{-1}E[\mu | \mu_{it} = v, \Omega],$$

and given the properties of the multivariate normal distribution:<sup>13</sup>

$$GIR(Y_t, \mu_{it} = v) = C(L)^{-1}\Omega_i\sigma_{ii}^{-1}v, \quad (9)$$

<sup>11</sup> A block recursive system or a “partially recursive” structure is one where the structural equations in one block have a recursive ordering in that block and the contemporaneous covariance matrix for structural shocks is diagonal. See Keating (1996) for a complete discussion.

<sup>12</sup> In equation (8), all non-zero deterministic components have been implicitly subtracted from  $y_t$ .

<sup>13</sup> See Dhrymes (1978, pp. 362-67) for the explicit derivation of the conditional expectation of a multivariate normal.

where  $\Omega_i$  is the  $i$ th column of  $\Omega$ . Although  $v$  could be any value, it seems appropriate to set it equal to its historical value: the standard error of the  $i$ th shock,  $\sigma_{ii}^{1/2}$ . This choice for the value of  $v$  corresponds to a unit shock of the historical shock.<sup>14</sup>

In general, the GIR in equation (9) will differ from the standard impulse responses noted in equation (3). But since the first column of the Choleski decomposition of  $\Omega$  has the form

$$R_1 = [\sigma_{11}^{-1/2}, \sigma_{21}\sigma_{11}^{-1/2}, \sigma_{31}\sigma_{11}^{-1/2}, \dots, \sigma_{k1}\sigma_{11}^{-1/2}]', \quad (10)$$

the GIR for the  $i$ th shock will be numerically equivalent to the impulse response function obtained using the Choleski decomposition when the  $i$ th variable is put first in the ordering.<sup>15</sup> Also note that they will be numerically equivalent when  $\Omega$  is diagonal, that is, when the system is subject to shocks that are independent.

The generalized variance decomposition (GVD) can be derived following a procedure similar to that discussed earlier. Namely, using the expression for  $y_{t+s}$  and  $y_{t+s|t}$  from equation (8), the forecasting error can be expressed as:

$$y_{t+s} - y_{t+s|t} = C_0^{-1}\mu_{t+s} + C_1^{-1}\mu_{t+s-1} + C_2^{-1}\varepsilon_{t+s-2} + C_s^{-1}\mu_{t+1},$$

where  $C_0 = I$ .

To discuss the contribution of each shock, it is again convenient to partition the square matrices  $C_j^{-1}$  into  $k$  column vectors, such that  $C_j^{-1} = [C_{j1}^{-1}, C_{j2}^{-1}, \dots, C_{jk}^{-1}]$ , and denoting the  $j$ th element of  $\mu$  by  $\mu_j$ , then:

$$y_{t+s} - y_{t+s|t} = \sum_{j=1}^k \{\mu_j(C_{j0}^{-1} + C_{j1}^{-1} + C_{j2}^{-1} + \dots + C_{js}^{-1})\}. \quad (11)$$

Recall that the VAR innovation  $\mu_j$  does not embody the historical correlations of the shocks as summarized by  $\Omega$ , so that using directly equation (11) to calculate variance decompositions would be misleading. GVD proposes to look instead at the average portion of MSFE associated with each historical shock contained in  $\mu$ , conditional on the value of the shock equal to its standard error. Assuming as before that  $\mu$  is multivariate normal,

$$E[y_{t+s} - y_{t+s|t}] = \sum_{j=1}^k \{\Omega_j \sigma_{jj}^{-1/2} (C_{j0}^{-1} + C_{j1}^{-1} + C_{j2}^{-1} + \dots + C_{js}^{-1})\}. \quad (12)$$

<sup>14</sup> This methodology is extended to more general types of shocks and to nonlinear models in Koop, Pesaran and Potter (1996), and to cointegrated systems by Pesaran and Shin (1997).

<sup>15</sup> For an explicit derivation of the Choleski decomposition shown in (10), see Hamilton (1994, pp. 87-92).

This expression (squared) can be used to calculate the portion of the variance associated with each of the historical shocks suitably scaled by the  $MSFE(y_{t+s})$ . Note that although these historical shocks contain the information in  $\Omega$ , they are not typically orthogonal. Thus, “squared values” of equation (12) will not add up to the  $MSFE(y_{t+s})$ , so the typical GVD will not add up to 100 percent unless the historical shocks are orthogonal. Also, note that the GVD for  $y_{t+s}$  for the  $j$ th shock will be numerically equivalent to the Choleski variance decomposition when the  $j$ th variable is first in the ordering.

To conclude this discussion, it is worth reiterating that while GIRs and GVDs are numerically equivalent in some specific instances to the standard impulse responses and variance decompositions, they are nonetheless conceptually different constructs. GIRs are intended to tell the analyst how the model behaves conditional on the historical information to a specific shock. Likewise, GVDs are intended to provide the “share” of the movements of a series to historical shocks. Neither GIRs nor GVDs intend to uncover the effect of a pure structural shock as in standard VAR analysis, and thus are not typically orthogonal. In a sense, generalized VAR analysis provides “stylized facts” about the model that consistently accounts for the dynamics and historical correlations present in the data, that in turn may be contrasted with the predictions of a theoretical model.

The next three sections present empirical evidence for Chile, Korea, Mexico and Turkey on the importance of various shocks in explaining short-run movements in inflation—as well as wages and output gap. Using the generalized VAR analysis described above, we examine the importance of each shock using variance decompositions and discuss the impact of three innovations on the variables of the system: *a*) an increase in the growth rate of nominal wages; *b*) a reduction in the rate of depreciation of the nominal exchange rate; and *c*) an increase in the rate of money growth.

### III. VAR MODEL

The specific variables included in the VAR model are inflation  $\pi$ , broad money growth  $\hat{m}$ , nominal exchange rate depreciation  $\hat{e}$ , nominal wage growth  $\hat{w}$ , the output gap GAP (measured as actual output minus trend, as a percentage of output), a time trend, and an error-correction term in the equations in which long-run homogeneity

must be satisfied.<sup>16</sup> Because our study focuses on the short-run links between the variables under consideration, we have abstracted from considering explicitly longer-run determinants of wages and prices (beyond those captured by a time trend), such as productivity and product market structure. We have also chosen to impose long-run homogeneity through error-correction terms between the levels of *a*) nominal money balances and prices; *b*) nominal wages and prices; and *c*) the nominal exchange rate and prices. Put differently, we impose the assumption that nominal shocks have no long-run effect on real money balances, real wages, and the real exchange rate.<sup>17</sup> We impose each restriction by entering an error-correction term in the relevant equation. For instance, in the money equation, we enter (with a one-period lag) the difference between the logarithms of money and prices. Because of the small size of our sample, and the lack of power of standard bivariate cointegration tests, we chose not to test explicitly for the validity of these restrictions, and decided instead to impose them a priori on theoretical grounds.

Appendix I provides precise data definitions, together with the results of augmented Dickey-Fuller and Phillips-Perron unit root tests. These tests show that all the series used in this study are stationary. The VAR model was estimated using quarterly data over the period 1979:Q4 through 1995:Q4 for Chile and Mexico, 1982:Q1 through 1995:Q4 for Korea, and from 1980:Q4 through 1994:Q4 for Turkey. The results presented in this paper are based on a model that included two lags; the lag selection process is discussed in Appendix I. To control for seasonality, the VAR model includes a full set of seasonal dummies.<sup>18</sup>

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<sup>16</sup> The use of a broad monetary aggregate alleviates the problem of assessing the link between money and prices associated with shifts between liquid and quasi-liquid monetary assets. In the case of Korea, our measure of broad money may actually not be comprehensive enough to account for recent portfolio shifts across various components of quasi-liquid assets. Such shifts were particularly large in 1996 and 1997. However, our sample period ends in 1995 and therefore our results are unlikely to be affected. The inclusion of a time trend in the model broadly captures “longer-run” effects that may have occurred during the sample period and our not explicitly model in the VAR. For example, in Chile underwent significant structural reform, broad based deregulation and steady reduction of inflation during the sample period. The time trend tend will pick up these “low frequency” movements of the equilibrium real money balances.

<sup>17</sup> Note that since the output gap is stationary, nominal shocks do not have any long-run effect on that variable. Thus, the VAR model implicitly assumes that potential output growth is unaffected by nominal shocks. This assumption is could be relaxed but since the focus of this paper is the short-run it is not unduly restrictive.

#### IV. GENERALIZED VARIANCE DECOMPOSITIONS

Table 1 presents the GVDs for the main variables of interest, namely inflation, wages, and the output gap. At forecast horizons of less than a year, historical shocks associated with inflation itself play a primary role in the movements of inflation. In addition, historical shocks associated with exchange rate depreciation play a substantial secondary role (in excess of 30 percent) in all countries except Korea. This may reflect the fact that, with the notable exception of Korea and to a lesser extent Chile, these countries have used the nominal exchange rate as a nominal anchor in their stabilization efforts during the period under study. The importance of shocks associated with wages varies considerably in our sample of countries: it ranges from a small influence in Chile, Korea and Turkey, to a much larger one in Mexico. The evidence for Mexico could be reflecting the effect of the so-called *Pactos Sociales* implemented during that period, which included an important incomes policy component (Santaella and Vela, 1996). In the case of Chile, the lack of significance of wages, and the importance of inflation shocks themselves, may be a reflection of the high degree of indexation prevalent during the sample period. Perhaps the most intriguing result is that historical shocks associated with money growth have at best a tertiary importance—even at long forecast horizons—in explaining movements in the rate of inflation in these countries.<sup>19</sup> It is possible that this result is reflecting the fact that the effects of nominal money growth are captured by other historical shocks, because monetary aggregates are responding endogenously to other shocks. Finally, historical shocks associated with the output gap is a significant determinant of inflation at medium and long forecast horizons only in Mexico.

Wage movements are greatly influenced by their historical shocks. At a forecast horizon of less than a year, these shocks typically explain in excess of 70 percent of fluctuations in the growth rate of nominal wages. Beyond a year, however, historical

<sup>18</sup> Various studies have found that external shocks, in particular oil price shocks, have played an important role in explaining developments in inflation in Korea; see, for instance, Corbo and Nam (1992) and Hoffmaister and Roldós (1996). Because our VAR model does not explicitly account for changes in oil prices, we also conducted a sensitivity test which consisted in estimating the model for Korea using a shorter sample (1985Q1-1995Q4). Our assumption is that restricting the sample in this way is sufficient to control for the (possibly lagged) inflationary effects of the second oil shock of 1979-80. The results obtained in this case for the impulse response functions were similar to those reported below. Those obtained for the variance decompositions displayed some differences, which are noted in the next section.

<sup>19</sup> This result appears to be at variance with those obtained by Dornbusch et al. (1990) for several developing countries, which suggest that the proportion of the variability of inflation accounted for by money growth innovations tend to increase over time.



shocks associated with inflation are important in all countries except Korea, explaining about 40 percent of wage movements in Chile and Mexico. In the case of Mexico, and even more so in the case of Korea, the output gap (which may be capturing fluctuations in labor demand) plays a significant role at longer horizons, explaining in the first case about 12 percent, and in the second about 22 percent, of the movements in wages. Historical shocks to exchange rate depreciation are important in Mexico and Turkey, but account for only a small percentage of the movements in that variable in Chile and Korea.

Output gap movements are also greatly influenced by their historical shocks. This is true both at a forecast horizon of less than a year, where they explain more than 80 percent of output gap movements, and at longer forecast horizons—except in Mexico where their importance declines to about 50 percent after two years. Interestingly, historical shocks associated with nominal exchange rate depreciation are important in explaining output gap movements at longer horizons in Turkey and particularly in Mexico, reaching about 10 and 30 percent respectively, within a year. Again, these results may be reflecting the impact of the exchange-rate based stabilization programs implemented during the estimation period in these countries. Note that, with the exception of Chile, historical shocks associated with inflation do not appear to be an important factor behind movements in the output gap. In Chile, these shocks explain about 10 percent of the movements in the deviations of output from trend at a forecast horizon of two years. Historical shocks to the nominal money growth rate have a significant impact only in Korea.<sup>20</sup>

## V. GENERALIZED IMPULSE RESPONSES

We now turn to an analysis of GIRs associated with the following shocks: wage growth, money growth, and exchange rate depreciation. The results focus on the following variables: inflation, output deviations from trend, nominal wage growth, and, for the first and third experiments, the level of the real wage. For the money growth shock, instead of the real wage, we consider the path of real money balances.

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<sup>20</sup> As noted earlier, the VAR model for Korea was also estimated using a shorter sample (1985Q1-1995Q4). The main differences with the results described above are that historical shocks associated with wages explain a larger share of inflation (up to 15 percent at a horizon of 24 quarters), historical shocks associated with nominal exchange rate changes explain a larger share of the output gap (up to 12 percent at 24 quarters), and historical shocks associated with inflation explain a much larger share of inflation (similar to the figure reported for Turkey).

GIRs and their one-standard error bands are illustrated in each case.<sup>21</sup> Recall that GIRs are obtained as the marginal distribution of each shock in the system and thus provide the dynamic responses of the variables in the VAR model that accounts for all of the information in the joint distribution. For instance, the GIR to a wage shock shows the evolution of variables in the model (wages, prices, exchange rate, and the output gap) corresponding to “historically correct” shock to wages that explicitly accounts for the historical occurrence of other shocks in the model. In this sense, these responses show what has happened historically when wages “move first,” that is, when wages shocks have occurred before other shocks.<sup>22</sup>

### A. WAGE SHOCK

As illustrated in Figure 1, an innovation in the rate of growth of nominal wages has a significant inflationary effect on impact in Korea, Mexico, and Turkey.<sup>23</sup> Over time, inflation displays considerable persistence in Mexico and, to a more limited extent, Korea. In Turkey, by contrast, the effect on inflation is short lived. In Chile the inflationary effect of the wage shock, both on impact and over time, is imprecisely measured. The path of real wages, consistent with the response of nominal wage growth and inflation, indicates that real wages increase significantly in all cases and return gradually to their initial level (except for Mexico, where the initial increase is followed by a temporary decline), as implied by the long-run homogeneity restriction.

The output effects of the wage shock are particularly interesting. As illustrated in the figure, on impact the wage shock is associated with a short-lived expansion in output (relative to its potential level) in Mexico and Turkey. In both countries, however, the expansion is followed by a contraction in activity, which is highly persistent in Mexico. In Korea, after a delay of about one quarter, output expands; the subsequent contraction is also short-lived. The case of Chile is the most striking: output expands

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<sup>21</sup> In all figures the dotted lines for the GIRs show one standard error band in each direction and are based on 1000 Monte Carlo replications. In each replication we sampled the VAR coefficients and the covariance matrix from their posterior distribution. From these replications we calculated the square root of the mean squared deviation from the impulse response in each direction. By construction, these bands contain the impulse response function but are not necessarily symmetric. See Kloek and VanDijk (1978) for details regarding the posterior distributions.

<sup>22</sup> This concept is related to that of the “temporal precedence” of Granger causality in the sense that the GIR shows the response of the system assuming that the corresponding shock is “first.” It should not, however, be interpreted as indicative of Granger causality.

<sup>23</sup> Throughout this discussion a “significant” change means that the interval defined by the error bands does not contain the value zero.



Figure 1. Generalized Impulse Responses to a (historical) Wage Shock

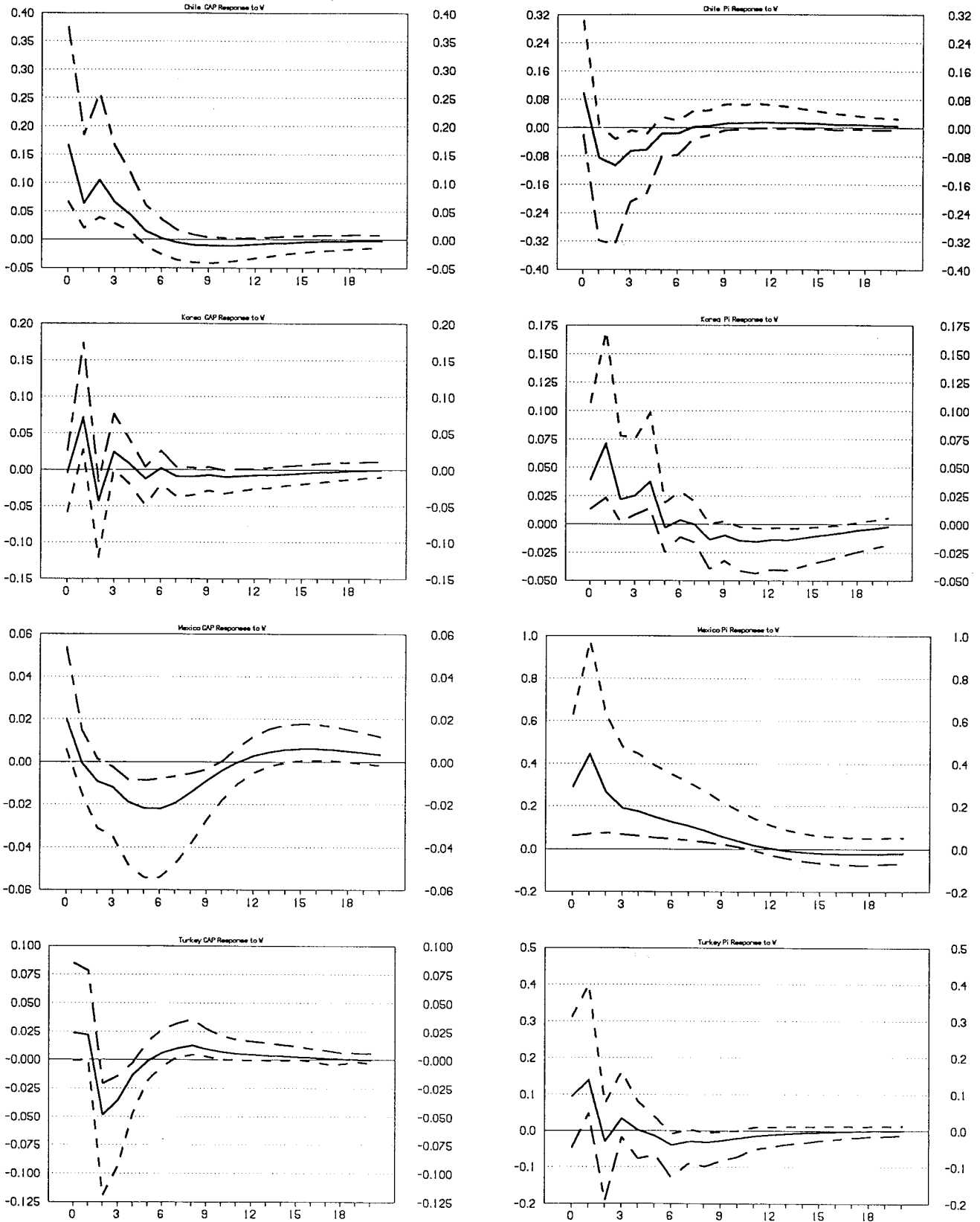
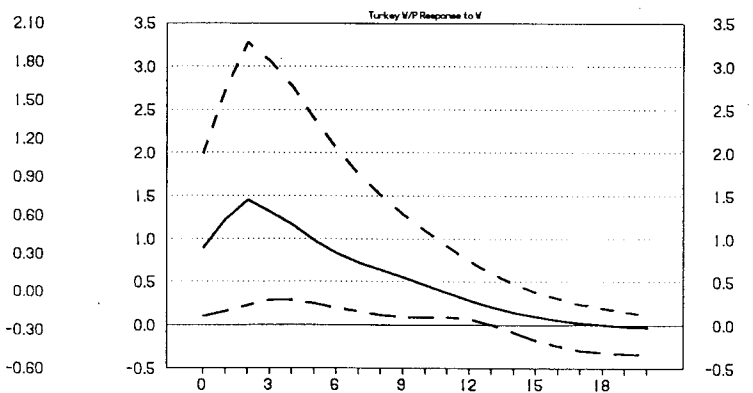
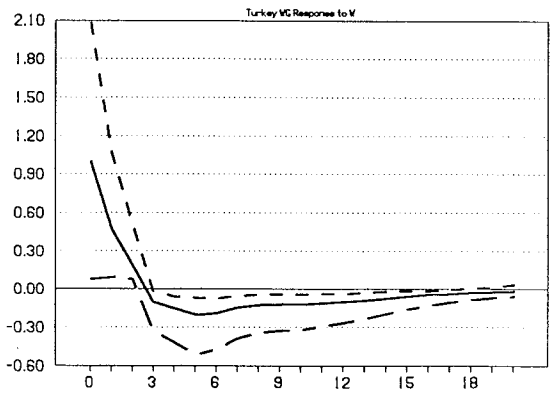
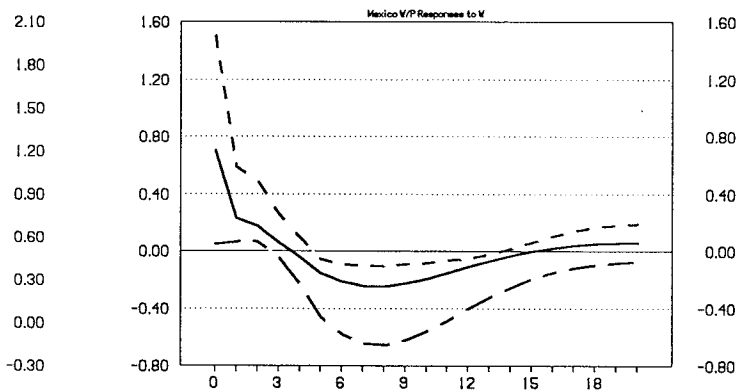
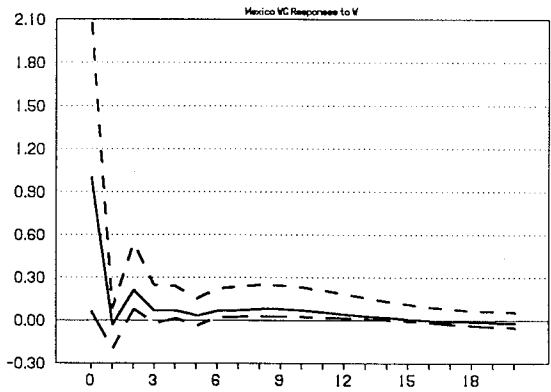
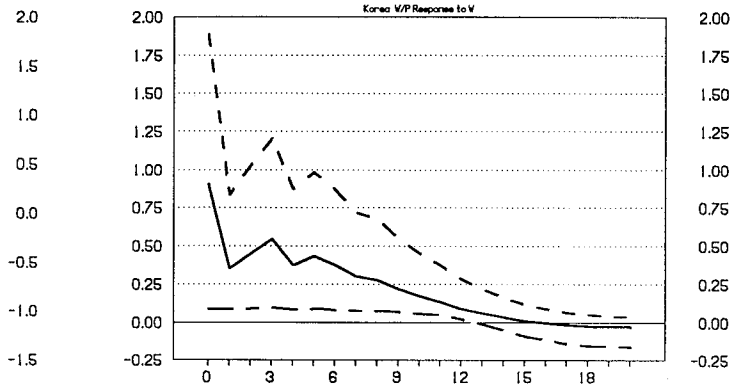
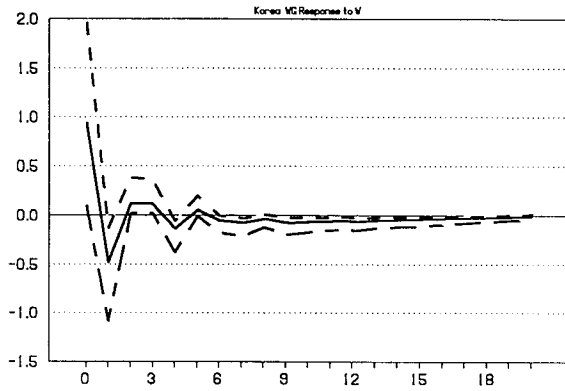
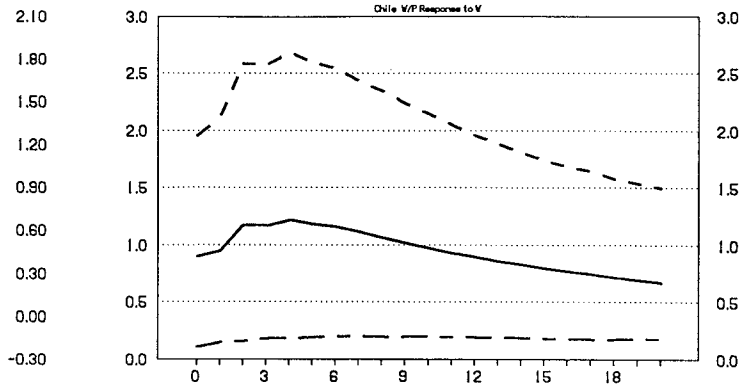
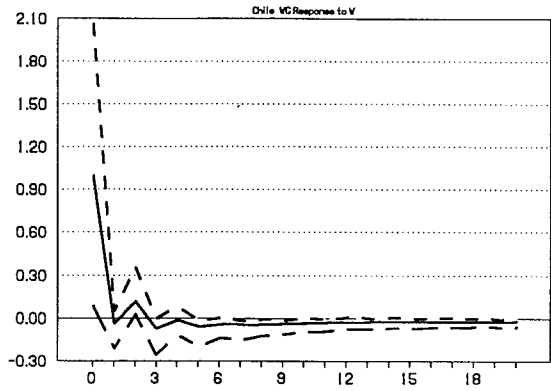


Figure 1 (Concluded). Generalized Impulse Responses to a (historical) Wage Shock



sharply on impact, and the expansionary effect lasts for about a year. There is no subsequent tendency for a recession to occur, as in the other countries.

The differences in output response across our sample of countries can be understood by contrasting the effect of supply and demand factors. As indicated earlier, an increase in the rate of growth of nominal wages is in all cases associated with a temporary increase in real wages—suggesting that inflation adjusts only with a lag to wage shocks. The increase in real wages, in turn, has supply- and demand-side effects. On the one hand, it tends to reduce the demand for labor and the supply of output. On the other, it increases disposable income, which tends (with a low marginal propensity to save) to stimulate aggregate demand.<sup>24</sup> The net effect on output, both on impact and over time, therefore depends on whether supply or demand factors dominate. The empirical results discussed above appear to suggest that the demand-side effect dominates on impact in the case of Turkey and Mexico, but that the supply-side effect comes into play fairly quickly. By contrast, in Chile (where the increase in real wages is the most persistent over time), demand factors not only dominate supply factors on impact, they continue to do so over the first few quarters.

## B. EXCHANGE RATE SHOCK

To understand the effects of an exchange rate shock in the present setting, it is worth beginning by briefly noting the analytical effects of a reduction in the nominal devaluation rate and the role of wage formation. Appendix II presents a simple model which illustrates the dynamics of real wages, output, and inflation in the presence of backward-looking nominal wage contracts. Essentially, the analysis shows that in the presence of such contracts, a permanent reduction in the nominal devaluation rate lowers inflation and leads at first to an increase in the real wage, followed by a gradual reduction over time, as contracts begin to reflect the lower path of the inflation rate. However, the initial increase in the real wage leads to a contraction in aggregate demand and output.<sup>25</sup>

The empirical effects of a negative (and temporary) exchange rate shock are illustrated in Figure 2. In all four countries, inflation falls significantly on impact and

<sup>24</sup> The supply-side effect is the one typically emphasized by orthodox economists. The demand-side effect (also known as the Keynes-Kalecki effect) has been advocated by New Structuralist macroeconomists, particularly Taylor (1983).

<sup>25</sup> As noted by various authors—see, for instance, Agénor (1996, p. 295)—the inverse relationship between the inflation rate and the real wage emerges only if the frequency of readjustments remains constant. If this frequency falls as a result of lower inflation, the correlation between prices and real wages may be positive.

Figure 2. Generalized Impulse Responses to a (historical) Exchange Rate Shock

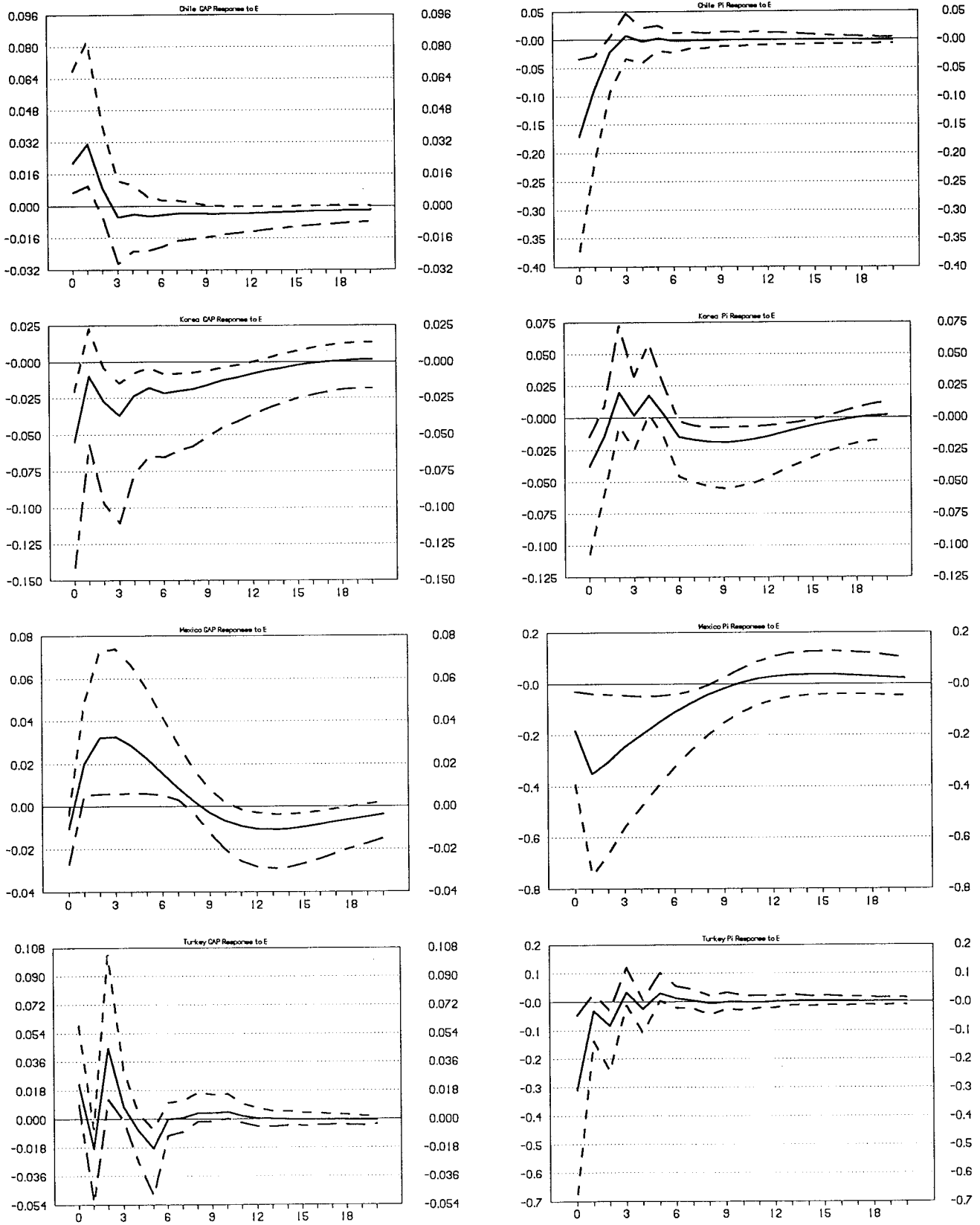
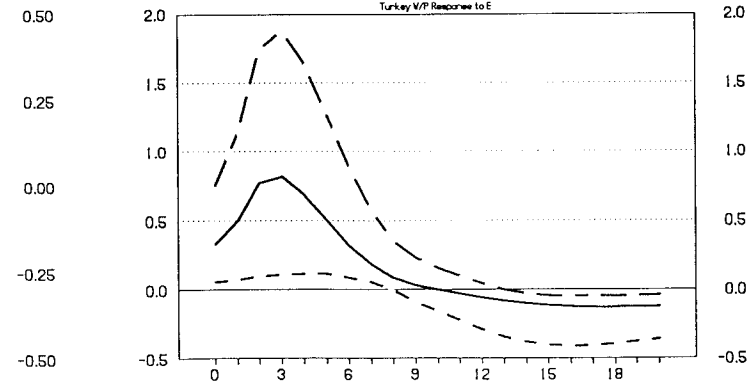
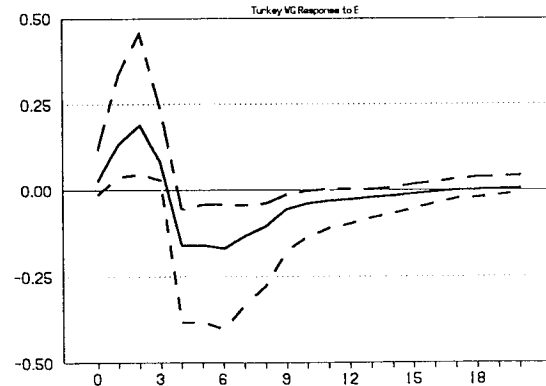
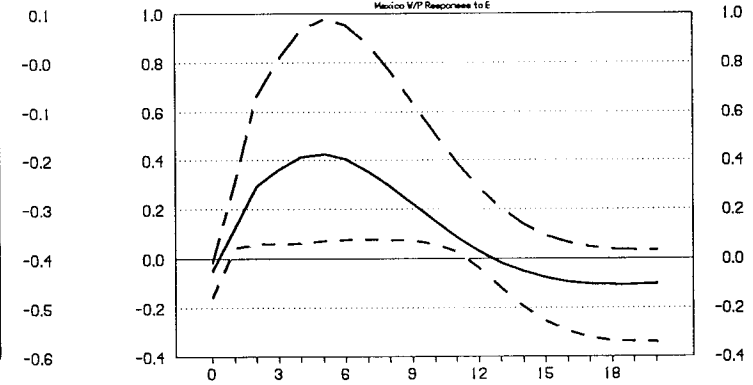
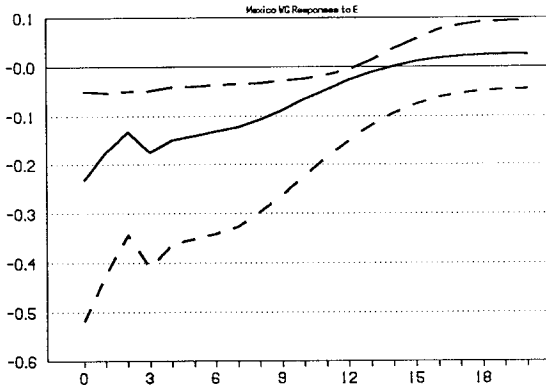
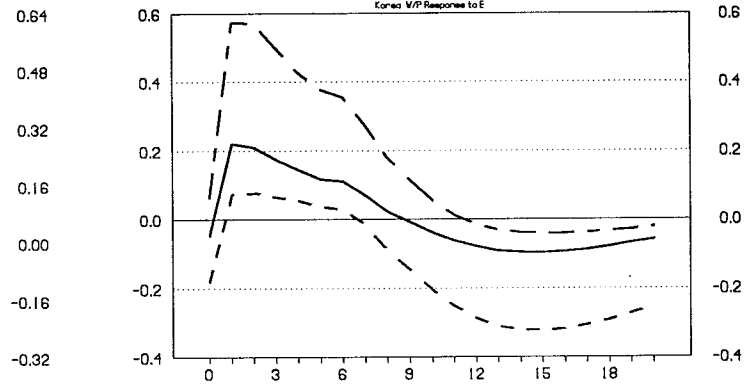
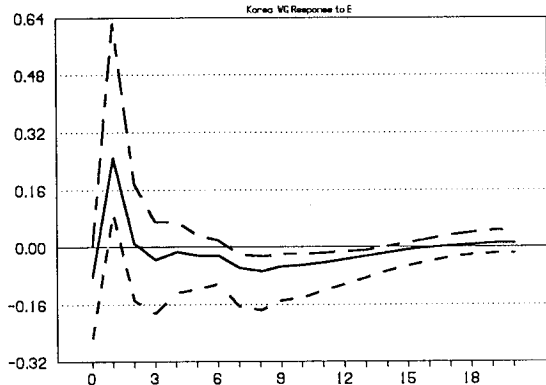
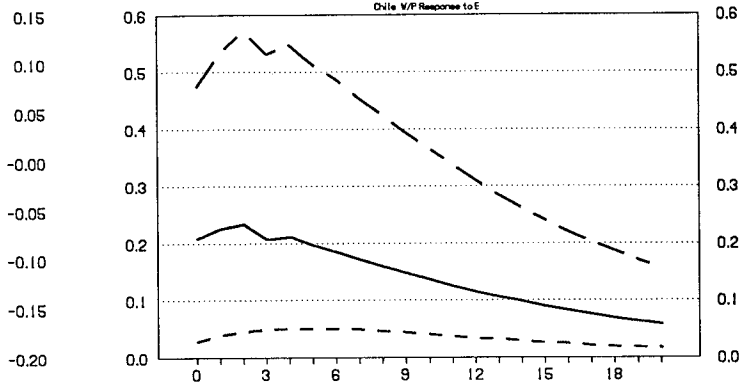
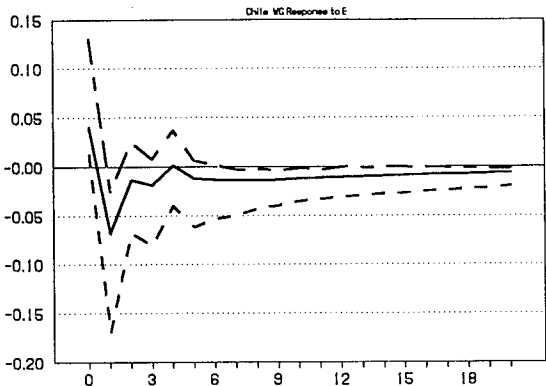


Figure 2 (Concluded). Generalized Impulse Responses to a (historical) Exchange Rate Shock



during the subsequent 2 quarters for Chile, Korea, and Turkey, whereas for Mexico, the reduction in inflation continues over a period of more than two years. As noted below, output increases in some cases, and falls in others. The initial reduction in inflation therefore indicates that the downward pressures on inflation associated with the reduction in the rate of depreciation of the nominal exchange rate (through a reduction in the price of imported inputs, for instance) dominate demand pressures—in those cases where output actually rises. This is consistent with our previous results on GVDs, which suggest that the relative importance of exchange rate shocks in the movements of inflation is large.

The rate of growth of nominal wages increases on impact in Chile and Turkey, falls in Mexico, with no discernible effect in Korea. One factor explaining the fall in wage growth in Mexico might be the sharp reduction in inflation, which would tend to moderate wage claims. For Chile and Turkey, factors explaining why nominal wage growth would increase in a situation where inflation is falling are not entirely clear. One possible explanation may be that the fall in inflation is perceived to be temporary (or to lack credibility), leading agents to anticipate a resurgence in inflation in the future and thus to demand higher wages now. This mechanism, however, cannot be properly tested in our methodological framework.

Real wages increase in Chile and Turkey on impact and, after a delay of one quarter, in Korea and Mexico. In all countries (particularly Chile), there is again considerable persistence over time. The path of real wages and output would appear to be consistent with the predictions of the model developed previously: in that interpretation, the drop in current output relative to potential would not be viewed as the result of a negative input price shock (associated with the reduction in the rate of nominal exchange rate depreciation), but rather the consequence of nominal wage inertia, which implies an increase in real wages induced by a reduction in inflation. This interpretation, in light of the GVDs discussed earlier, would appear to be particularly relevant for Mexico.<sup>26</sup>

On impact, a fall in the rate of depreciation of the exchange rate has a positive and significant output effect in Chile and Turkey. This expansionary effect occurs despite an increase in real wages, which again suggests the predominance in the very short run of the demand-side effect of changes in real wages, relative to the supply-side effect—as discussed earlier in the context of a nominal wage shock. However, in both countries, the expansion in output is short lived. In Korea and Mexico, output falls on impact; but whereas the initial contraction is quickly followed by an expansion

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<sup>26</sup> Note that Figure 6 shows a slight reduction in the real wage in Mexico on impact. However, it is barely significant.

in the case of Mexico (with output increasing significantly over a period of almost 6 quarters), in Korea, the contractionary effect persists over a period of about 3 years.

### C. MONEY SHOCK

The effects of a shock to monetary growth on inflation in the short run has received much attention in monetary economics, and is particularly interesting in the present context. An increase in the rate of nominal money growth has often been observed to result in an increase, on impact, in the economy's stock of real money balances (or, equivalently, to a reduction in velocity). A well-documented example is the case of Peru in the second half of the 1980s (Dornbusch et al., 1990, pp. 8-9). This empirical phenomenon has been explained as resulting from the existence of a lag between the change in the nominal supply of money and the response of inflation (or nominal income). An increase in the money growth rate, in the presence of these lags, would raise both nominal and real money balances since prices do not respond immediately.<sup>27</sup>

A temporary increase in the rate of growth of the nominal money stock is illustrated in Figure 3. In order to focus on the behavior of real money balances, we have replaced the real wage (shown in previous figures) with the response of the level of real money balances that is consistent with the responses of inflation and nominal money growth. The results indicate in all cases an expansionary effect on output, which persists in Chile and Korea. For instance, in Korea a positive monetary shock stimulates output over a horizon of 6 quarters.

The effects of the shock on inflation are fairly rapid and transient—except for Mexico. On impact, inflation rises significantly in Chile, Korea, and Mexico. In Turkey, whereas inflation rises sharply on impact, there is no discernible effect afterward.

Nominal wage growth responds significantly to the innovation in nominal money growth in Chile, Mexico and Turkey, but not in Korea. Real money balances seem to follow the predicted path discussed above (an initial increase followed by a gradual reduction over time) in all countries. Movements in real balances are more persistent in Chile and Mexico than in Korea and Turkey. In particular, in Korea there is no significant effect on real money balances after 2 years.

<sup>27</sup> A formal model explaining this behavior (and emphasizing the endogeneity of money supply and the frequency of price adjustment) is developed by Dornbusch et al. (1990). As noted by these authors, Cagan-type models of money demand with adaptive or rational expectations—as well as models of gradual adjustment of real money balances to their desired level—predict a reduction, rather than an increase, in real money balances, following an increase in the rate of growth of the nominal money stock.

Figure 3. Generalized Impulse Responses to a (historical) Money Shock

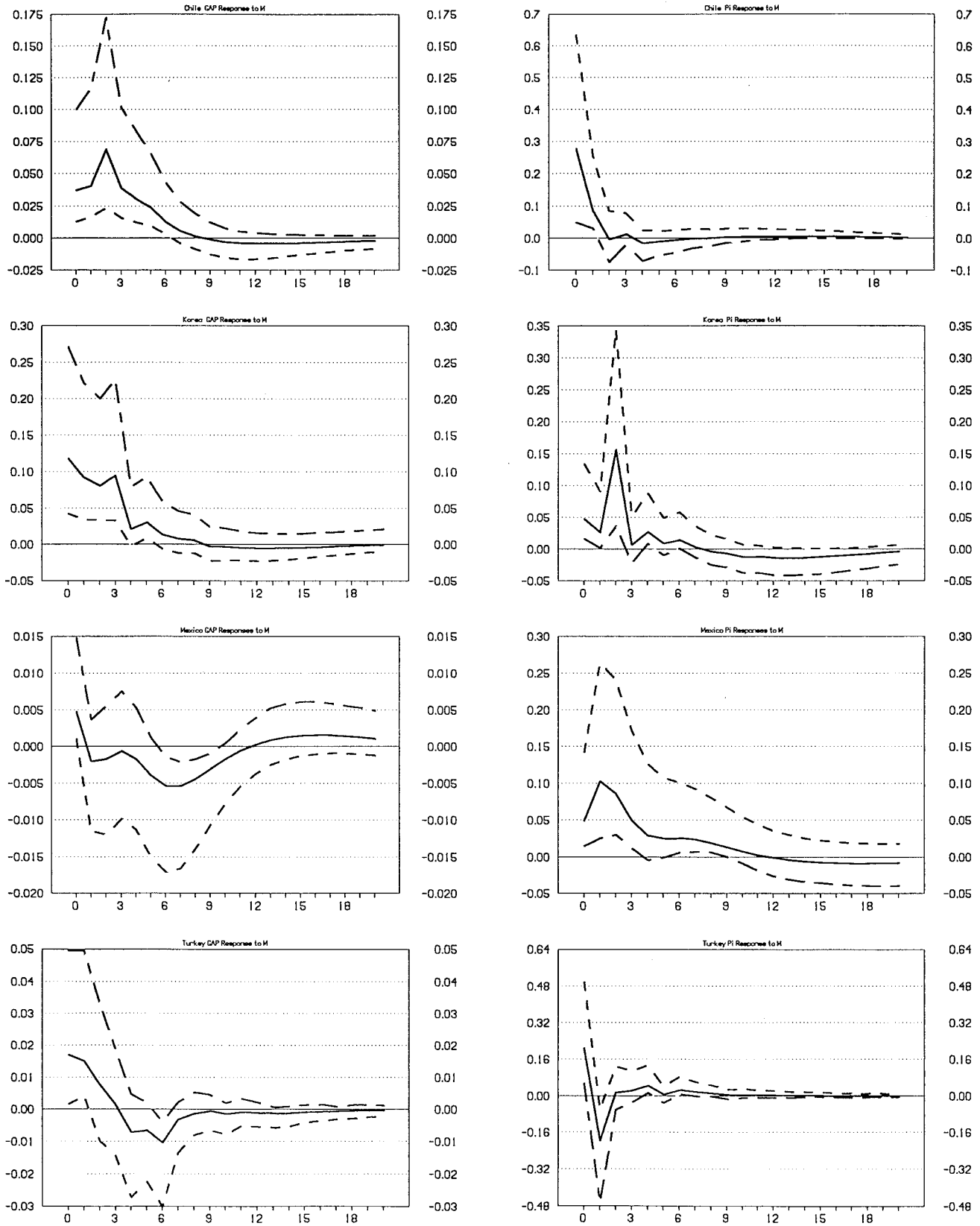
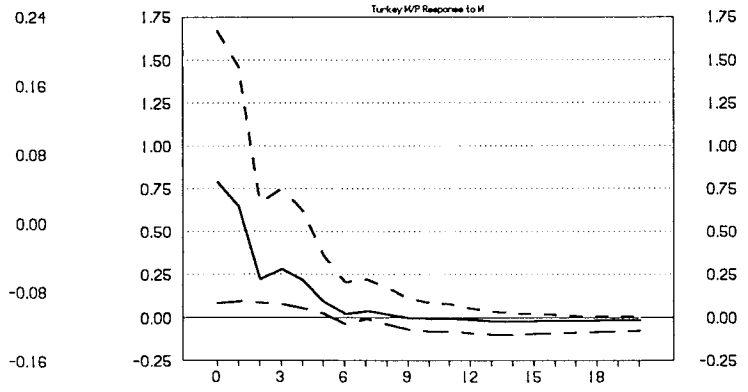
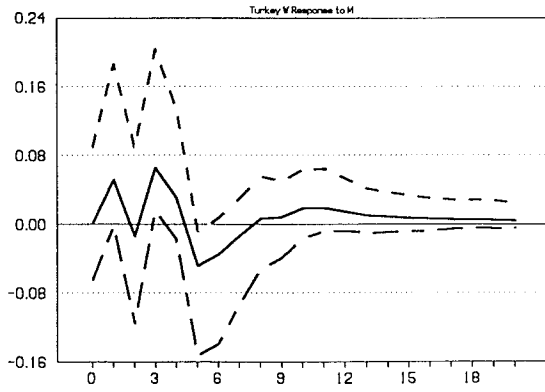
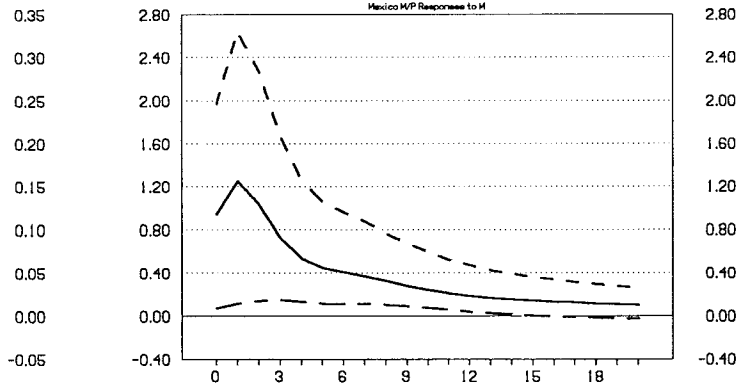
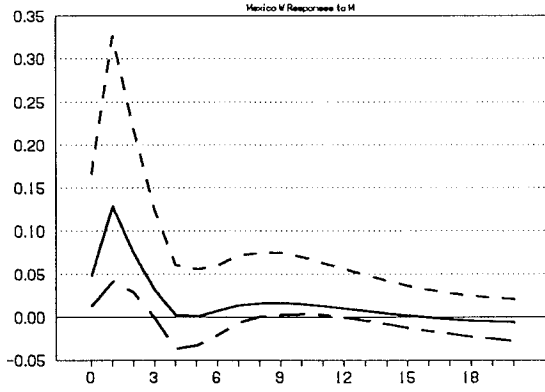
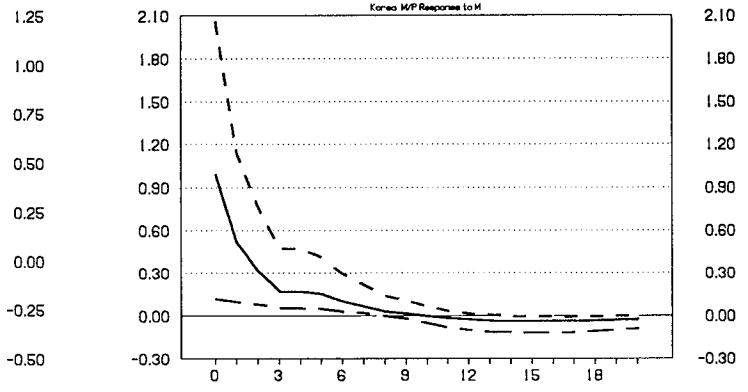
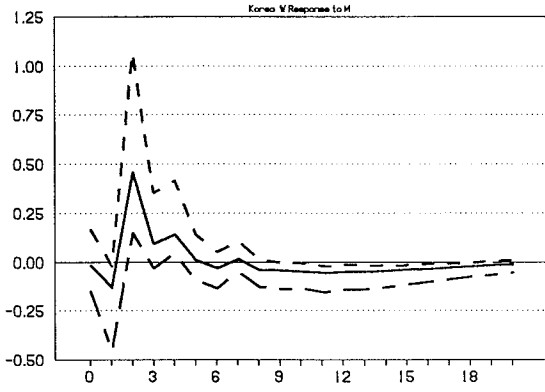
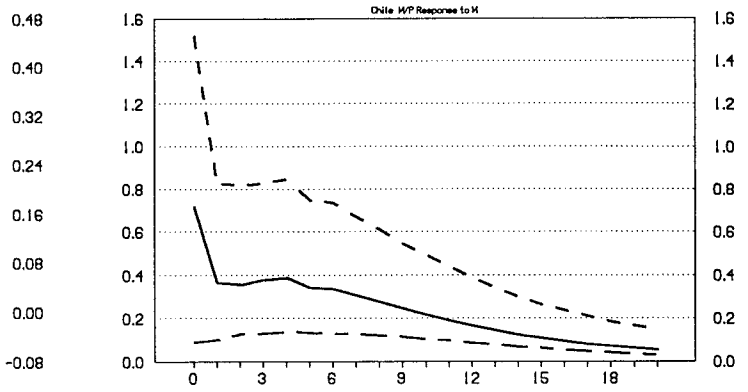
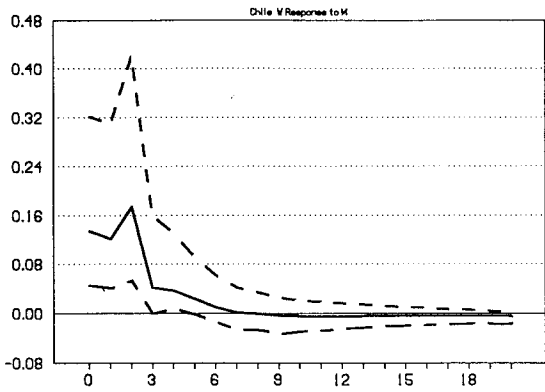




Figure 3 (Concluded). Generalized Impulse Responses to a (historical) Money Shock



## VI. SUMMARY AND CONCLUSIONS

The purpose of this paper has been to examine the short-run links between money growth, exchange rate depreciation, nominal wage growth, cyclical movements in output, and inflation in four middle-income developing countries: Chile, Korea, Mexico, and Turkey. The methodology was based on generalized VAR techniques. Variance decompositions showed, in particular, that wage movements were greatly influenced by their own shocks at short horizons, with inflation shocks playing a more important role at longer horizons in all countries except Korea. They also indicated that at short forecast horizons, historical shocks associated with inflation itself are the main factor explaining movements in inflation. In addition, nominal exchange rate depreciation also plays a substantial role in all countries except Korea. Shocks associated with wages have little impact in Chile, Korea, and Turkey, and a much larger influence in Mexico.<sup>28</sup> Shocks associated with the output gap are important determinants of inflation at medium and long forecast horizons only in Mexico; and shocks associated with money growth have at best a tertiary importance—even at short forecast horizons—in explaining movements in the rate of inflation in all countries except Mexico. As argued earlier, one possible interpretation of this result is that it may reflect the fact that, during the period under study, the effects of nominal money growth may have been captured by other shocks, because monetary policy was accommodative.

Generalized impulse responses showed that an innovation in the rate of growth of nominal wages generates an increase in inflation in all countries—an effect that displays considerable persistence over time in Mexico and Korea. Real wages also increase in all countries. Output, however, falls in some cases and rises in others, reflecting conflicting demand- and supply-side effects associated with the rise in wages. A fall in the rate of depreciation of the exchange rate leads to a reduction in inflation on impact in all countries—reflecting mostly the direct, downward inflationary pressures associated with the reduction in the rate of depreciation of the nominal exchange rate. In Chile and Turkey the shock is associated with a transitory increase in output (despite a persistent increase in real wages), whereas in Korea and Mexico output falls on impact. Finally, a temporary increase in the rate of growth of the nominal money stock has an expansionary effect on output, and only a short-lived impact on inflation in Chile, Korea, and Turkey. In Mexico the effect on output is

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<sup>28</sup> Montiel (1989) found that the dominant factor in the inflationary process in Israel over the period 1976–85 appears to have been nominal wage growth. To a lesser degree, this appeared to have been also the case over the same period for Argentina and Brazil.

expansionary but short-lived, whereas the effect on inflation is positive on impact and persistent over time. Nominal wage growth increases significantly in Chile, Mexico and Turkey, but not in Korea. Real money balances in Korea (and to a lesser extent Mexico and Turkey) follow the path often described in other empirical studies of monetary dynamics in developing countries, namely, an initial increase followed by a gradual reduction over time.

The thrust of the analysis suggests that accounting for wage shocks may be important for understanding the inflationary process in developing countries.<sup>29</sup> It is also interesting to note that for all the countries in the sample (with the exception of Korea) wages appear to react significantly over time to price shocks—reflecting perhaps backward indexation of nominal wages. By contrast, in Korea, for example, wages appear to react more to shocks in the output gap than to shocks in inflation.<sup>30</sup> These stylized facts may be suggesting that as developing countries mature into more advanced countries, business cycles factors may tend to exert greater influence on labor market and wages, making inflation more consistent with an output gap model of inflation. This tentative conclusion should not, of course, be construed as a general statement about the inadequacy of output gap models in developing countries, although our results for Chile are consistent with those obtained by Roldós (1997). In particular, Coe and McDermott (1997), using a methodology substantially different from ours, have provided evidence that the gap model works well in Asian countries—even in those with relatively low levels of development.

Our results suggest that accounting for the behavior of real wages is also important to understand the short-run fluctuations in output. Because wage movements tend to affect both the supply and the demand sides of the economy, the short-run effects of policy shocks on activity may be difficult to ascertain a priori. For instance, with backward-looking nominal wage contracts, a disinflation program based on a reduction in the devaluation rate may raise real wages sufficiently to entail (in the presence of a low elasticity of labor demand) an expansion in output—as illustrated by our results. A direct shock to wages may have similar effects. Such movements in output may or may not be compatible with the balance-of-payments objectives formulated in the context

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<sup>29</sup> Likelihood ratio test suggest that wages provide statistically significant information (at conventional significance levels) in the VAR model for Korea ( $\chi^2_8 = 27.3$ , significant at the 1 percent level) and Mexico ( $\chi^2_8 = 13.3$ , significant at the 10 percent level). In Chile, the likelihood ratio test suggest that they wages are “borderline” significant ( $\chi^2_8 = 12.3$ , significant at the 15 percent level), whereas in Turkey the test is not significant ( $\chi^2_8 = 9.8$ ).

<sup>30</sup> These results may reflect the limited impact of trade unions on aggregate wages in Korea, as documented in a variety of studies. See for instance Fields (1994).

of the program. The potential response of real wages must therefore be gauged at the design stage of a stabilization program; and whether or not such response may require the use of additional policy instruments needs to be considered.

### Data Sources and Unit Root Tests

The data used in this study are at a quarterly frequency and cover the period 1979:Q1 to 1995:Q4. The variables are measured as follows. Inflation  $\pi$  is the rate of increase in consumer prices.  $\hat{m}$  is the rate of growth of broad money. GAP is the ratio of current industrial output to potential output, estimated with the Hodrick-Prescott filter (see Hodrick and Prescott, 1997).<sup>1</sup>  $\hat{w}$  is the rate of growth of nominal wages in the manufacturing sector, and  $\hat{e}$  the rate of depreciation of the nominal effective exchange rate, as calculated by the IMF. All rates of growth are calculated at an annual rate. All data are obtained from the IMF's *International Financial Statistics*.

To determine the number of lags to include in the near-VAR model, we started by calculating standard lag-length tests, that is, Hannan-Quinn (HQ) and Schwarz. These tests compare the cost of increasing the lag length (reduced degrees of freedom) to the benefit (increased information extraction from the data). As noted in the text, the VAR models included a full set of seasonal dummies. Using a maximum lag length of six, for all countries the HQ test suggested using two lags and the Schwarz suggested using just one, with the exception of Chile where both tests suggested using one lag. We decided to err on the side of more information—albeit at a potential cost of less precision—and used two lags for all VAR models to ensure that information in the data was not ignored.

The time series properties of the series used in the VAR models are summarized in Table A1. Whereas the Augmented Dickey-Fuller (ADF) unit root test provides mixed evidence for a number of series, the Phillips-Perron (PP) test suggests that, with the exception of inflation in Mexico, the null hypothesis of a unit root can be rejected for all the series at conventional significance levels. For inflation in Mexico the ADF test marginally rejects the unit root hypothesis at a significance level of 10 percent. We proceeded, nonetheless, to include inflation in our VAR model for Mexico, due to the well known fact that the ADF and PP tests are typically not powerful enough to discriminate between a highly persistent series from a true unit root process.

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<sup>1</sup> A standard smoothing parameter equal to 1600 was used in the computations of the filtered series. See Agénor and Hoffmaister (1996) and the references therein for a more detailed discussion of the properties of the Hodrick-Prescott filter.

Table A1  
Order of Integration: Unit Root Test Statistics

		ADF test	PP test
	$k$	test statistic	
Chile			
$\pi$	4	-4.061**	-5.217***
$\hat{m}$	0	-8.867***	-8.993***
GAP	4	-5.173***	-6.259***
$\hat{w}$	0	-5.131***	-5.398***
$\hat{e}$	0	-4.905***	-4.814***
Korea			
$\pi$	2	-1.743	-4.641***
$\hat{m}$	3	-2.774	-8.783***
GAP	4	-4.775***	-4.354***
$\hat{w}$	3	-2.871	-8.907***
$\hat{e}$	0	-5.717***	-5.928***
Mexico			
$\pi$	0	-2.924	-2.701
$\hat{m}$	0	-5.961***	-6.562***
GAP	4	-4.521***	-3.697**
$\hat{w}$	3	-2.059	-8.578***
$\hat{e}$	0	-4.329***	-4.682***
Turkey			
$\pi$	0	-6.481***	-6.483***
$\hat{m}$	3	-3.773*	-11.824***
GAP	4	-3.705*	-13.104***
$\hat{w}$	0	-3.837*	-3.819*
$\hat{e}$	4	-5.471***	-7.258***

Notes: Variables are as defined in the text. The estimation period is 1979Q1-1995Q4 for Chile, and Mexico. For Korea, the estimation period is 1982Q2-1995Q4. For Turkey, the estimation period is 1980Q2-1995Q4 for all series except the output gap, for which it is 1982Q2-1995Q4. Asterisks \*, \*\* and \*\*\* denote rejection of the null hypothesis of a unit root at the 10%, 5% and 1% significance levels. Critical values are from McKinnon (1991).

### Disinflation, Output and Real Wages

This Appendix uses a modified version of the simple dynamic model presented by Agénor (1996) to analyze the effects of an exchange rate shock (a reduction in the nominal devaluation rate) on real wages, output, and inflation in the presence of backward-looking nominal wage contracts.

Suppose that the economy produces a nonstorable good, which is an imperfect substitute to the foreign good. Domestic output  $y$  is inversely related to the real product wage,  $\omega = w/p$ , where  $w$  denotes the nominal wage and  $p$  the price of the domestic good:

$$y = y(\omega), \quad y' < 0. \quad (1)$$

Consumption  $c$  depends positively on income and negatively on the expected *long-run* value of the relative price of the domestic good,  $z^*$ :

$$c = c(y, z^*), \quad 0 < c_y < 1. \quad (2)$$

$z^*$  must be consistent with the relative price for which the market for domestic goods clears in the long run:

$$c[y(\tilde{\omega}), z^*] = y(\tilde{\omega}),$$

from which we have, using (1):

$$z^* = c_z^{-1}(1 - c_y)y'\tilde{\omega} = \Phi(\tilde{\omega}), \quad (3)$$

with  $\Phi' > 0$ . Thus, an increase in the long-run value of the real wage, by increasing excess demand for the domestic good (since it reduces output supply by more than it reduces consumption), leads to an increase in the long-run expected relative price.

Suppose that changes in the rate of inflation are positively related to movements in the real exchange rate (which are determined by the difference between the devaluation rate and the prevailing rate of inflation) and excess demand for domestic goods:<sup>1</sup>

$$\dot{\pi} = \kappa(c - y) + \theta(\pi - \varepsilon), \quad (4)$$

where  $\kappa, \theta > 0$ .

With backward-looking wage contracts, the nominal wage  $w$  depends only on past levels of prices:

<sup>1</sup> The first effect may be associated with the impact of exchange rate changes on the domestic price of imported inputs and final consumption goods. This price equation, in which a reduction in the rate of depreciation (relative to the prevailing rate of inflation) lowers the rate of change of the inflation rate (rather than lowering inflation itself) was apparently first introduced by Dornbusch (1982).

$$w = \rho \int_{-\infty}^t e^{-\rho(t-k)} p_k dk,$$

where  $\rho$  is a discount rate. Differentiating this equation with respect to time yields

$$\dot{w} = \rho(w - p). \quad (5)$$

Given the definition of the real wage, its rate of change over time can be written as

$$\dot{w}/w = -\rho(1 - \frac{1}{w}) - \pi. \quad (6)$$

A linear approximation to equations (4) and (6) yields the following dynamic system in  $\omega$  and  $\pi$ :

$$\begin{bmatrix} \dot{\omega} \\ \dot{\pi} \end{bmatrix} = \begin{bmatrix} -\rho/\tilde{\omega} & -\tilde{\omega} \\ -\kappa(1 - c_y)y' & -\theta \end{bmatrix} \begin{bmatrix} \omega - \tilde{\omega} \\ \pi - \tilde{\pi} \end{bmatrix} + \begin{bmatrix} 0 \\ \kappa c_{z^*} z^* + \theta \varepsilon \end{bmatrix}, \quad (7)$$

where a “~” denotes steady-state values, and  $\tilde{\omega} = (1 + \varepsilon/\rho)^{-1}$  denotes the steady-state level of the real wage. The steady-state solution is characterized by  $\dot{\omega} = \dot{\pi} = 0$  and goods market equilibrium.<sup>2</sup> Thus, inflation and the rate of growth of nominal wages must be equal to the devaluation rate in the steady state ( $\dot{w}/w = \tilde{\pi} = \varepsilon$ ).

Figure A1 shows the long-run equilibrium of the model. In the North-East panel, curve  $\dot{\pi} = 0$  shows the combinations of the inflation rate and the real wage for which the inflation rate does not change over time, whereas curve  $\dot{\omega} = 0$  shows the combinations of  $\pi$  and  $\omega$  for which the real wage does not change. Curve  $y^s$  in the North-West quadrant shows the inverse relationship between output and the real wage (equation (1)). The consumption function (equation (2)) is represented in the South-West panel (for a given value of the long-run expected relative price  $z^*$ ). The long-run equilibrium values of inflation and the real wage are obtained at point  $E$ , with output determined at point  $A$  and consumption (which equals output) determined at point  $B$ .

Consider now the effect of a disinflation program that takes the form of a permanent, unanticipated reduction in the devaluation rate, from  $\varepsilon^h$  to  $\varepsilon^s < \varepsilon^h$ . The dynamics associated with this shock are illustrated in Figure A2. The curve  $\dot{\pi} = 0$  shifts to the left. Neither inflation nor the real wage changes on impact. Whereas the real wage rises monotonically throughout the adjustment process (which takes the economy from point

<sup>2</sup> Stability of the dynamic system (7) requires that the determinant of the matrix  $\mathbf{A}$  of coefficients be positive, and that its trace be negative:

$$\det \mathbf{A} = \theta\rho/\tilde{\omega} - \tilde{\omega}\kappa(1 - c_y)y' > 0, \quad \text{tr} \mathbf{A} = -(\theta + \rho/\tilde{\omega}) < 0.$$

These conditions are always satisfied here.



Figure A1  
Steady-State Equilibrium

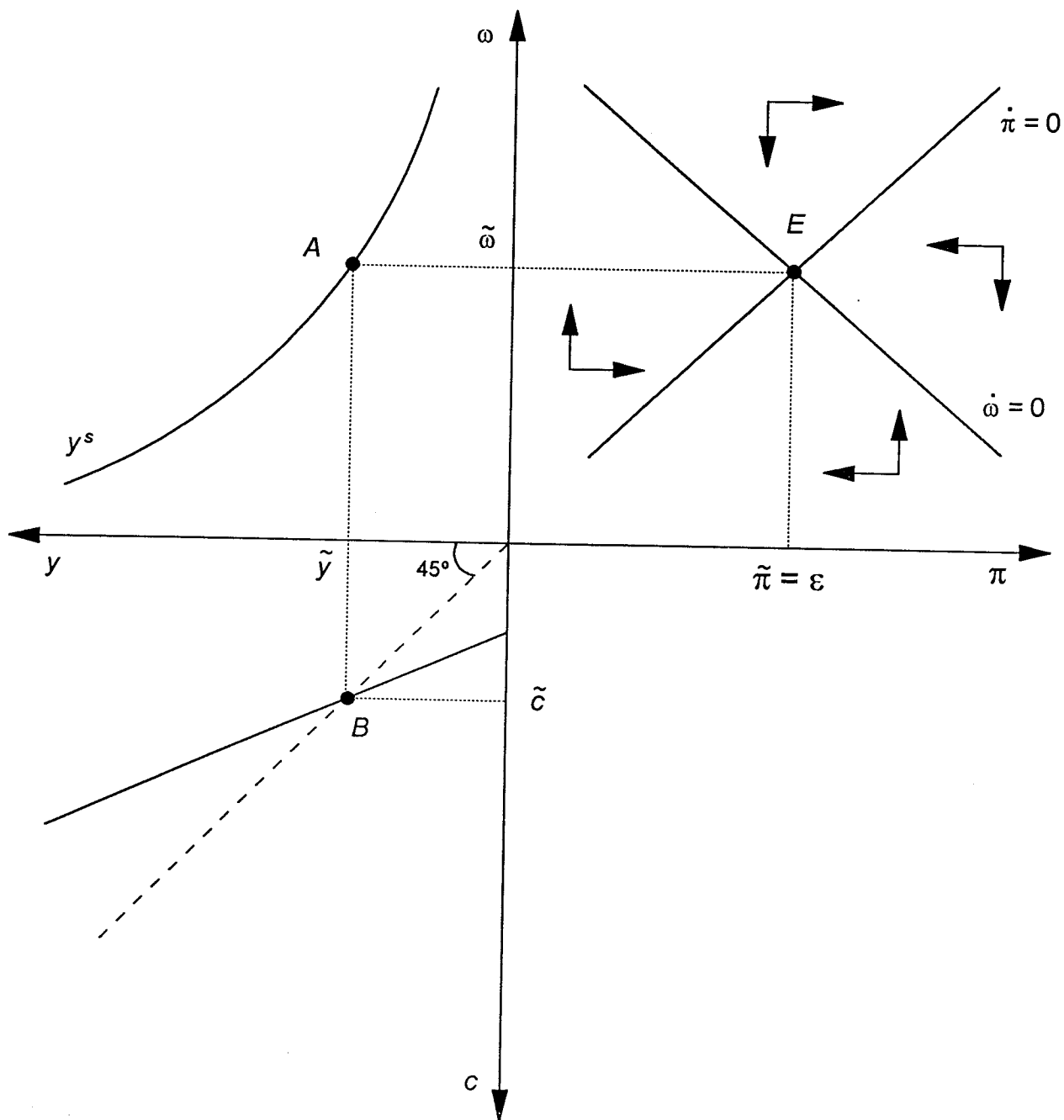
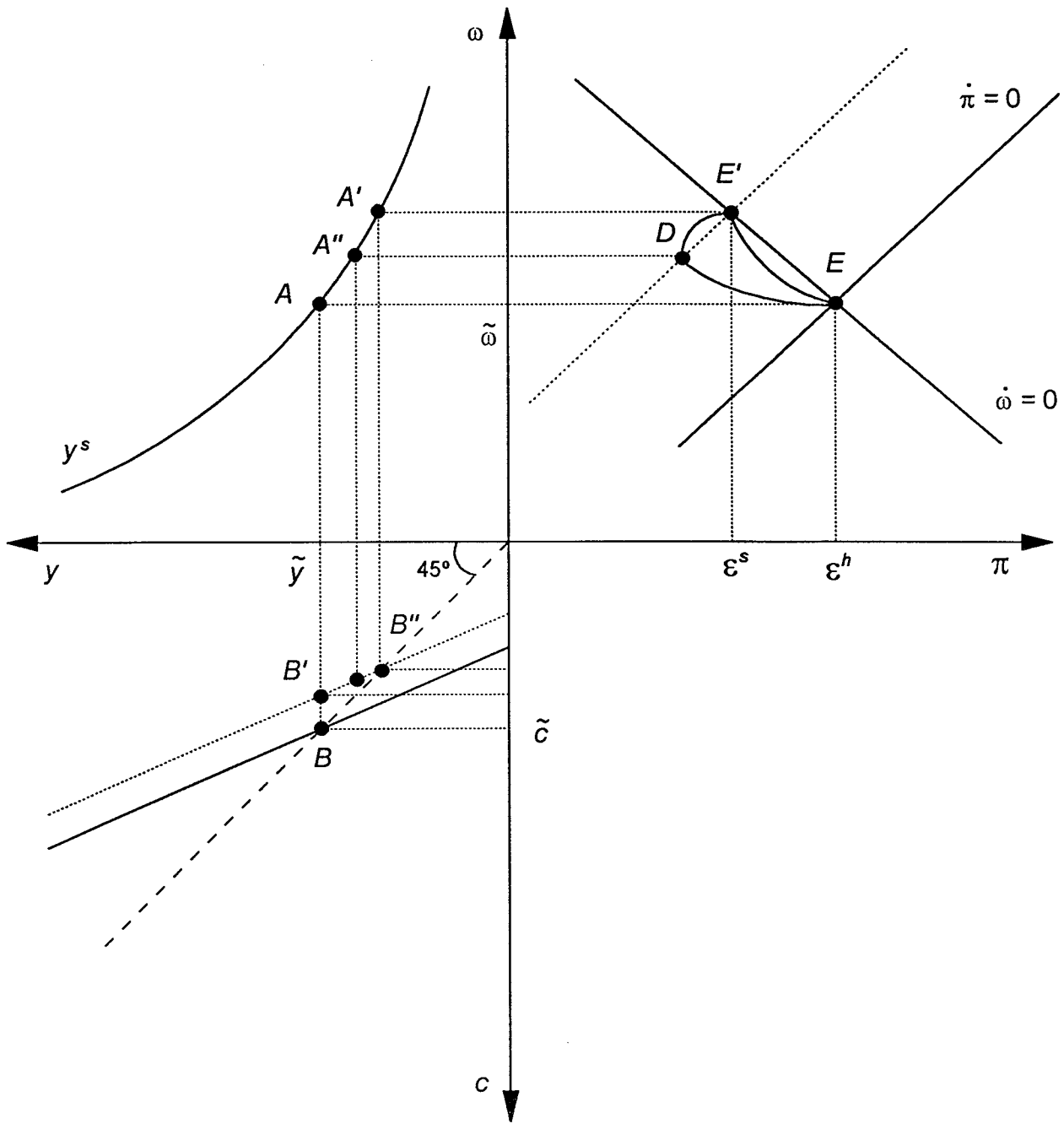


Figure A2  
Reduction in the Devaluation Rate



*E* to point *E'*), inflation may either fall continuously or may fall at first and increase in a second stage. Mimicking the path of real wages, output falls continuously from point *A* to *A'* in the North-West quadrant. The long-run increase in the real wage is reflected in a reduction in the expected long-run relative price  $z^*$ , thereby shifting downward the consumption function—in such a way that equilibrium of the market for domestic goods is maintained in the long run (point *B'*). The fall in consumption on impact (from point *B* to point *B'*), with output unchanged at its initial steady-state value, tends to create excess supply of domestic goods—thereby increasing the downward pressure on the inflation rate resulting from the reduction in the devaluation rate. Thus, the adjustment process to a cut in the devaluation rate leads to a gradual increase in real wages and falling output and consumption.<sup>3</sup>

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<sup>3</sup> By contrast, with forward-looking contracts a permanent reduction in the devaluation rate leads to an initial downward jump followed by a continuous fall in real wages, a consumption boom and an expansion in output (Agénor, 1996). The consumption and output boom that occurs under forward-looking contracts in this model accords well with the evidence on exchange rate-based stabilization programs; see Agénor and Montiel (1996, Chapter 10).

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