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THE MACROECONOMICS OF EXCHANGE-RATE AND PRICE-LEVEL INTERACTIONS: EMPIRICAL EVIDENCE FOR WEST GERMANY

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ABSTRACT

This paper studies the evidence on the conditional covariances between the German wholesale price level and the Deutsche mark exchange rate in the short run and in the long run. I rely both on an unrestricted time-series model, and on a structural Mussa-Dornbusch model. The results from unrestricted estimates indicate that the volatility of changes in the nominal exchange rate much exceed the volatility of the inflation rate both in the short run and in the long run. This implies a very high correlation between changes in the nominal and real exchange rate, and a correlation between the inflation rate and changes in the exchange rate that never exceeds .4--with 95% probability. The results from the structural estimates and sensitivity analysis indicate that perfect price flexibility is strongly rejected, and that the model tends to make sticky prices play a crucial role in explaining the evidence. Since the overidentifying restrictions implied by the structural model are rejected, I conclude that we still do not have a fully satisfactory explanation of observed extreme sluggishness of aggregate price levels.

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1. Introduction

The joint dynamics of prices and exchange rates are at the core of open-economy macroeconomic models. Exchange-rate fluctuations affect prices of imported intermediate and final goods, wage formation, and hence costs of production. Thus aggregate price dynamics are crucially influenced by exchange-rate effects. At the same time, the dynamic properties of the price level determine the extent to which goods prices react to shocks in the asset markets, ad the extent to which these shocks are reflected in exchange-rate fluctuations.

This paper uses data on West Germany to estimate and interpret the covariation of the exchange rate and the price level over different time-horizons. I follow the approach of Robert Flood [1981], who pointed out a number of empirical regularities, and asked which models best explain the empirical facts. At the same time, I extend Flood's analysis by studying the covariance between prices and exchange rates both in the short run and in the long run, and by directly estimating the parameters of macro models to identify more clearly the empirical limitations of existing theories. My empirical analysis concentrates on two main questions: the correlation between exchange rates and prices, and the variance of nominal exchange rates and prices.

A number of authors have recently addressed the issue of exchange-rate effects on aggregate price levels. The empirical evidence for the US ranges from estimates of reduced-form equations (as Rudiger Dornbusch and Stanley Fischer [1986], Robert Gordon [1985]), to estimates of equations describing firms' price-setting behavior (Wing Woo [1984]), and estimates of structural equations (Jeffrey Sachs [1985]). These results, while highlighting important correlations in the data, cannot be directly used for prediction purposes mainly because of the endogeneity of the exchange rate. Forecasts of the comovements of the price level and the exchange rate based on reduced-form equations with the exchange rate of the right hand side are likely to be biased. Similarly, the evidence from structural equations (estimated with instrumental variables procedures) alone is not very useful for forecasting purposes: structural estimates yield different predictions depending on the rest of the macroeconomy, and, once again, on the nature of exogenous shocks. Predictions on the comovements of the exchange rate and the price level over time can only be made after estimating all of the relevant feedbacks between these two variables, and the typical covariance matrix of exogenous disturbances: this is the main objective of this paper.

While unrestricted statistical models can provide unbiased estimates of covariance of prices and exchange rates over time, they cannot be used to falsify sticky- or flexible-prices theories of exchange-rate dynamics, because--under certain conditions--both families of models yield similar predictions. For this reason I estimate a structural model which subsumes both flexible- and sticky-prices as special cases, and use structural parameters' estimates to interpret the evidence from unrestricted regressions.

In section 2 I outline the basic channels that affect the comovements of prices and exchange rates with a macro model that emphasizes price setting by monopolistic competitors. The model also serves to identify the variables that form my dataset, whose trend properties I analyze in section 3. In section 4 I study the covariance matrix of prices and exchange rates in the short run and in the long run, obtained from an unrestricted vector autoregression (VAR) that includes the time series in section 3. To interpret the evidence, in section 5

I estimate the parameters of the model of section 2, and test the restrictions it imposes on the data. The model implies a number of constraints on the vector-autoregressive representation of the variables in the system: this constrained VAR is used to compute another set of dynamic covariance matrices of prices and exchange rates, which I compare to the unrestricted ones. Section 6 contains a summary of the results and some concluding remarks.

2. Price Level and Exchange Rate Linkages: A Benchmark Model

Sachs [1985] lists three major effects of exchange-rate changes on the price level:[1] the competitiveness effect, the direct effect, and the wage inflation effect. The first arises from shifts in the demand for domestic output associated with exchange-rate fluctuations. A depreciation of the exchange rate, other things equal, increases demand for domestic output, and leads to an increase in prices, to the extent that supply equations are upward sloping. The direct effect arises from the use of imported intermediate goods in the production of domestic output: an increase in the price of foreign currency increases domestic-currency costs of production, and is--in part--passed through into higher output prices. Finally, the wage inflation effect works though the determination of nominal wages and their effect on production costs. An exchange-rate depreciation increases the price of imported

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[1] In what follows the price level is taken to be the domestic-currency price of home output.

consumption goods, and brings about an increase in nominal wages:[2] these increases in costs of production are also passed though into higher output prices.

I capture the three effects by assuming that the economy is populated by many monopolistic competitive firms, all with identical technology and demand functions, and all producing differentiated goods.[3] The demand function for good i is, in logs:

$$\begin{aligned} q_{it} = & -\gamma[p_{it} - \lambda p_{dt} - (1-\lambda)(p^*_{ct}+e_t)] + \\ & d[m_t - \lambda p_{dt} - (1-\lambda)(p^*_{ct}+e_t)] + fq^*_t + hg_t + n_{1t} \end{aligned} \tag{1}$$

where q_i is output of the firm i, p_i the price charged by that firm, p_d the aggregate domestic price level, p^*_c the foreign consumption price level, e the price of foreign currency in terms of domestic currency, m the nominal money stock,[4] q^* a foreign demand shock, represented by foreign output, g a domestic demand shock, represented by domestic fiscal policy, and n_1 an exogenous innovation. In equation (1) foreign goods prices, and therefore the exchange rate, enter domestic residents' deflator with a weight of $(1-\lambda)$.

Firms' marginal costs are (in logs):

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[2] As workers attempt to maintain the purchasing power of wages.

[3] A version of this model, which did not include fiscal policy, was estimated by myself and Julio Rotemberg [1988]. Prices and exchange rates linkages in the context of alternative partial equilibrium models of industrial organization are studied by Dornbusch [1987].

[4] This specification is similar to the one used by John Taylor [1982].

$$\beta q_{it} + (1-\alpha_1-\alpha_2)p_{dt} + \alpha_1 w_t + \alpha_2(e_t+p^*_{Nt}) + n_{2t} \tag{2}$$

Factors of production are labor, with cost equal to w, imported intermediate goods, with foreign currency price equal to p^*_N, and domestic output. n_2 represents an exogenous (negative) productivity shock. The three effects discussed by Sachs can be clearly identified in the expression for the profit-maximizing price charged by firm i:[5]

$$\begin{aligned}\bar{p}_{it} = (1+\beta\gamma)^{-1}\{&[\beta\lambda(\gamma-d)+1-\alpha_1-\alpha_2-\lambda\alpha_1]p_{dt} + [\beta(1-\lambda)(\gamma-d)+(1-\lambda)\alpha_1+\alpha_2]e_t + \\ &[\beta(1-\lambda)(\gamma-d)+(1-\lambda)\alpha_1]p^*_{ct} + \alpha_2 p^*_{Nt} \\ &+ \alpha_1 k_t + \beta dm_t + \beta f q^*_t + \beta h g_t + n_t\}\end{aligned} \tag{3}$$

where $n_t = n_{2t} + \beta n_{1t}$, and k represents the real wage in terms of a consumption basket.[6] Equation (3) shows that, since all firms are identical, all firms

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[5] These effects are proportional to the terms in the coefficient of e in equation 3. $\beta(1-\lambda)\gamma$ represents the competitiveness effect: with upward sloping supply, $\beta>0$, and demand shifts by the elasticity times the weight of foreign goods in domestic consumption. α_2 represents the direct effect: the share of imported intermediate goods in domestic production. $(1-\lambda)\alpha_1$ is the wage inflation effect: given the real wage, an increase in the exchange rate brings about an increase in costs proportional to the share of imported goods in the consumption deflator, times the share of labor in total costs. Exchange-rate changes have an additional effect on goods demand, through the deflator for real balances, represented by the term $-d\beta(1-\lambda)$ in the coefficient of e_t.

[6] In the econometric estimation that follows, I assume that the real consumption wage is only affected by innovations in the other forcing variables in the model, but is not affected by demand, productivity, or velocity shocks. However, the real <u>product</u> wage is endogenous.

would charge the same profit-maximizing prices.

To allow for the presence of price stickiness, I specify a price dynamics equation derived from Rotemberg [1982]:

$$p_{it} = cp_{it-1} + \rho c_t p_{it+1} + (1-c-\rho c)\bar{p}_{it} \tag{4}$$

where c represents a cost of price adjustment--a parameter that can be estimated, and ρ is a constant discount factor. c is a transformation of the quadratic cost of price adjustment of Rotemberg [1982]: if C stands for Rotemberg's cost of price adjustment, $c=C/[1+(1+\rho)C]$.[7] Equation (4) implies that the current price charged by firm i is a weighted average of the discounted value of the price it expects to charge at time t+1, ${}_tp_{t+1}$, of last period's price, and of $\bar{p}_{it}$. With c=0 the model implies perfect price flexibility. The aggregate version of (4) is equivalent to the wage dynamics equation of Taylor [1980], as parametrized by Guillermo Calvo [1983] to model staggered price setting.[8] Notice that the larger the cost of changing prices c, the smaller is the effect of current information on the current price level.

Finally, the feedback from the price level to the exchange rate is modelled after Robert Mundell [1968], J. Marcus Fleming [1962], Michael Mussa [1976] and Dornbusch [1976]: I assume that the exchange rate is forward-looking, and is

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7 Thus c ranges between 0 and $1/(1+\rho)$. Equation (4) is a first order condition from firms' value maximization problem, where profits depend both on the level and the rate of change of p.

8 Rotemberg [1987] shows that equivalence. See the papers by David Backus [1984] and myself [1988a] for discrete-time empirical applications of the Taylor-Calvo model.

determined by current and expected future excess demand for money and foreign interest rates. Equilibrium in the money market is specified as follows:

$$m_t - \lambda p_{dt} - (1-\lambda)(e_t+p^*_{ct}) = aq_t - b[i^*_t + {}_te_{t+1} - e_t] + n_{3t} \tag{5}$$

where q stands for the index of domestic output, i^* is the foreign interest rate, ${}_te_{t+1}$ is the the expectation of e at time t+1, conditional on information at time t, and n_3 represents a money demand disturbance. In equation (4), by assuming that the opportunity cost of holding money is represented by the uncovered return on foreign assets, I relegate any time-varying risk premium in the foreign exchange market to the disturbance term n_3. This is not entirely inappropriate, since the extensive empirical tests of various versions of international capital asset pricing models have so far been unsuccessful in providing a reliable model for the dynamics of the risk premium.[9] Thus n_3 includes both a velocity shock and a random risk premium. Given that the real wage and the price of imported materials are assumed to be exogenous, firms are not constrained in the inputs markets. Therefore aggregate output is obtained by summing over the firms i the demand equations in (1):

$$q_t = \gamma(1-\lambda)[e_t+p^*_{ct}-p_{dt}] + d[m_t-\lambda p_{dt}-(1-\lambda)(e_t+p^*_{ct})] + fq^*_t + hg_t + n_{1t} \tag{6}$$

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[9] See, for example, the tests by Jeffrey Frankel [1982], and Robert Hodrick and Sanjay Srivastava [1984]. On the other hand, the apparent negative correlation between the risk premium and the expected change in the exchange rate documented by Eugene Fama [1984] and Hodrick and Srivastava may not be consistent with this assumption.

Changes in p_d affect money market equilibrium both through the deflator for real balances and through aggregate demand: thus the contemporaneous feedback from the price level to the exchange rate is ambiguous.[10]

For the purpose of studying the covariance of prices and exchange rates over time, knowledge of the parameters in the model above is clearly not enough. We need to compute the reduced form of the system in (1)-(6), and estimate the typical pattern of disturbances affecting the economy, since, as is well known, the covariance of endogenous variables in a model depends on the source of exogenous shocks. A few preliminary observations, however, can be usefully made without explicitly writing down the full solution of the model. To simplify the exposition, and without loss of generality, let for the moment p_c^* be constant and equal to zero. Then the following identity relates the variance of the real exchange rate--defined as $e-p_d$--the variance of the price level, the variance of the nominal exchange rate, and the correlation between exchange rates and prices:

$$\mathrm{Var}(p_d) = \mathrm{Var}(e-p_d) - \mathrm{Var}(e) + 2\sigma_e\sigma_{pd}\mathrm{Corr}(p_d,e) \qquad (7)$$

The identity in (7) says that the variance of the nominal price level is small, when the variance of the real and nominal exchange rate are of similar

[10] Notice that with $\gamma>1$ an increase in p_d--keeping other things equal--has actually ambiguous effects on excess demand for money, since presumably the share of domestic goods in the consumption deflator is larger than that of foreign goods.

magnitudes, and the covariance between the nominal exchange rate and the price level is small. Further manipulation of (7) shows that in this case the correlation between the nominal and the real exchange rate is large. Conversely, the correlation between the nominal exchange rate and the price level is large, when the variance of the real exchange rate is small relative to the variance of the nominal exchange rate and of the price level.

It is illuminating to discuss the covariance of prices and exchange rates in the two cases where prices are perfectly flexible--$c=0$--and where prices are sticky--$c>0$. When $c=0$, the real exchange rate can be solved for from the equilibrium in the goods market, and is a function of real variables affecting demand and cost functions. The nominal exchange rate is in turn determined in the asset markets, after solving forward the difference equation implicit in (5). If, again for simplicity, the forcing variables follow the same first-order stochastic process the exchange-rate equation would be:

$$e_t = (1+b)^{-1}K[m_t - n_{3t} + bi^* + \lambda(e_t - p_{dt}) - aq_t] \tag{8}$$

where the constant K is a function of b and the autoregressive coefficient of forcing variables. q is itself a function of the real exchange rate and variables affecting goods supply and demand.[11] Equation (8) says that, if goods

[11] Since these real variables are determined independently of the nominal exchange rate, I can use the semi-reduced form as in equation (8) for the purpose of this exposition. A real exchange rate depreciation has conflicting effects on equilibrium output, since it increases demand, but at the same time it increases marginal costs through the "wage inflation effect," thus tending to reduce supply.

market shocks that depreciate the real exchange rate either do not affect output significantly or decrease output, and if the variance of exogenous shocks in the money market is small relative to the variance of the real exchange rate, then the nominal and the real exchange rate are positively correlated, and, from equation (7) the variance of the price level is smaller than the variance of the exchange rate. This result was pointed out by Flood [1981], and Maurice Obstfeld and Alan Stockman [1985].[12] The sticky-prices model yields similar predictions, but for entirely different reasons. Sticky nominal prices imply a small variance of the price level in the short-run, and, as a result, a high correlation of the nominal and the real exchange rate.

Therefore flexible and sticky-prices models can yield similar predictions on the comovements of prices and exchange rates over time, and cannot be falsified simply by the analysis of unrestricted estimates. For this reason in the next section I will both offer unrestricted evidence, and attempt to interpret this evidence using structural estimates.

3. The Data and their Trend Properties

I use monthly data on West Germany, during the generalized-floating period, from June 1973 to July 1987. Sources are IMF International Financial Statistics and OECD Main Economic Indicators. World final goods prices, as well as world

[12] In the models used by Flood and Obstfeld and Stockman money demand depends on real income, and output is exogenous: then the correlation between the nominal and real exchange rate is always nonnegative.

activity are, respectively, the OECD consumption price index, and the OECD industrial production index. The Data Appendix reports all details necessary to reconstruct the data set in this paper.

One major issue to be faced in the estimation regards the trend properties of the data. As is well known,[13] most macroeconomic time series have very high autocorrelation coefficients, such that they cannot be statistically distinguished from univariate random walks. Table 1 presents tests of the hypothesis that each individual series can be represented as a stationary process in the first differences. The two columns on the left report statistics for the hypothesis $H_0:\phi_1=0\ \delta_1=1$, in the following univariate model:

$$X_t = \phi_0 + \phi_1 t + \delta_1 X_{t-1} + \delta_2(X_{t-1}-X_{t-2}) + \delta_3(X_{t-2}-X_{t-3}) + \delta_4(X_{t-3}-X_{t-4}) + u_t$$

where X is the log of each series in the dataset, and t is a linear time trend. The statistic for the null hypothesis is algebraically identical to the F statistic for that linear constraint in an ordinary-least-squares regression. Its distribution is reported by David Dickey and Wayne Fuller [1981]. For all variables except q^* the null hypothesis is not rejected at the 95 percent confidence level. In the case of q^* the null is not rejected at the 97.5 percent level. Table 1 also reports James Stock and Mark Watson's [1986] q^f test of a stochastic trend in each series:[14] in all cases the results fail to

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[13] See, for example, Charles Nelson and Charles Plosser [1982].

[14] Autoregressions of first differences of all variables that include a linear trend indicate that preprocessing in the form of linear detrending is not needed.

reject the null hypothesis. Thus the univariate tests in table 1 suggest that all series need to be differenced to achieve stationarity.

First-differencing all variables is not appropriate, however, when some of them are cointegrated. This happens when, while individual time series are non-stationary, certain linear combinations of them are stationary. In the model illustrated in section 2 and 3 cointegration arises if, for example, demand and productivity disturbances were stationary stochastic processes. Then the reduced form expressions for p_d and e would contain stationary disturbances, indicating that the exchange rate, the price level, and the other variables in the model have (in the terminology of Stock and Watson [1986]) common stochastic trends. Similarly, the forcing variables in the model might have common stochastic trends. Robert Engle and Clive Granger [1987] have shown that, in the presence of common stochastic trends,[15] the vector-autoregressive representations that this paper relies on are inappropriate. I perform tests to detect the presence of common stochastic trends, in several different groupings of the variables in the dataset. These tests, due to Stock and Watson [1986], are based on the eigenvalues of certain coefficient matrices in the common-trend representation of the series in each grouping. When the number of common trends is less than the number of variables, these coefficient matrices have rank that is smaller than the number of variables.

Table 2 reports the results. The column on the right shows that the null hypothesis that the number of common trends equals the number of variables is

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[15] Stock and Watson [1986] show that if some series have stochastic trends in common, their multivariate representation is cointegrated, as defined by Engle and Granger [1987].

not rejected at the 95 percent confidence level for all groups in the table, except in the case of k and q^*. In that case, the null hypothesis would not be rejected at 97.5 percent or higher levels. Notice that the test for the absence of a common stochastic trend among final-goods prices and the nominal exchange rate is not rejected, i.e. the real exchange rate follows a random walk according to the data. Although John Huizinga [1987] finds some evidence of stationarity in his own real exchange rate data, he reports a very slow rate of mean reversion.[16]

In conclusion, the evidence presented in this section suggests that first-differencing all the data appears to be an acceptable procedure to achieve stationarity, and that consistent estimates parameters of both the unrestricted and the restricted model could be recovered using differenced data.[17]

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16 Evidence that the real exchange rate is approximated by a random walk does not, <u>per se</u>, invalidate the hypothesis that prices are sticky. If the determinants of the long-run equilibrium exchange rate follow random walk processes, so would the real exchange rate. See Stockman [1987] for a discussion of the implications of alternative theories on the behavior of the real exchange rate, and Kenneth West [1987] for an example where near-random walk behavior of real variables is actually induced by the presence of sticky prices.

17 Nelson and Heejoon Kang [1981] discuss the problems arising when random walks are detrended with linear non-stochastic trends. In a small sample, tests of difference-stationarity have little power against a trend-stationary model where the indeterministic component has an autoregressive coefficient arbitrarily close to 1. This problem is raised by Nelson and Plosser [1982] and is discussed by Bennet McCallum [1986]. See also Paul Samuelson [1976] for a critical discussion of the random walk assumption. The main results of this paper turn out to be unaffected by the method used to achieve stationarity.

4. Predicting the Short- and Long-Run Covariances of the Price Level and the Exchange Rate: Unrestricted Estimates

In this section I present empirical evidence on the joint dynamics of inflation rates and changes in the exchange rate base on an unrestricted statistical model which includes all variables appearing in sections (2) and (3). The model is compactly written as follows:

$$x_t = Gx_{t-1} + u_t \tag{9}$$

where $x_t' = [p_{dt}, p_{dt-1}, \ldots, e_t, e_{t-1}, \ldots, p^*_{ct}, p^*_{ct-1}, \ldots, p^*_{Nt}, p^*_{Nt-1}, \ldots, k_t, k_{t-1}, \ldots, q^*_t, q^*_{t-1}, \ldots, m_t, m_{t-1}, \ldots, g_t, g_{t-1}, \ldots, i^*_t, i^*_{t-1}, \ldots]$ (all variables are now log-differences), and u_t is a vector of i.i.d. normal disturbances. The vector x contains, for each variable, a number of lags equal to the order of the autoregression minus 1. The order of the VAR is chosen to minimize residual variance, and any systematic autocorrelation of residuals.

Using (8) recursively, I write the deviation of variables in x from their conditional expectation, ℓ periods ahead:

$$x_{t+\ell} - E(x_{t+\ell} \mid \theta_t) = x_{t+\ell} - G^{\ell} x_t = \sum_{j=0}^{\ell-1} G^j u_{t+\ell-j} \tag{10}$$

Where θ_t is information available at time t. Then I use (9) to derive the covariance matrix of ℓ-period-ahead innovations:

$$E\ [x_{t+\ell} - G^{\ell} x_t][x_{t+\ell} - G^{\ell} x_t]' = \sum_{j=0}^{\ell-1} G^j \Sigma (G\ )'^j \tag{11}$$

where Σ is the contemporaneous covariance matrix of the u's. If the system is stable, the eigenvalues of the matrix G are less than 1 in modulus, and the covariance matrices converge to a given steady-state value. Equation (11) provides the conditional covariance matrix of all variables in the model over different time horizons, given the estimated reduced form equation, and the estimated typical pattern of exogenous shocks. It thus takes into account the simultaneous determination of endogenous variables, as well as their dependence on the different underlying shocks. The relevant dynamics are induced by the powers of the matrix G.

Table 3 contains summary statistics for the system (9). I have chosen to estimate a fifth-order VAR because of the relatively large number of variables entering the system. The results I report below would be very little affected by the number of lagged variables included in each equation. The estimates of the matrix G and the covariance matrix of innovations Σ are then applied recursively as in equation (11) to compute the conditional covariance matrix between the exchange rate and the price level over time. I concentrate on the standard error of the inflation rate, the standard error of the change in the nominal exchange rate, the correlation between the inflation rate and changes in the nominal exchange rate, and the correlation between changes in the nominal and the real exchange rate.

For all of these statistics I compute 90 percent confidence bounds following Bradley Efron's [1982] "bootstrap" method:[18] Given the estimates of G,

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[18] This appears more appropriate than the normal approximation since these statistics are highly nonlinear functions of the original parameters, as suggested by equation (11). See David Runkle [1986] for a discussion and an application to a similar context.

I get an estimate of T disturbance vectors u_t in equation (9), where T is the size of the original sample. I attach probability 1/T to each estimated u_t, and generate 300 artificial samples of x, by drawing from the estimated u's and applying recursively equation (9), conditional on the initial estimates of G, and the sample realization of x_0. For each artificial sample, I reestimate all parameters, together with the statistics of interest here. These empirical distributions are used to compute confidence bounds.

Figure 1 plots the 90 percent confidence bounds for the standard errors of the inflation rate and of the change in the nominal exchange rate from 1 to 24 months ahead.[19] Standard errors are expressed in percent per annum. The figure highlights the most important feature of the data, namely, as noted by Flood [1981] and others, the volatility of the (changes in the) nominal exchange rate much exceeds the volatility of the inflation rate. The standard error of the inflation rate ranges from 1.19 to 1.499 on a 1-month horizon, and from 2.28 to 3.10 on a 24-month horizon. By contrast, the standard error of the change in the nominal exchange rate ranges from 10.63 to 13.34 on a 1-month horizon, and from 14.82 to 18.15 on a 24-month horizon. Thus the difference in volatility between the nominal price level and the nominal exchange rate is not just a short-run phenomenon, but is also present in the long-run, unconditional estimates.[20] Figure 2, which plots the confidence bounds for the correlation

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[19] These statistics converge relatively fast to the long-run, unconditional values.

[20] The "empirical regularities" literature of Mussa [1979] Flood [1981] and others concentrates on the covariance of one-step-ahead forecast errors exclusively.

between changes in the real and the nominal exchange rate, confirms the finding of figure 1: the correlation between the real and the nominal exchange rate exceeds 99 percent (with 95 percent probability), both in the short run and in the long run, although it appears to be higher at higher frequencies.

Figure 3 plots the correlation between the inflation rate and changes in the nominal exchange rate: the correlations range between 10 and 40 percent in the long run. This result suggests the importance of fluctuations of the real exchange rate, which prevent prices and exchange rates from moving more closely together even in the long run.

4. Evidence from the Structural Model

In this section I estimate the parameters of the model in section 2, and use them to interpret the evidence based on the unrestricted time-series model of the previous section.

I estimate a system that includes the aggregate demand equation (6), the money demand equation (5), and the equation describing equilibrium price dynamics (4)--after aggregation across firms i, and solving for $\bar{p}_{dt}$, which equals a constant times marginal costs. The three equations are estimated jointly, by replacing the conditional expectations of the future price level and exchange rate with their actual realizations. This procedure, as Robert Cumby, Huizinga and Obstfeld [1983] argue, produces composite disturbances in the price dynamics and output demand equations, that are the sum of the structural disturbances and future surprises: thus the disturbances in the price dynamics and money demand equations follow a first-order moving-average process. By applying the generalized instrumental variables method of Lars Hansen [1982], Hansen and Kenneth Singleton [1982] and Cumby-Huizinga-Obstfeld [1983], I take into account both the first order moving-average process of the disturbances, and the possible presence of conditional heteroskedasticity in the disturbances.

Table 4 reports the results. The estimation was carried out assuming given values of λ (the share of domestic goods in the consumption deflator), ρ (the rate of discount of firms' profits), and α_1 and α_2 (the shares of labor and imported intermediates in the aggregate production function), and using lagged values of the forcing variables as instruments. All parameters, except β an d, are significantly different from zero and of the expected sign. Since β is insignificantly different from zero, the hypothesis of constant returns to scale cannot be rejected. The point estimates of the elasticity of demand for domestic output γ and of the output elasticity of money demand a are quite high--4.3 and 1.8 respectively--but are not significantly different from lower, and more reasonable, values. Finally, the most notable result of table 4 is the size and significance of c, the parameter representing the cost of changing prices: its value of 0.48 is more than 30 standard errors away from 0, leading to a strong rejection to the flexible-prices version of the model.

The table also reports autocorrelations of the estimated disturbances at different lags. Although we still do not know the distribution of these statistics, it is useful to compare them to the what the thoretical model predicts. While--as Cumby, Huizinga and Obstfeld [1983] show--the model predicts high values for the first order autocorrelation coefficients in the price-dynamics and money demand equations, it does not impose any *a priori* restriction on the value of the autcorrelation coefficients in the output demand equation. The table shows, contrary to the predictions of the model, that the autocorrelation coefficients in the price dynamics equation remain quite high after the first lag. In the money demand equation it is also hard to detect a drop of the autocorrelation coefficient after the first lag. Finally, the estimated disturbances in the output demand equation do not display any particular pattern or high values.

As Hansen [1982] pointed out, the model and the rational expectations assumption jointly imply overidentifying restrictions that can be tested. The restrictions are that the inner products of the instruments and the disturbances in excess of the estimated parameters should be close to zero according to a certain metric. Hansen's J statistic, reported in Table 4, is distributed as χ^2 with degrees of freedom equal to the number of instruments times the number of equations, less the number of parameters to be estimated. The value of the statistic implies a rejection of the null hypothesis at high confidence levels. Thus while there is apparently strong evidence against the flexible-prices version of the model, suggesting that the differences in volatility of prices and exchange rates in the short run might not be due exclusively to real shocks, the sticky-model is also not fully consistent with the evidence.

The rejection of overidentifying restrictions, however, does not indicate where the sticky-prices model fails in explaining the comovements of exchange rates and prices. For this reason, I compute the restrictions imposed by the sticky-prices model on the reduced-form (9), and derive the resulting dynamic covariance matrices between p_d and e, which I compare with the unrestricted ones. To obtain these restrictions, I solve to eliminate output from the system of three equations whose estimates are reported in table 4, and obtain a system of difference equations in p_d and e. These equations are solved using the now-standard methods outlined by Hansen and Thomas Sargent [1980], which involve the assumption that agents know the stochastic processes of the forcing variables, and use this knowledge to form expectations about their future paths, based on past realizations of these variables. The result is a system of linear equations in p_d, e, and the seven forcing variables, with nonlinear constraints across the coefficients of the vector-autoregressions of the forcing variables (which I estimate again including 5 lags for each variable)[21] and the coefficients of the reduced froms of p_d and e:

$$x_t = \Gamma x_t + \Pi x_{t-1} + \Psi u_t \tag{12}$$

Where x was defined above, The matrix Γ has nonzero elements, since contemporaneous values of the forcing variables enter the reduced-form equations for the endogenous variables.

Equation (12) implies the following constrained VAR:

$$x_t = G x_{t-1} + D u_t \tag{13}$$

where $$G = (I-\Gamma)^{-1}\Pi$$

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21 According to the model, lagged values of prices and the exchange rate are not to be included in the reduced-form representation of the forcing variables.

$$D = (I-\Gamma)^{-1}\Psi$$

The covariances between p_d and e can now be computed as in equation (11). The results of these calculations are reported in figures 4 and 5. Figure 4 plots the correlation between changes in the nominal and the real exchange rate, and, for reference, the 90% confidence bounds obtained from the unrestricted model. While the correlation between changes in the nominal and real exchange rate implied by the restricted model also exceed .9 over all time horizons, they are clearly outside the confidence region implied by the unrestricted model. As theory predicts, the restricted correlation between nominal and real exchange rates changes is highest in the short run, and declines with the length of the horizon. The difference between the restricted and unrestricted estimates is also apparent in the estimates of the relative magnitudes of the standard errors of inflation rates and exchange rate changes.[22] While in the case of the unrestricted estimates the ratio of the standard error of the exchange-rate change and the rate of inflation ranges from 7.1-11.2 (1-month ahead) to 4.8-8.0 (24-months ahead), the same ratio implied by the constrained model ranges from 3.7 (1-month ahead) to 2.4 (24-months ahead). Thus the model I estimate implies a larger relative volatility of the rate of inflation, than what is observed in the data. Finally, figure 5 plots the correlation between the inflation rate and the changes in the nominal exchange rate. The figure shows that the unrestricted estimates fall within the 90% confidence bounds implied by the

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22 Since the VAR is estimated by minimizing the residual variance, the levels of the standard errors of prices and exchange rates are much smaller when computed from the unrestricted estimates.

model. It also shows that the correlation between exchange-rate changes and the inflation rate tends to increase with maturity: in other words, the restricted model tends to ascribe some of the short-run variability of the real exchange rate to short-run nominal rigidities.

The role of short-run nominal rigidities can also be assessed by performing sensitivity analysis. Decreasing c from .48 to .1 increases the correlation between the inflation rate and changes in the nominal exchange rate to .76 in the 1-month horizon and .85 in the 24-month horizon. Furthermore, the variance of the nominal price level actually exceeds the variance of the nominal exchange rate in this case. Thus the model tends to ascribe a very important role to price stickines in explaining the observed comovements of exchange rates and prices.

6. Summary of the Evidence and Concluding Remarks

This paper has studied the evidence on the conditional covariances of (changes in) the German wholesale price index and the Deutsche mark exchange rate. The results from unrestricted estimates indicate that the volatility of changes in the nominal exchange rate much exceed the volatility of the inflation rate both in the short run and in the long run. This implies a very high correlation between nominal and real exchange rate changes. A high volatility of the real exchange rate is also associated with a relatively low long-run correlation between the inflation rate and exchange-rate changes, which--with

95% probability--never exceeds .4.

Both sticky- and flexible-prices models could be consistent with these results. In order to offer an intepretation of the evidence, I estimate a Mussa-Dornbusch structural model, that allows for a wide variety of real shocks--including shocks in productivity, demand, real wages, fiscal policy, and real imported-materials prices--and allows to test perfect price flexibility as a nested hypothesis. Most of the estimated parameters are significant and of reasonable magnitudes. The hypothesis of perfect price flexibility is rejected with a t statistic of 34. Furthermore, sensitivity analysis shows that the estimated model tends to ascribe a crucial role to price stickiness in the pattern of covariances of the exchange rate and the price level. At the same time, however, the overidentifying restrictions are soundly rejected.

Rejections of the overidentifying restrictions do not indicate the source of misspecification: by computing the dynamic covariance matrices between the inflation rate and changes in the exchange rate implied by the model, I show that the model predicts a higher volatility of the inflation rate, relative to the volatility of the nominal exchange rate, than what is actually observed in the data.

These results should be strong enough to persuade empirical researchers to look for an explanations of the observed sluggishness of the aggregate price level. On one hand, the real side of the model could be enriched, in a way to make shocks have a larger impact on the real exchange rate. This might give less of a crucial role to price stickiness as an explanation of the evidence.

On the other hand, recent theoretical models of costly price adjustment[23] might lead to empirical applications that fit the data more satisfactorily.

[23] Like those of Michael Parkin [1986], N. Gregory Mankiw [1985], Lawrence Ball and David Romer [1987].

Appendix: The Data

p_d	International Financial Statistics line 63: wholesale price index.
e	International Financial Statistics line rf: average dollar exchange rate.
p_c^*	OECD Main Economic Indicators line I134750009H country # 950 (total OECD): consumer price index.
p_N^*	OECD Main Economic Indicators line I147100T9H Basic materials price index - industrial goods - imported. (West Germany)
q	International Financial Statistics line 66c: Industrial Production - Adjusted.
q^*	OECD Main Economic Indicators line I12100009K country # 950 (total OECD): industrial production in manufacturing.
k	Deutsche Bundesbank Bullettin, Statistical Appendix, Series 4, Table 10: Wages and Salaries per Man-Hour, divided by International Financial Statistics line 64 (consumer price index).
i^*	International Financial Statistics line 60c country 111: (US Treasury Bills Rate).
m	International Financial Statistics line 34bc (M1, adjustec
g	Computed as the log of a weighted average of government spending and government debt, with weights equal to $[r-\lambda\theta+(1-\lambda)p]/(r+p)$, and $p/(r+p)$, with r (real interest rate) = .02, θ (utility discount factor) = .04, p (reciprocal of average age in the economy) = 1/45. See Giovannini [1988b]. International Financial Statistics line 82 (Government Spending) and line 88 (Government Debt).

Table 1:

Univariate Unit-Root Tests
(monthly data, 1973:6 to 1987:6)

Series	Dickey-Fuller Test:		Stock-Watson Test:	
	D.W.	Φ_3	q^f	Marg. Signf.
p_t	1.85	3.66	0.073	70.00
e	1.99	1.07	-0.687	53.50
p_c^*	2.01	5.85	-1.584	38.25
p_N^*	2.01	3.47	-0.845	50.25
i^*	1.99	1.52	-0.898	49.25
q^*	2.01	6.45	-2.481	28.00
m	2.08	3.57	0.183	72.75
k	1.99	3.96	-3.655	18.75
q	1.99	3.18	0.024	69.00
g	1.97	1.17	-3.992	16.75

Notes: The two columns on the left report statistics from the following regression:

$$X_t = \phi_0 + \phi_1 t + \delta_1 X_{t-1} + \delta_2(X_{t-1}-X_{t-2}) + \delta_3(X_{t-2}-X_{t-3}) + \delta_4(X_{t-3}-X_{t-4}) + u_t$$

Φ_3 tests the restriction $\phi_1=0$, $\delta_1=1$. Its distribution is reported by Dickey and Fuller [1981]. The two columns on the right report the Stock-Watson q^f test, described by Stock and Watson [1986].

Table 2:

Testing for Common Trends
(monthly data, 1973:6 to 1987:6)

Variables	Hypothesis Tested (# of common trends)	q^f	Marg. Signf.
$p_c^*\ p_N^*$	2 vs. 1	-1.852	85.25
$m\ i^*$	2 vs. 1	-6.866	41.75
$m\ g$	2 vs. 1	-2.644	78.25
$m\ i^*\ g$	3 vs. 2	-4.924	90.25
$k\ q^*$	2 vs. 1	-18.310	4.00
$k\ q\ q^*$	3 vs. 2	-18.620	18.25
$k\ q\ q^*\ g$	4 vs. 3	-20.799	34.75
$m\ i^*\ p_c^*\ p_N^*$	4 vs. 3	-26.108	17.50
$i^*\ p_N^*\ g$	3 vs. 2	-3.525	96.00
$q\ q^*\ i^*\ p_c^*\ p_N^*$	5 vs. 4	-30.067	23.75
$m\ i^*\ p_c^*\ p_N^*\ k\ q^*$	6 vs. 5	-40.737	16.75
$e\ p_d$	2 vs. 1	-5.195	55.50
$e\ p_d\ p_c^*$	3 vs. 2	-5.339	88.00
$e\ p_d\ m\ i^*$	4 vs. 3	-10.174	86.75
$e\ p_d\ p_c^*\ p_N^*$	4 vs. 3	-1.993	99.75
$e\ p_d$ pc g	4 vs. 3	-5.475	99.00
$e\ p_d\ k\ q^*$	4 vs. 3	-18.377	45.50
$e\ p_d\ m\ i^*\ p_c^*\ p_N^*$	6 vs. 5	-42.427	13.50

Table 3:

Unrestricted Vector Autoregression
Summary Statistics

Equation	$\bar{R}^2$	D.W.	S.E.E (percent)
p_d	.620	1.92	.133
e	.046	1.99	1.166
p_c^*	.595	2.02	.092
p_N^*	.210	1.94	.935
i^*	.185	2.01	.078
k	.240	2.00	.663
q^*	.329	2.17	.300
m	.151	1.95	.404
g	.242	1.99	3.991

Sample: December 1973 to June 1986.
Degrees of freedom: 118.

Table 4:

Estimates of the Structural Parameters

$$p_{dt} = cp_{dt-1} + \rho c_t p_{dt+1} + (1-c-\rho c)(\ \beta q_t + (1-\alpha_1-\alpha_2)p_{dt} + \alpha_1[k_t+\lambda p_{dt}+(1-\lambda)(e_t+p^*_{ct})] + \alpha_2(e_t+p^*_{Nt}) + n_{2t}\ ) \quad \text{(i)}$$

$$q_t = \gamma(1-\lambda)[e_t+p^*_{ct}-p_{dt}] + d[m_t-\lambda p_{dt}-(1-\lambda)(e_t+p^*_{ct})] + fq^*_t + hg_t + n_{1t} \quad \text{(ii)}$$

$$m_t - \lambda p_{dt} -(1-\lambda)(e_t+p^*_{ct}) = aq_t - b[i^*_t + {}_te_{t+1}-e_t] + n_{3t} \quad \text{(iii)}$$

Parameter	Value	Standard Error
a	1.780	(0.860)
b	0.355	(0.172)
c	0.479	(0.014)
d	0.357	(0.322)
f	0.762	(0.346)
h	0.287	(0.083)
β	-0.540	(0.356)
γ	4.323	(1.722)

Sample autocorrelations of the residuals at the first 5 lags:

(i)	-0.54	-0.19	0.47	-0.25	-0.19
(ii)	-0.20	-0.17	0.13	-0.04	0.10
(iii)	-0.14	-0.12	0.13	0.04	-0.07

J Statistic [$\chi^2(13)$]: 43.0853

Sample: December 1973 to June 1986. The estimates are computed assuming ρ = 0.99594 (5% p.a. discount rate), λ=0.8 (average share of imports in spending in the 1973-87 period equals 0.2), and α_1=0.6, α_2=0.14 (from Michael Bruno and Jeffrey Sachs [1985]).

Fig. 1: Inflation–Rate and Exchange Rate Volatility

90% Confidence Bounds (Using Percent Changes)

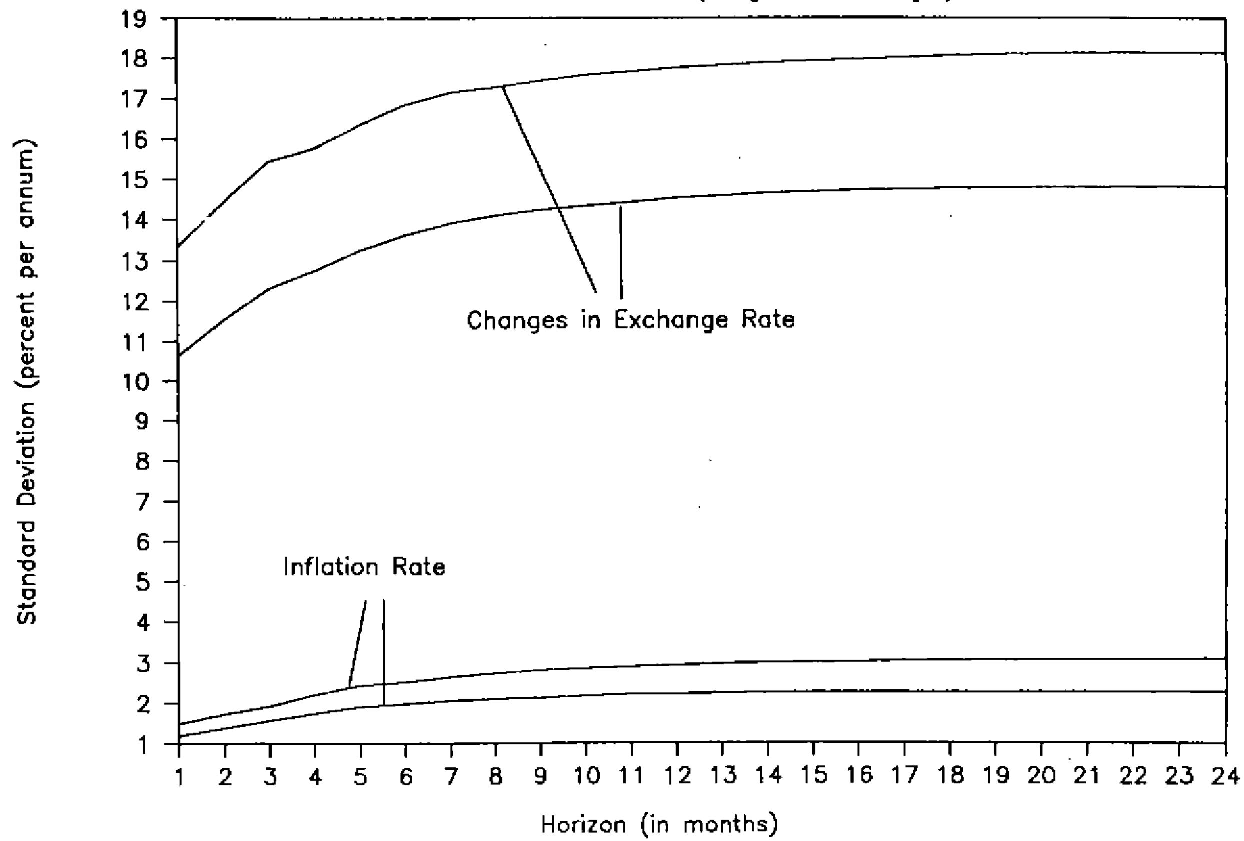


Fig. 2: Correlations between the Nominal and the Real Exchange Rate

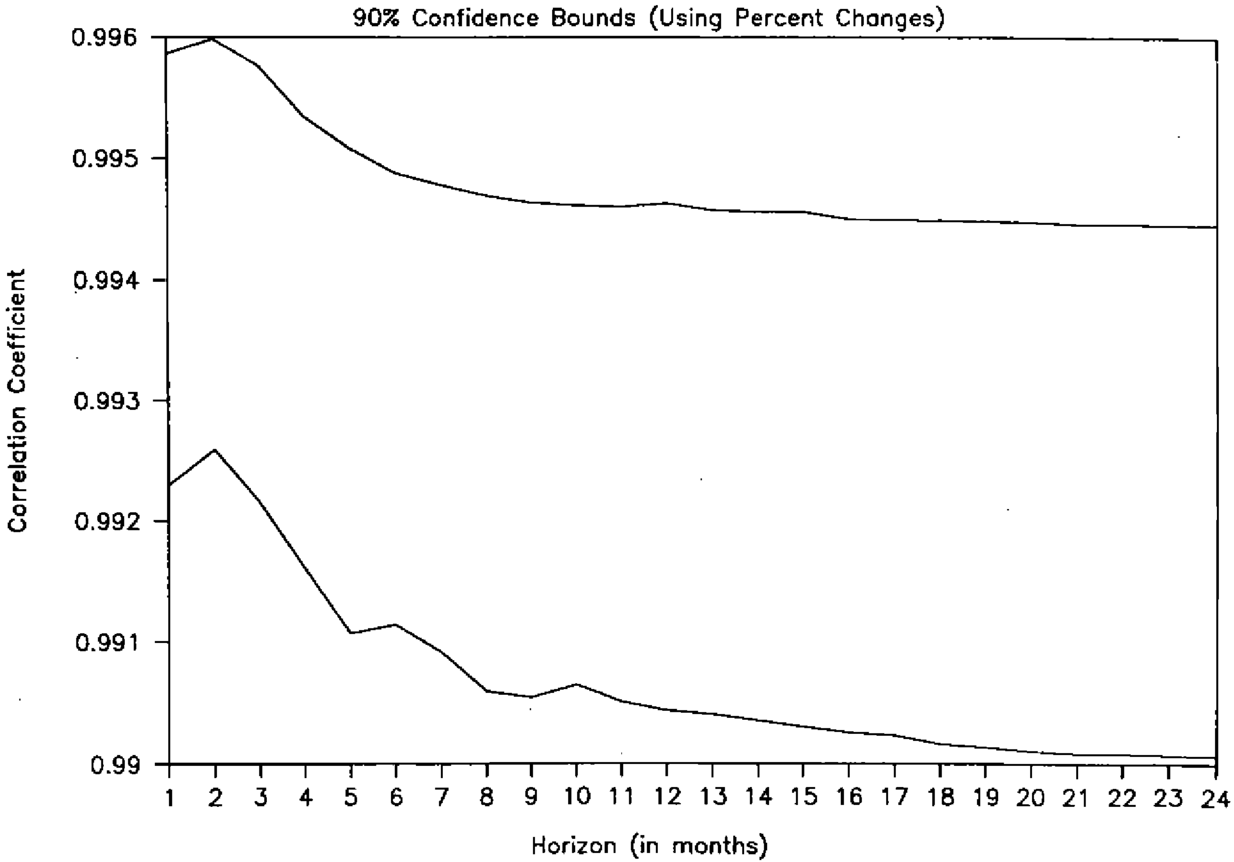


Fig. 3: Correlations between Inflation and Exchange-Rate Changes

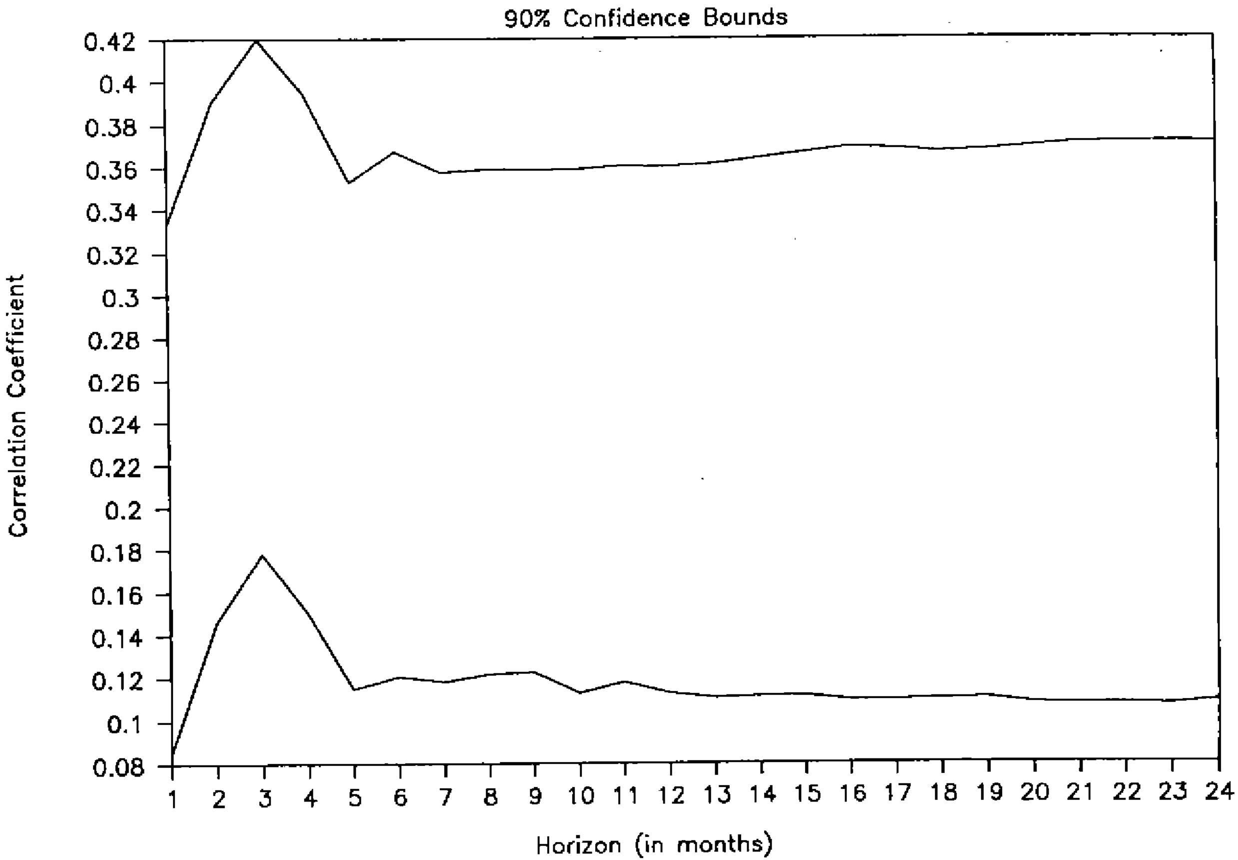


Fig. 4: Correlations between Real and Nominal Exchange Rates

Comparing Restricted and Unrestricted Estimates (Using Percent Changes)

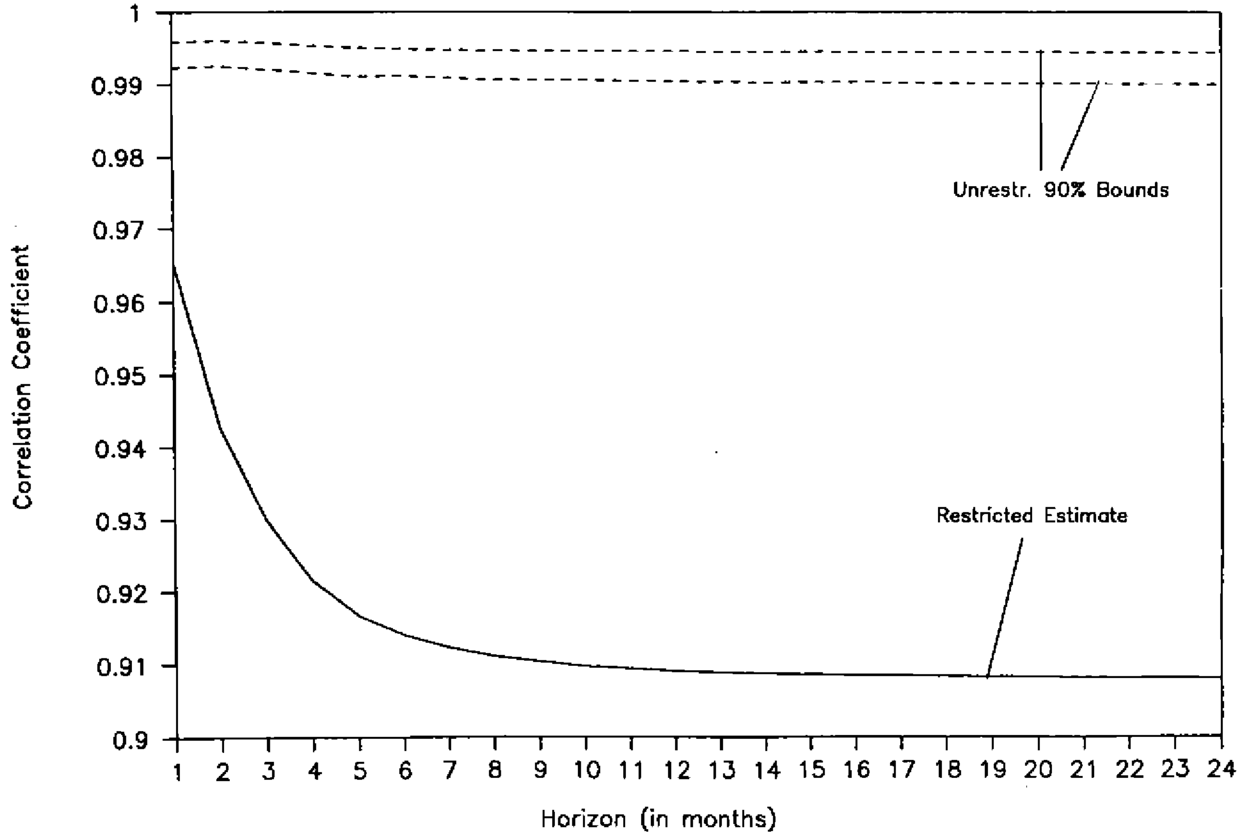
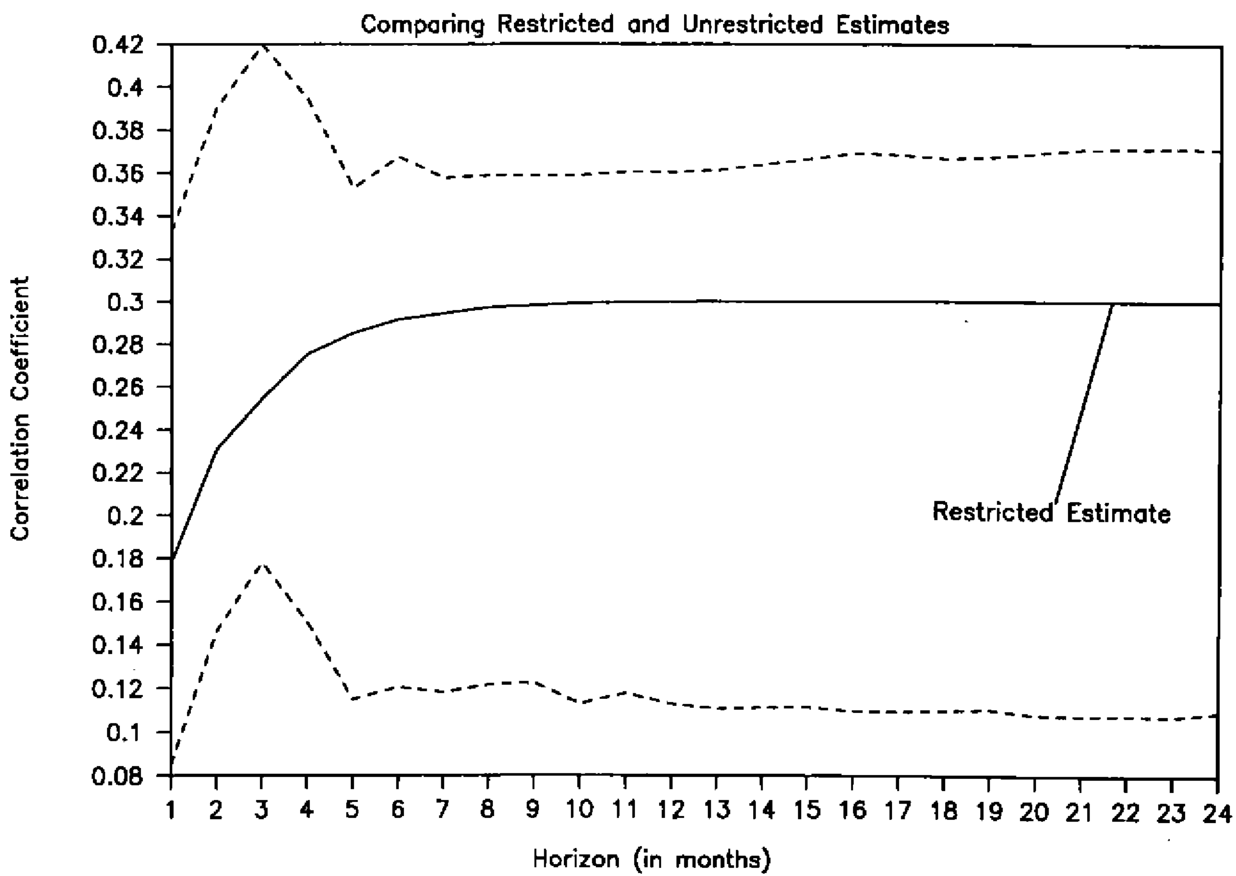


Fig. 5: Correlations between Inflation and Exchange-Rate Changes



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