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**PREMIA IN FORWARD FOREIGN EXCHANGE
AS UNOBSERVED COMPONENTS**

by Theo E. Nijman, Franz C. Palm
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Theo E. Nijman*, Franz C. Palm** and Christian C.P. Wolff**

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ABSTRACT

In this paper, we reconsider the signal-extraction approach to measuring premia in the pricing of forward foreign exchange, put forward in Wolff (1987). We argue that modeling the difference between the forward rate and the associated future spot rate as an ARMA or ARIMA model for the premium buried in a white noise forecast error has the additional advantage of yielding testable restrictions on the observed series, while at the same time indicating whether the parameters of the model for the premium are identified. We present solutions to the identification problem, show how the premium can be estimated in a direct way provided the identification problem is solved and illustrate the results using the series analyzed in Wolff (1987). The results confirm the earlier finding that premia in forward exchange exhibit a certain degree of persistence over time.

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1. INTRODUCTION

The pricing of forward foreign exchange has received considerable attention in the finance literature in recent years. Time-varying premia in the pricing of forward contracts have been documented and characterized in a number of studies, e.g. by Baillie and Bollerslev (1990), Domowitz and Hakkio (1985), Fama (1984), Hansen and Hodrick (1980, 1983), Hodrick and Srivastava (1986) and Wolff (1987). The empirical results in these studies can be interpreted against the background of equilibrium models of international asset pricing such as Adler and Dumas (1983), Fama and Farber (1979), Solnik (1974) and Stulz (1981).

Along with other authors, Wolff (1987) divides the forward foreign exchange rate observed at time t for currency to be delivered at $t+1$ into an expected future spot rate component and a premium component

$$F_t^{t+1} = E[S_{t+1}|t] + P_t \quad (1)$$

where F_t^{t+1} is the natural logarithm of the forward rate at time t for an exchange at $t+1$, $E[S_{t+1}|t]$ is the conditional or rational expectation of the log of the spot rate at $t+1$, conditional on all information available at t and P_t is a premium term. Subtracting S_{t+1} from both sides of equation (1) and defining $v_{t+1} \equiv E[S_{t+1}|t] - S_{t+1}$, we obtain

$$F_t^{t+1} - S_{t+1} = P_t + v_{t+1}, \quad (2)$$

where $[v_t]$ is an uncorrelated, zero-mean sequence with variance σ_v^2 . As noted in Wolff (1987), equation (2) states that the forecast error resulting from the forward rate as a prediction of the future spot rate consists of a premium component and a white-noise error term due to the arrival of new information between t and $t+1$ concerning the spot rate at $t+1$. The premium

component P_t is referred to as the signal that will be assumed to be generated by an ARMA or ARIMA-model. As noted in Wolff (1987), the process for P_t can be derived from a model of international asset pricing. For instance, under the assumptions made by Domowitz and Hakkio (1985) that the utility functions are of the Cobb-Douglas form, that the log of the initial endowments with money and goods in both countries follow independent AR(1)-processes and the additional assumptions that domestic and foreign currency have a GARCH(1,1) and a homoskedastic disturbance, respectively, Lucas' (1982) model leads to a first order autoregression for P_t . The component v_{t+1} is a noise added to the signal which has to be extracted. The advantage of the signal extraction approach proposed by Wolff (1987) is that the expectation of S_{t+1} given information up to period t does not have to be modeled. One can for instance abstract from announcement effects which affect both F_t^{t+1} and $E[S_{t+1}|t]$ without affecting P_t .

The problem that one faces involves checking the order to the process for P_t , identifying and estimating its parameters and extracting a signal from the observations on $y_{t+1} \equiv F_t^{t+1} - S_{t+1}$.

In this paper we argue that modeling the difference between the forward rate and the associated future spot rate as an ARMA or ARIMA model has the advantage of yielding testable restrictions on the observed series while at the same time indicating whether the parameters of the model for the premium are identified.

The paper is organized as follows. In section 2, we discuss the specification, identification and estimation problem of models in which the premium is modeled as an unobserved component. Section 3 contains the results of an empirical study of premia as unobserved components in the three bilateral exchange rates analyzed by Wolff (1987). In section 4, the spot rate is assumed to be generated as a random walk. The premia then become obser-

vable and are modeled using a univariate time series process. Section 5 concludes this paper.

2. MODELING THE PREMIUM AS AN UNOBSERVED COMPONENT

To specify the model for y_t completely, let us assume that P_t is generated by the following ARMA(p,q) model

$$\phi(L)P_t = \theta(L)a_t, \quad (3)$$

where $\phi(L)$ and $\theta(L)$ are polynomials in the lag operator L of degree p and q respectively and a_t is a white noise with mean zero and constant variance σ_a^2 . The roots of $\phi(L)$ are outside the unit circle, those of $\theta(L)$ are different from the roots of $\phi(L)$ and are outside or on the unit circle. Substitution of expression (3) into (2) yields the following model for y_t

$$\phi(L)y_t = \theta(L)a_{t-1} + \phi(L)v_t, \quad (4)$$

which is an ARMA (p^*, q^*) model with $p^* \leq p$ and $q^* \leq \max(p, q)$. If the first difference of P_t is generated by an ARMA(p,q) process of the form in (3), the resulting model for y_t will be an ARIMA ($p^*, 1, q^*$) process with $p^* \leq p$ and $q^* \leq \max(p, q+1)$. The right-hand side of model (4) can be expressed as an MA(q^*) process with innovation ϵ_t yielding the following expression for y_t

$$\phi(L)y_t = \omega(L)\epsilon_t, \quad (5)$$

where $\omega(L)$ is a polynomial of degree q^* in L and ϵ_t has mean zero and variance σ_ϵ^2 . The ARMA model in (5) can be analyzed using univariate time series methods. For instance, the order of p^* and q^* can be determined from the estimated autocorrelation and partial autocorrelation functions of the series y_t . Also, we can test for the presence of a unit root in y_t using

the Dickey-Fuller test or a variant of it, if there are reasons to believe that the process for P_t is integrated of order one.

A comparison of (4) and (5) immediately shows the nature of the identification problem. Under the usual regularity conditions the parameters of (5) are identified, i.e. we are able to determine unique values for the parameters of $\phi(L)$, $\omega(L)$ and σ_ϵ^2 from the second order moments of y_t . The question now arises whether it is possible to uniquely determine the parameters in $\theta(L)$, σ_a^2 and σ_v^2 from those in $\omega(L)$ and σ_ϵ^2 once $\phi(L)$ is identified. This is possible if $p \geq q+1$, a result which follows from Hotta (1989) [for similar results in a different context see e.g. Palm and Nijman (1984)].

To illustrate this point, assume that $p=0$ and $q=1$. Then expressions (4) and (5) specialize to become respectively

$$y_t = (1-\theta L)a_{t-1} + v_t \quad (6)$$

and

$$y_t = (1-\omega L)\epsilon_t \quad (7)$$

For the second moments of y_t , we then have

$$E y_t^2 = \sigma_a^2 (1+\theta^2) + \sigma_v^2 = \sigma_\epsilon^2 (1+\omega^2) \quad (8.a)$$

$$E y_t y_{t-1} = -\theta \sigma_a^2 = -\omega \sigma_\epsilon^2 \quad (8.b)$$

$$E y_t y_{t-i} = 0, \quad i > 1 \quad (8.c)$$

The three unknown parameters θ , σ_a^2 and σ_v^2 cannot be uniquely determined from the information in $E y_t^2$ and $E y_t y_{t-1}$ (or equivalently from σ_ϵ^2 and ω). This implies that the log-likelihood function does not have a well defined maximum. The values for the maximum of the likelihood function reported in Wolff (1987) for the dollar/mark and the dollar/yen rates therefore have to

be an artifact of numerical approximations in finite samples or imprecision caused by the complexity of the Kalman filter algorithm. As $|\theta| \leq 1$, we can determine a lower bound $\sigma_a^2 \geq E y_t y_{t-1}$ for the variance σ_a^2 of the innovation of the premium.

When $p=1$ and $q=1$, expressions (4) and (5) become respectively

$$(1-\phi L)y_t = (1-\theta L)a_{t-1} + (1-\phi L)v_t \quad (9)$$

and

$$(1-\phi L)y_t = (1-\omega L)\epsilon_t. \quad (10)$$

The second moments of y_t satisfy the equations

$$E y_t^2 - \phi E y_t y_{t-1} = (1+\theta^2)\sigma_a^2 + (1+\phi^2)\sigma_v^2 = (1+\omega^2)\sigma_\epsilon^2 \quad (11.a)$$

$$E y_t y_{t-1} - \phi E y_{t-1}^2 = -\theta\sigma_a^2 - \phi\sigma_v^2 = -\omega\sigma_\epsilon^2. \quad (11.b)$$

$$E y_t y_{t-i} = \phi E y_{t-1} y_{t-i}, \quad i > 1, \quad (11.c)$$

from which the parameters θ , σ_a^2 , σ_v^2 and ϕ can not be uniquely determined.

It is easily checked that ϕ can be derived from knowledge of the second moments, but that θ , σ_a^2 and σ_v^2 cannot be uniquely determined from them.

However, in many cases, it is possible to derive upper and lower bounds for the variances σ_v^2 and σ_a^2 . In the appendix, we show that if $\phi \geq \omega \geq 0$ or if $\omega \leq \phi \leq 0$, the variance of the spot rate innovation σ_v^2 has an upper bound

$$0 \leq \sigma_v^2 \leq \sigma_\epsilon^2 (1 + \omega)^2 / (1 + \phi)^2, \quad (12.a)$$

and the variance σ_a^2 satisfies the following upper and lower bounds

$$\sigma_\epsilon^2 [(1 + \omega^2)\phi - (1 + \phi^2)\omega] / (1 + \phi)^2 \leq \sigma_a^2 \leq \sigma_\epsilon^2. \quad (12.b)$$

Note that if we let ϕ go to zero, we obtain again the MA(1) model for y_t with lower bound $-\omega\sigma_\epsilon^2$ for σ_a^2 and upper bound $(1 + \omega^2)\sigma_\epsilon^2$ on σ_v^2 .

When $\theta=0$, i.e. when $q=0$, the condition $p \geq q+1$ is satisfied and the parameters σ_a^2 , σ_v^2 and ϕ can be uniquely determined from the second moments of y_t provided stability and invertibility conditions are imposed.

The identification problem can be solved by introducing additional information of the following kind: (a) a priori restriction, e.g. assuming $\theta=0$ in equation (9), (b) using additional sample information, e.g. analyzing the two series F_t^{t+1} and S_{t+1} in a bivariate framework, a solution that will be adopted in section 3, (c) in the case where the variables are not normally distributed, third and higher moments may contain information that allows to identify the model. Solution (a) is not attractive as it is based on prejudice. Solution (b) uses additional sample information. It is expected to increase the efficiency or power of the statistical methods used. Note however that it depends on the assumption that the optimal forecast of S_{t+1} depends on present and past values of S_t and F_t^{t+1} only. An investigation of (c) is imperative in order to minimize the risk of misspecification.

3. EMPIRICAL RESULTS FOR THE PREMIUM AS UNOBSERVED COMPONENT

In this section we study the premia for the three bilateral exchange rates analyzed in Wolff (1987): the dollar/pound, dollar/mark, and dollar/yen exchange rates. Thirty-days forward rates and subsequently observed spot rates are taken from the Harris Bank Data Base supported by the Center for Studies in International Finance at the University of Chicago. The rates are Friday closes, sampled at four-week intervals covering the period April 6, 1973 to July 13, 1984.

As discussed in Wolff (1987) the autocorrelation and partial autocorrelation functions for $y_t = F_t^{t+1} - S_{t+1}$ suggest an ARMA(1,1) model in the dollar/pound case and MA(1) models in the dollar/mark and dollar/yen cases.

Estimates of these models are presented in Table 1. Large sample standard errors are given within parentheses. In all cases a very insignificant constant has been removed prior to estimation. As shown in the previous section, the ARMA(1,1) model for y_t in the dollar/pound case is consistent with either an AR(1) or an ARMA(1,1) model for the premium. If one is willing to restrict θ a priori to be equal to zero, that is to choose the AR(1) model, both σ_v^2 and σ_a^2 can be identified from ϕ , ω and σ_ϵ^2 . The implied estimates of σ_a^2 , the variance of the innovation in the premium, and σ_v^2 , the variance of the innovation in the exchange rate, are $2.77 * 10^{-4}$ and $3.06 * 10^{-4}$ respectively, which are of course very close to the estimates reported by Wolff (1987) who estimated the same model along different lines. If the premium is not restricted to be generated by an AR(1) model but is allowed to be generated by an ARMA(1,1) model, equations (12.a) and (12.b) yield bounds on the two variances for the dollar/pound case: $0.60 * 10^{-4} < \sigma_a^2 < 6.15 * 10^{-4}$ and $0 < \sigma_v^2 < 4.36 * 10^{-4}$.

The MA(1) models which are identified for y_t in the dollar/mark and dollar/yen cases are only consistent with MA(1) models for the premium. In these cases the structural parameters are not identified, but upper and lower bounds which are derived from (12.a) and (12.b) with $\phi = 0$ are reported in Table 1. Note that, as in Wolff (1987), in all cases there is evidence of nonnegligible time variation of the premium, although the claim in the previous paper that the two variances roughly coincide has to be dropped.

Table I Least Squares Estimates of ARMA Model (5) for $F_t^{t+1} - S_{t+1}$ and Implied Bounds on Structural Parameters (note that σ_ϵ^2 is an upper bound on σ_a^2).

	ϕ	ω	σ_ϵ^2 ($\times 10^4$)	Lower bound σ_a^2 ($\times 10^4$)	Upper bound σ_v^2 ($\times 10^4$)
Dollar/Pound	.458 (.248)	.228 (.265)	6.15	0.60	4.36
Dollar/Mark	0	-.167 (.086)	9.48	1.58	9.74
Dollar/Yen	0	-.195 (.082)	8.28	1.62	8.60

In the next section, additional assumptions on the process of S_t will be introduced with the aim to identify the model for P_t for the dollar/mark and the dollar/yen rates and to check the appropriateness of the AR(1) process for the forward premium in the dollar/pound rate.

4. A BIVARIATE MODEL FOR THE SPOT RATE AND THE FORWARD PREMIUM

As indicated in section 2, identification of the parameter of a moving average process for P_t can be achieved by modeling the expectation of the spot rate conditional on past information. To this end, we complete the model (1) for the premium by specifying the process generating the spot rate. Assuming that the log of the spot rate follows a random walk,

$$S_{t+1} = S_t + v_{t+1}, \quad (13)$$

the premium in (1) becomes an observable

$$F_t^{t+1} - S_t = P_t \quad (14)$$

and its serial correlation structure can be analyzed. If the first differences of S_{t+1} exhibit autocorrelation, e.g. if S_{t+1} is generated by an $ARI(p,1)$ process, we get a bivariate model for S_{t+1} and F_t^{t+1} with cross-equation restrictions

$$S_{t+1} = Y(L)S_t + v_{t+1} \quad (15)$$

$$\begin{aligned} F_t^{t+1} &= Y(L)S_t + F_t \\ &= Y(L)S_t + a_t - \theta a_{t-1}, \end{aligned} \quad (16)$$

where $Y(L) = Y^*(L)\Delta + 1$, $Y^*(L)$ is a polynomial of degree p in L which satisfies the stability condition, Δ is the difference operator, and where the premium is assumed to be a first order moving average. Equations (15) and (16) can be consistently estimated as respectively a univariate AR model and dynamic regression model with MA disturbances, provided $v_{t+1} \perp a_t$, for all t and t' . Joint estimation of (15) and (16) is expected to be more efficient than single equation estimation, except when S_{t+1} is generated by a random walk and v_{t+1} and a_t are independent for all t and t' .

Autocorrelations and partial autocorrelations for ΔS_t and for $x_t = F_t^{t+1} - S_t$ are given in Table II.

For the dollar/mark and the dollar/yen spot rates, a random walk model appears to be appropriate as none of the autocorrelations and partial autocorrelations is significantly different from zero at the five percent level. For the dollar/pound spot rate, an $ARI(1,1)$ model could be entertained as well, as the first serial correlation coefficient is significantly different from zero. An augmented Dickey-Fuller test applied to a regression of S_t on S_{t-1} , an intercept, ΔS_{t-1} and ΔS_{t-2} , (see e.g. Fuller (1976)) yields the values $-.567$, -1.248 and -1.455 for the t -statistic of the dollar/pound, dollar/mark and dollar/yen spot rates respectively.

Table II Autocorrelations and Partial Autocorrelations of ΔS_t and $x_t = F_t^{t+1} - S_t$, 4/6/1973 - 7/13/1984.

	lag	Dollar/Pound		Dollar/Marx		Dollar/Yen	
		ΔS_t	x_t	ΔS_t	x_t	ΔS_t	x_t
<u>Autocorrelation</u>	1	.193*	.855*	.138	.801*	.136	.821*
	2	.105	.734*	-.031	.566*	-.015	.632*
	3	-.046	.629*	.025	.391*	.040	.540*
	4	-.058	.528*	-.114	.287*	.019	.439*
	5	.053	.439*	-.087	.230*	.087	.315*
	6	.099	.371*	-.084	.245*	-.014	.257*
	7	.080	.310*	.012	.281*	-.090	.186*
	8	.034	.265*	.113	.346*	.058	.094
	9	.040	.240*	-.017	.415*	-.060	.111
	10	.070	.228*	.126	.436*	.047	.185*
	11	.093	.206*	.092	.378*	-.058	.195*
	12	.174*	.174*	-.066	.296*	.033	.181*
<u>Partial Autocorrelation</u>	1	.193*	.855*	.138	.810*	.135	.821*
	2	.071	.014	-.051	-.261*	-.034	-.129
	3	-.082	-.006	.038	.076	.048	.186*
	4	-.045	-.041	-.128	.032	.006	-.118
	5	.089	-.014	-.050	.032	.088	-.073
	6	.085	.020	-.079	.171*	-.040	.110
	7	.028	-.009	.039	.045	-.080	-.146
	8	-.000	.024	.094	.182*	.076	-.029
	9	.045	.047	-.055	.116	-.087	.255*
	10	.072	.045	.130	.008	.077	.094
	11	.063	-.034	.039	-.071	-.087	-.023
	12	.135	-.047	-.052	-.000	.085	-.017

"*" indicates statistical significance at the five percent level.

These values are not significantly different from zero at the usual significance levels. Therefore, the hypothesis of a unit root cannot be rejected yielding additional evidence for the appropriateness of the random walk model for these series.

Under the additional assumption that the marginal process for S_t is a random walk, and that the expectation of S_{t+1} conditional on past and present

values of S_t coincides with the expectation conditional on all past and present information, the premium $P_t = x_t = F_t^{t+1} - S_t$ is observable.

Its autocorrelations and partial autocorrelations are reported in Table II as well. The autocorrelations and partial autocorrelations for the dollar/pound premium are consistent with an AR(1) model selected by Wolff (1987) and used in section 3. This finding is further evidence in favor of the outcome from the analysis of the premium as an unobserved component, although the estimate $\hat{\phi} = .855$ is larger than the estimate obtained in the preceding section.

For the dollar/mark and the dollar/yen rates, the autocorrelations slowly die out. An AR(1) model for both the dollar/mark and the dollar/yen rates appears to be consistent with the empirical evidence given in Table II. For the dollar/mark rate, an AR(2) model could also be entertained. Given that several partial autocorrelations are significantly different from zero, an ARMA(1,1) model has also been estimated for the dollar/mark and the dollar/yen series (see Table III). In Table III, estimates for the AR(1) model for x_t are also reported. As the findings of section 3 indicated that an MA(1) process might be appropriate for the dollar/mark and dollar/yen premia, estimates for this model are also given in Table III. Asymptotic t-values are given between parentheses.

For the dollar/pound series, the moving average parameter in an ARMA(1,1) model is not significant at the five percent level. Therefore, the AR(1) model reported in Table III has been selected. The estimated parameter values in Table III confirm the earlier result in Wolff (1987) that premia in forward rates exhibit a certain degree of persistence over time. In fact, the autoregressive coefficient values are somewhat higher than those reported in the earlier study.

Table III Maximum-Likelihood Estimates of the Model for $x_t = F_t^{t+1} - S_t$.

Series	Model	Intercept	ϕ	θ	σ_a^2	$\chi^2(36)$
Dollar/Pound	AR(1)	-.002 (1.843)	.856 (19.869)		35×10^{-5}	21.07
Dollar/Mark	AR(1)	.003 (5.456)	.814 (16.875)		16×10^{-5}	49.77
	ARMA(1,1)	.003 (6.505)	.722 (10.530)	.307 (3.260)	14×10^{-5}	35.45
	MA(1)	.0031 (14.600)		.718 (12.435)	22×10^{-5}	234.65*
Dollar/Yen	AR(1)	.002 (1.396)	.822 (17.415)		99×10^{-5}	41.97
	ARMA(1,1)	.002 (1.683)	.720 (10.517)	.308 (3.273)	95×10^{-5}	39.14
	MA(1)	.0022 (4.130)		.736 (13.107)	143×10^{-5}	173.14*

*** indicates statistical significance at the five percent level.

For all three series, a χ^2 -test based on 36 serial correlations of the residuals, is significant even at a zero percent level for the MA(1) model and indicates that this model is inappropriate. It can easily be verified that $\sigma_v^2 = 5 \times 10^{-4}$ together with the estimates of ϕ , θ and σ_a^2 in Table III would yield $\omega = 0.60$ in the dollar/mark case and $\omega = 0.43$ in the dollar/yen case. Parameter values such as $\phi = 0.72$ and $\omega = 0.60$ are not at all incompatible with insignificance of the parameters in the ARMA(1,1) model because of cancelling roots as well as an estimate of -0.168 for the moving average parameter in the MA(1) model. Similar results hold true for the dollar/yen case.

5. CONCLUDING REMARKS

In this paper we reconsider a methodology put forward in Wolff (1987) to identify and measure premia in the pricing of foreign exchange. The methodology models the premia as unobserved components generated by an ARMA(p,q) process and buried in white noise. In some cases, more specifically when $p < q+1$, the parameters of the process of the premia cannot be identified from observations on $F_t^{t+1} - S_{t+1}$. Only inequality restrictions for some unidentified parameters can be derived. To solve the identification problem additional information has to be used. We analyze the serial correlation structure of the spot price, and model the conditional expectation of S_{t+1} given information up to time period t assuming it to be equal to the conditional expectation of S_{t+1} given $S_t, S_{t-1}, \dots$ only. If this additional assumption is not plausible, one has to model $E(S_{t+1} | t)$ in terms of all the variables which are used to form expectations. The definition of the premium which equals the forward rate minus the expected spot rate yields a second equation of the model. The equations for the spot rate and for the premium can be analyzed separately or jointly. In the empirical part, we report the results of an equation by equation analysis of the spot rate and the forward premium for the rates analyzed by Wolff (1987). For the dollar/pound rate, our analysis confirms the choice of an AR(1) process for P_t , made by Wolff (1987). For the dollar/mark and the dollar/yen rate, the results for the series $F_t^{t+1} - S_t$ suggest an ARMA(1,1) model for P_t . The parameters of this model are identified by using the random walk specification for ΔS_t to obtain $E[S_{t+1} | t] = S_t$ which allows us to measure the premium by $F_t^{t+1} - S_t$. Our empirical results confirm the earlier result in Wolff (1987) that premia in the pricing of forward rates show a certain degree of persistence over time. However, his conclusion that the innovation variance of the premium σ_a^2 and the variance of the innovation in the spot rate are of the same order of magnitude is not confirmed.

Compared with the Kalman filter technique, the use of the ARMA representation of the unobserved component model has the following additional advantages: (a) it yields directly testable restrictions on the second moments of the observed series, (b) the conditions for identification can be easily verified, (c) provided the parameters are identified, they can be estimated fairly easily using the method of asymptotic least squares (ALS), which first estimates the unrestricted ARMA model for y_t and then determines estimates of the parameters of the processes for the unobserved components P_t and v_t by minimizing a quadratic distance between unrestricted estimates of the parameters of the process for y_t and the restricted parameters [see e.g. Kodde et al. (1990)], (d) for stationary processes an estimate for the signal P_t can be obtained using the Wiener-Kolmogorov filter [see e.g. Sargent (1987)].

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APPENDIX: Derivation of bounds for the innovation variances of the premium σ_a^2 and of σ_v^2 when the premium is generated by an ARMA(1,1) model.

The equations for the second moments of y_t in (11.a) and (11.b) are linear in σ_v^2 and σ_a^2 . It is straightforward to solve them and express σ_v^2 and σ_a^2 as a function of the unknown parameter θ and the identifiable parameters ϕ , ω and σ_ϵ^2 to obtain

$$\sigma_v^2 = \sigma_\epsilon^2 [(1 + \theta^2)\omega - \theta(1 + \omega^2)] / [(1 + \theta^2)\phi - \theta(1 + \phi^2)] \quad (\text{A.1})$$

$$\sigma_a^2 = \sigma_\epsilon^2 [(1 + \omega^2)\phi - \omega(1 + \phi^2)] / [(1 + \theta^2)\phi - \theta(1 + \phi^2)]. \quad (\text{A.2})$$

As the two variances have to be positive and as θ is restricted to lie inside the unit interval, the estimates of ϕ , ω and σ_ϵ^2 generally contain some information on σ_v^2 , σ_a^2 and θ . For instance, if $\phi \geq \omega \geq 0$ or if $\omega \leq \phi \leq 0$, the nonnegativity of the variances σ_v^2 and σ_a^2 requires that $\theta \leq \omega$ which implies that

$$0 \leq \sigma_v^2 \leq \sigma_\epsilon^2 (1 + \omega)^2 / (1 + \phi)^2 \quad (\text{A.3})$$

and

$$\sigma_\epsilon^2 [(1 + \omega^2)\phi - (1 + \phi^2)\omega] / (1 + \phi)^2 \leq \sigma_a^2 \leq \sigma_\epsilon^2. \quad (\text{A.4})$$

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