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Monetary Policy and Dividend Growth  
in Germany  
Long-Run Structural Modelling  
versus Bounds Testing Approach

von

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# **Monetary Policy and Dividend Growth in Germany: Long-Run Structural Modelling versus Bounds Testing Approach\***

**Running Title: Monetary Policy and Dividend Growth in Germany**

by

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## **Abstract**

This paper examines the long-run relationship between monetary policy and dividend growth in Germany. For this purpose, we test for cointegration between both variables in the period 1974 to 2003. However, problems related to spurious regression arise from the mixed order of integration of the series used, from mutual causation between the variables and from the lack of a long-run relationship among the variables of the model. We address these problems by applying the bounds testing approach to cointegration in addition to a more standard long-run structural modelling approach. In principle, both procedures are capable of dealing with the controversial issue of the exogeneity of monetary policy vis-à-vis dividend growth. However, the structural modelling approach still leaves a certain degree of uncertainty about the integration properties of the interest rate and the dividend growth. Hence, we feel legitimized to refer to the bounds testing procedure and to conclude that in the longer term short-term rates drive stock returns but not vice versa.

*JEL Classifications:* C22, E52, G12.

*Keywords:* Bounds testing approach to cointegration, Structural modelling, ARDL models, Monetary policy, Stock returns.

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## **I. Introduction**

This paper deals with the impact of monetary policy on stock returns in Germany. It sheds some light on the more general debate on monetary policy and stock returns, that is whether: (a) the central bank as a monopolistic supplier of base money can influence stock returns in a systematic fashion; and (b) if this is the case, whether asset prices should be used as monetary policy indicators. While part (b) of the current debate has been at the centre of theoretical and empirical research for some years now, part (a) still lacks a thorough empirical backing.<sup>1</sup> In principle, it is acknowledged that there are two main channels through which a central bank can influence asset prices. First, the central bank is able to determine short-term interest rates, which act as a benchmark for short-term returns and are used for discounting the assets' future income streams. Thus, the central bank is able to affect asset prices via agents' expectations about the future path of money market rates (short-run impact).

Second, the long-run perspective about future inflation has an impact on the current prices of long-term assets, since nominal long-term returns usually contain an inflation premium. Given that monetary policy determines inflation in the long run, it has a strong impact on asset prices via inflation expectations (long-run impact). However, the short run and the long run are intertwined since, for instance, changes in inflation expectations should cause a break in the sequence of expected short-term rates. This interconnection may serve as the first hint that the use of an error-correction modelling framework, which enables us to model this link between the short and the long run, is highly suitable in this context.

In order to tackle these important questions, we test for the significance of a cointegration relationship between the short-term interest rate (i.e., monetary policy) and stock returns which should ultimately affect stock prices as well. For this purpose,

we apply the bounds testing approach to cointegration originally proposed by Pesaran, Shin and Smith (1996, 2001) and compare the results with those using more standard econometric procedures to estimate the impact of monetary policy on stock returns. As a result, the bounds testing methodology appears to be particularly useful in the current application in at least two dimensions.

First, as claimed for instance by Durham (2003) and Rigobon and Sack (2004), estimating the response of asset prices to changes in monetary policy is complicated by the endogeneity of policy decisions and by the fact that the 'event-study' approach typically used in this context requires a much stronger set of assumptions than ours. We show that the response of asset prices to changes in monetary policy can be singled out and identified based on the procedure proposed by Pesaran, Shin and Smith (1996, 2001) and Pesaran and Shin (1999), respectively.

Hence, in contrast to common instrumental variables procedures, this methodology is capable of addressing the controversial issue of (*lack of*) *exogeneity* of the monetary policy variable. It enables us investigate to the up to now far less explored side of the relationship between monetary policy and the stock market: how stock returns react to changes in monetary policy (Durham, 2003 and Rigobon and Sack, 2004). In this respect, our contribution reaches beyond investigations of asset price booms and monetary policy which look at correlations leaving aside the important question of 'causality' and 'exogeneity' (see, e.g., Detken and Smets, 2004) and for this purpose use a different approach than the heteroscedasticity-based approach applied by Rigobon and Sack (2004).

Second, determining the order of integration of interest rates and stock returns is not an issue when using the bounds test procedure even if there is *no clear information on the integration properties* of the underlying variables. Thus, whether variables should be

introduced in differenced or level form is highly questionable, for instance, within the framework of the Johansen procedure but not within the bounds testing approach to cointegration. Finally, this approach compensates for not applying structural break unit root tests to individual financial time series.

The paper proceeds as follows. Section II discusses our way of modelling monetary policy impacts on stock prices. In section III, we apply the bounds testing procedure on monthly data for Germany. In order to check whether the bounds testing procedure is superior to other approaches, we also compare the empirical results with those obtained from an extended Johansen exercise in section IV. In this section, we analyse cointegration models which contain alternatively the one-month money market rate and the stock returns as an exogenous I(1) variable. Section V concludes and discusses some implications for the current debate about the impacts of monetary policy on asset prices in general.

## **II. Modelling monetary policy impacts on stock returns**

Modelling the relation between the short-term interest rate and the stock market performance, we take a rather pragmatic view. In the tradition of the Capital Asset Pricing Model (CAPM), we assume that there is a linear relation between the stock market performance measure and a risk free interest rate – which is interpreted as the central bank short-term interest rate – plus a risk premium which is assumed to be stationary (time-invariant):

$$(1) \quad r_t = \beta \cdot rf_t + \phi + \varepsilon_t,$$

where  $r_t$  is the return measure in period  $t$ ,  $rf_t$  the central bank short-term interest rate,  $\phi$  the risk premium and  $\varepsilon_t$  is the noise variable.

Assuming that the short-term interest rate of the central bank actually determines the risk free rate, and, in addition, that the risk premium is a stationary variable, the central bank can be expected to have a systematic impact on stock returns. Put another way, equation (1) would suggest that stock returns and central bank rates are cointegrated.

While it is difficult to assign all of the weight of the  $\beta$  coefficient to central bank policies, it is straightforward to assume that using short-term money market rates as the rf variable monetary policy is dominating  $\beta$ . Although central banks do not directly set the most widely watched indicator of short-term monetary conditions, namely the one-month interest rate, they can nevertheless determine pretty much its evolution. Initially, we have based our analysis on three different future stock return measures (i.e., dependent variables  $r_i$ ), namely (i) the annualised one-month continuously compounded stock returns ( $h$ ); (ii) the annualised one-month dividend growth rates in percent ( $\Delta d$ ); and (iii) the difference between the two ( $h - \Delta d$ ).

*(i) Stock price changes ( $r_i = h$ )*

Stock returns and central bank rates could be cointegrated, if a rise in short-term interests systematically reflects the central bank's policy of adjusting the price of money to improved growth/profit expectations as mirrored by rising stock prices. Either the central bank simply responds passively to the economic environment, or a higher short-term rate is evidence of monetary policy efforts to slow down the economy. In such a case, the central bank takes pre-emptive action against bubbles during the upswing as emphasised for instance by Cecchetti, Genberg and Wadhvani (2002) and follows an "active", or "anti-cyclical" policy approach.

*(ii) Dividend growth ( $r_i = \Delta d$ )*

In principle, the same considerations as with respect to our proxy (i) are valid. However, in the context of dividend growth rates it is important to note that dividends as dependent variables might suffer from a drawback, namely firms' "dividend policy". In the second half of the period under review, firms reduced their share of dividend in relation to total profits quite heavily. This finding could be explained by investors expecting high returns from retained earnings. So whereas actual dividend declined, future expected cash flows might have been increased, thereby translating into rising stock prices. That is to say, firms' dividend policy might have blurred the information content of dividend (growth) in the sample under review.

*(iii) Stock price change minus dividend growth ( $r_i = h - \Delta d$ )*

Again, the same arguments as in (i) apply.

What does the above model show? In empirical terms, the monetary policy variable should not, a priori, be excluded when analysing a *long-term* relationship between the stock return and its determinants. However, some readers might have a strong prior belief that monetary policy shocks cannot have permanent effects on stock returns (see, e.g., European Central Bank, 2002, p. 46). We now apply time series econometric techniques to solve this issue.

### **III. Testing for the existence of long-run stock market relations**

#### **III.1 Stylized facts**

We investigate the empirical relation between short-term interest rates (i.e. monetary policy) and stock returns in Germany over the period August 1974 to September 2003. Following the seminal study by Rigobon and Sack (2003), we use monthly data which were in our case provided by Datastream Primark and calculated three alternative future stock return measures: (i) the annualised one-month continuously compounded stock

returns ( $h$ ); (ii) the annualised one-month dividend growth rates in percent ( $\Delta d$ ); and (iii) the difference between these two return measures ( $h - \Delta d$ ).<sup>2</sup>

The performance measures are calculated over two different holding periods, namely 3 and 12 months. Since we leave lag orders constantly at 12 in our estimations with an eye on the monthly frequency of our data set, the use of lag-orders of higher than 12, e.g. 24, 36 and 48 might be problematic. We use average return measures as – against the backdrop of the rational valuation formula – the forecast performance of current stock prices should generally be better for long-term return measures since these make up a larger part of the stock markets' calculated equilibrium price and, moreover, should be less susceptible to one-off shocks and “peso effects” than highly volatile short-term returns.<sup>3</sup>

After having ensured that there is no problem of “reverse causation”, i.e. that the short-term money market rate really is the ‘forcing variable’ these different stock return measures are then regressed on the one-month money market rate. We experimented with some other proxies of monetary policy, but we finally decided to use the one-month money market rate  $ilm$  (i.e., the DM rate until the end of 1998 and the euro rate from 1999 on). Further details on the series are given in the annex.

For a broad-brush view on the data and to identify possible correlations, Figure 1 shows a cross-plot of one of our stock market return measures, namely German dividend growth, against the one-month money market rate. The chart indicates a significant relationship between  $ilm$  and  $\Delta d_{48}$ . What matters for our empirical work, however, is that the overall relationship between these figures reveals a clear negative relation - rather than being vertical or horizontal. Figure 2 shows the variables under review over time. At first glance, it appears that the one-month money market rate was leading dividend growth by around a double-digit number of months both when interest rates were going



up and when they were falling. Observers might conclude from this apparent relationship that, in Germany, monetary policy “causes” dividend growth - a hypothesis that we seek to test more rigorously in the remaining sections of this paper.

- Figures 1 and 2 about here -

## III.2. Testing for cointegration: The ARDL bounds testing approach

### III.2.1. Theoretical background

An important problem inherent in the usual residual-based tests and even in some system-based tests for cointegration is given by a decisive precondition. One should know with certainty that the underlying regressors in the model, i.e. our monetary policy variable, are integrated of order one (I(1)). However, given the low power of unit root tests there will always remain a *certain degree of uncertainty with respect to the order of integration* of the underlying variables. For this reason, we now make use of the bounds testing procedure proposed by Pesaran, Shin and Smith (1996, 2001) to test for the existence of a linear long-run relationship, when the orders of integration of the underlying regressors are not known with certainty. The test is the *standard Wald or F statistic* for testing the significance of the lagged levels of the variables in a first-difference regression. The involved regression is an error-correction form of an ARDL model in the variables of interest.

More specifically, in the case of an unrestricted error-correction model (ECM), regressions of  $y$  on a vector  $x$ , the procedure as a first step involves estimating the following model:

$$\Delta y_t = a_{0y} + a_{1y} \cdot t + \phi y_{t-1} + \delta_1 x_{1,t-1} + \delta_2 x_{2,t-1} + \dots + \delta_k x_{k,t-1} +$$
$$(2) \quad \sum_{i=1}^{p-1} \psi_i \Delta y_{1,t-i} + \sum_{i=0}^{q_1-1} \varphi_{1i} \Delta x_{1,t-i} + \sum_{i=0}^{q_2-1} \varphi_{i2} \Delta x_{2,t-i} + \dots + \sum_{i=0}^{q_k-1} \varphi_{ik} \Delta x_{k,t-i} + \xi_{ty}$$

with  $\phi$  and  $\delta$ 's as the long-run multipliers,  $\Psi$ 's and  $\varphi$ 's as short-run dynamic coefficients,  $(p,q)$  as the order of the underlying ARDL-model ( $p$  refers to  $y$ ,  $q$  refers to  $x$ ),  $t$  as a deterministic time trend,  $k$  as the number of 'forcing variables', and  $\xi$  uncorrelated with the  $\Delta x_t$  and the lagged values of  $x_t$  and  $y_t$ .

As a second step, one has to compute the usual F-statistic for testing the joint significance of  $\phi = \delta_1 = \delta_2 = \dots = \delta_k = 0$ . However, the asymptotic distributions of the standard Wald or F statistics for testing the significance of the lagged levels of the variables are non-standard under the null hypothesis that there exists no long-run relationship between the levels of the included variables. Pesaran and his co-authors provide *two sets of asymptotic critical values*; one set assuming that all the regressors are  $I(1)$ ; and another set assuming that they are all  $I(0)$ . While these two sets of critical values refer to two polar cases, they actually provide a band covering all possible classifications of the regressors into  $I(0)$ ,  $I(1)$  (fractionally integrated or even mutually cointegrated).

In view of this result, we have as a third step which is to use the appropriate *bounds testing procedure*. The test is consistent. For a sequence of local alternatives, it follows a non-central  $\chi^2$ -distribution asymptotically. This is valid irrespective of whether the underlying regressors are  $I(0)$ ,  $I(1)$  or mutually cointegrated. The recommended proceedings based on the F-statistic are as follows. One has to compare the F-statistic computed in the second step with the upper and lower 90, 95 or 99 percent critical value bounds ( $F_U$  and  $F_L$ ). As a result, three cases can emerge. If  $F > F_U$ , one has to reject  $\phi = \delta_1 = \delta_2 = \dots = \delta_k = 0$  and hence conclude that there is a long-term relationship between  $y$  and the vector of  $x$ 's. However, if  $F < F_L$ , one cannot reject  $\phi = \delta_1 = \delta_2 = \dots = \delta_k = 0$ . In this case, a long-run relationship does not seem to exist.

Finally, if  $F_L < F < F_U$  the inference has to be regarded as inconclusive and the order of integration of the underlying variables has to be investigated more deeply.

The above procedure should be *repeated* for ARDL regressions of *each* element of the vector of x's on the remaining relevant variables (including y) in order to select the so called 'forcing variables'. For example, in the case of  $k = 2$ , the repetition should concern the ARDL regressions of  $x_{1t}$  on  $(y_t, x_{2t})$  and  $x_{2t}$  on  $(y_t, x_{1t})$ . If it is determined that the linear relationship between the relevant variables is in fact not 'spurious', one could for instance proceed and still estimate coefficients of the long-run relationship by means of the ARDL-procedure.

### **III.2.2. Application to German stock market data**

Since the choice of the orders of the included lagged differenced variables in the unrestricted ECM specification can have a significant effect on the test results, models in the stock returns ( $h$ ,  $\Delta d$  or  $h-\Delta d$ , in logs) and the one-month money market rate ( $ilm$ ) are estimated for the orders  $p = q = 2, 3, 4, \dots, 12$ . Finally, in the absence of a priori information about the direction of the long-run relationship between  $h$ ,  $\Delta d$  or  $h-\Delta d$  and the monetary policy variables, we estimate unrestricted ECM regressions of  $h$ ,  $\Delta d$  or  $h-\Delta d$  (as the respective dependent variables  $y$ ) on the "vector" of monetary policy variables ( $x$ ) as well as the reverse regressions of  $x$  on  $y$ . More specifically, in the case of the unrestricted ECM regressions of  $y$  on  $x$ , we re-estimate equation (2) using monthly observations over a maximum sample ranging from August 1974 to September 2003. In view of the monthly nature of observations we set the maximum orders to 12, i.e. we estimate eq. (1) for the order of  $p = q_1 = q_2 = 12$  over the same sample period. It is important to note already at this early stage of investigation that we have to choose  $p$  and  $q$  *quite liberally* in order to endogenise the stock returns.<sup>4</sup>

Since we are interested in the impact of the money market rate, namely of  $ilm$ , but take it for granted that the constant (i.e., the stationary risk premium) also influences stock returns, we distinguish between *three different definitions of stock returns* (cases  $h$ ,  $\Delta d$  and  $h-\Delta d$ , in each of these cases monetary policy stance is approximated by the short-term interest rate  $ilm$  as implied by theory:

- *Model 1*: ( $h$ ,  $ilm$ , intercept), means:  $h$ ,  $ilm$  and a constant included in the long-run relationship,
- *Model 2*: ( $\Delta d$ ,  $ilm$ , intercept), means:  $\Delta d$ ,  $ilm$  and a constant included in the long-run relationship, and
- *Model 3*: ( $h-\Delta d$ ,  $ilm$ , intercept), means:  $h-\Delta d$ ,  $ilm$  and a constant included in the long-run relationship.

The models 1, 2, and 3 each portray an important implication of the theoretical model derived in section II, namely that there is cointegration between monetary policy and stock returns. It is also connected with a second implicit idea inherent in the model insofar as it allows monetary policy to slow down the adjustment to a new stock market equilibrium in the wake of a shock. The core implication of the model derived above is that the one-month money market rate determines German stock returns in the short *and* in the long run. In summary, thus, we would like to highlight that our modelling approach is strictly *guided by theory*.

We now use the data to tell us which of the above cases fits the German stock market data best.<sup>5</sup> Table 1 displays the empirical realisations of the F-statistics for testing the existence of a long-run relationship between the stock return and the one-month money market rate (model 1:  $r_i = h$ , model 2:  $r_i = \Delta d$ , and model 3:  $r_i = h-\Delta d$ ). In all of these cases, the underlying equations pass the usual diagnostic tests for serial correlation of

the residuals, for functional form misspecification and for non-normal and/or heteroscedastic disturbances.

The 90, 95 and 99 percent lower and upper critical value bounds of the F-test statistic dependent on the number of regressors and dependent on whether a linear trend is included or not are originally given in Table B in Pesaran, Shin and Smith (1996) and usefully summarised in Pesaran and Pesaran (1997, Annex C, Statistical Tables, Table F). The critical value bounds for the application without trend are given in the middle panel of this Table F at the 90 percent level by 4.042 to 4.788, at the 95 percent level by 4.934 to 5.764 and at the 99 percent level by 7.057 to 7.815. However, we dispense with the specification assuming a linear trend, since it does not make sense for German interest rates and stock returns. We took the upper bound critical values from these intervals and tabulate them in Table 1 as the relevant conservative benchmarks to check the significance of the cointegration relationships.

According to the empirical realisations of the F-values in Table 1, we find that the null hypothesis of no long-run relationship in the case of unrestricted ECM regressions of the log of stock returns on the one-month money market rate is *rejected* in four cases at  $\alpha = 0.1$  and in one of these cases even at the 5 percent level.

- Table 1 about here -

Overall, our results parts of which are displayed in Table 1 provide *some evidence in favour of the existence of a long-run relationship* between the (future) stock returns (as measured by  $h$ ,  $\Delta d$  or  $h-\Delta d$ ) and the one-month money market rate and the estimated constant, i.e. the risk premium. This is valid at least if we approximate stock returns by the variable  $\Delta d$  and use moving-average (MA) orders of 3 or 12. For all other

specifications of the stock returns, namely  $h$  and  $(h-\Delta d)$ , we do not find any cointegrating relationships except for  $h-\Delta d$  (MA=12).

But in view of the potential endogeneity of monetary policy with respect to stock market performance, it is not possible to know a priori whether monetary policy, i.e. the 1-month money market rate, is the 'long-run forcing' variable for the average future stock return performance.<sup>6</sup> Since we attach the highest importance to this point (although it has not been tackled frequently in the literature so far), we have considered all possible regressions and substituted the change in the stock return  $dh$ ,  $d(\Delta d)$  or  $d(h-\Delta d)$  as the dependent variable in eq. (2) by the change in the one-month money market rate  $d(i1m)$ , in order to test whether this relationship is *spurious* in respect to not actually capturing the 'correct direction of causation'. As such, we must ensure that the future stock return is not among the forcing variables.

The empirical results based on the reversed test equations are displayed in the second column of Table 1. In the case of  $r_i = \Delta d$  and for moving averages of 3 or 12 months, we find that the *direction* of this relation is most likely to be *from the one-month money market rate to the future stock returns*. Hence, we think it is justified to consider the short-term interest rate  $i1m$  as the 'long-run forcing' variable for the stock returns  $\Delta d$ . Analogously, the one-month money market rate  $i1m$  can be regarded as the 'long-run forcing' variable for the explanation of the variable  $\Delta d$  if MA=12. As a consequence, in this case the parameters of the long-run relationship can now be estimated using the ARDL procedure discussed in Pesaran and Shin (1999).

Experimenting with dummies coded as one from October 1987 onwards, from July 1990 on, from August 2001 on and from September 2001 on did not change the results substantially. Moreover, we do not think one needs to be particularly concerned about

small sample issues in our context since our estimations are based on monthly data covering about 30 years, i.e. on more than 300 observations – even for cointegration analysis, this is not a particularly small sample.<sup>7</sup> Hence, one frequently claimed additional important advantage of the bounds test within the ARDL framework over the main alternatives such as the Johansen approach to cointegration is not decisive in our context: that it has better small sample properties. Following Narayan and Smyth (2003), we thus do not apply critical values for small sample sizes which have been made available by Narayan and Smyth (2004, 2004a) but trust in those delivered by Pesaran, Shin and Smith (1996, 2001) for a sample size of 1000 observations and tabulated in Pesaran and Pesaran (1997), pp. 478f.<sup>8</sup>

#### **4. Long-run structural modelling – An application of the Johansen procedure**

To check for robustness and in order to convince the reader that applying ARDL models is really worth the effort, we have also moved to some complementary tests for cointegration on the basis of model 2, the one with the best fit according to Table 1.

In this section, we will use our data set and run through the usual steps involved in developing a cointegrating vector-autoregressive (VAR) model. To do this, we have to test for unit roots of the relevant variables and have to make sure that all variables that enter the VAR are integrated of order one (section IV.1). Then we choose the appropriate lag length using an unrestricted VAR (section IV.2) and a specific treatment of the deterministic elements, e.g. restricted trends and unrestricted intercepts. Finally, we choose the number of cointegrating vectors using the Johansen tests (section IV.3). However, one of the innovations of this paper is that the approach taken involves two alternative ways to treat monetary policy in the context of the often mentioned potential endogeneity problem with respect to stock returns. Initially we follow the literature and assume that the VAR model does not contain any exogenous variables. Later on, we give up this assumption and

analyse a cointegration model which contains the one-month money market rate as an exogenous I(1) variable. When using cointegration analysis in the Johansen-framework (Johansen, 1991 and 1995), we would first need to establish that all the underlying variables are I(1). However, such pre-testing results may adversely affect the test results based on cointegration techniques (Cavanaugh, Elliot and Stock, 1995 and Pesaran, 1997).

#### **IV.1. Unit root tests**

The cointegrating VAR procedure as is generally the case when applying this test presumes that all the variables under investigation are integrated of order one (I(1)), and that we already know the nature of the unconditional mean of the variables in the underlying VAR model, namely whether the variables contained have non-zero means or are trended, and whether the trend is linear. Therefore, it is important that all the variables used are tested for unit roots, using the augmented Dickey-Fuller (ADF) test, and that the nature of the trends in the variables are ascertained, for example by plotting each of them against time (see Figure 2).

Hence, as a first step, we have to determine the orders of integration of the dividend growth rate  $\Delta d3$  and the one-month money market rate  $i1m$ . For this purpose, we test for unit roots in the individual time series by means of a battery of unit root tests. Tables 2 and 3 present the results from the standard ADF t-statistic. Considering the monthly frequency of our observations we admit the levels of the variables to be autoregressive processes of order 12. Hence, we set the maximum orders of lagged first differences in the variables to a maximum of 11 (Pesaran, Shin and Smith, 1996, p. 16 f.). The sample length is chosen as 1974M8 to 2003M9 (with M = months) which corresponds to the average sample of most of the models estimated in this paper. The following tables denote the results of ADF-tests *for the levels* of the variables. The results for the first differences are available on request. The optimal lag orders for the unit root tests are se-



lected by the Akaike (AIC) or the Schwarz information criteria (SIC).

- Tables 2 and 3 about here -

From Tables 2 and 3 it is immediately clear that the null of a unit root in the one-month money market interest rate cannot be rejected if one uses the ADF-test. But for German stock returns, the AIC is minimised at a lag length of 1, while the SBC and the HQC are minimised at the lag length of 2 (see Table 2). It follows that if this standard procedure of optimal lag length selection is undertaken, then the null hypothesis of non-stationarity must be rejected clearly.

However, other types of unit root tests whose results are available on request conveyed a slightly different picture in the sense that evidence of stationarity was less unambiguous. Moreover, our ADF unit root test results also show that the probability for German dividend growth rates  $\Delta d$  to be integrated of order one increases with the moving-average (MA)-order (1, 3, 12, 24, 36, 48) of the dividend growth rate variable  $\Delta d$ . Furthermore, the results in Tables 2 and 3 show that the results are not robust against changes in the sample length. Especially, the order of the ADF-test equation often plays a crucial role in the empirical analysis and, when selecting it, special care must be taken to ensure that it is high enough to make sure that the disturbances in the model are not serially correlated. Hence, we one could also feel legitimised to take the lower ADF-test statistics at a higher lag order as a basis to judge about the non-stationarity of dividend growth. Anyway, the annualised one-month dividend growth rate  $\Delta d_3$  appears to be if at all borderline  $I(0)/I(1)$ , i.e. possibly nearly integrated of order one.

A visual inspection of the dividend growth rates does not help us to make a final judgment, since Figure 2 does not say too much about the unit root property. Rather it begs the question why unit root structural tests are not used, given that it cannot be excluded

ex ante that the series display some breaks (see Figure 2).<sup>9</sup> Given that the variables employed by us tend to be  $I(0)$  and/or  $I(1)$  and the bounds test is applicable irrespective of whether or not the variables are  $I(1)$ , the bounds test appears highly suitable in our context from this angle as well (Islam, 2004, p. 996-997, Narayan and Smyth, 2004, 2004a). From this perspective, applying the bounds test procedure gives credence to the empirical analysis. Moreover, the bounds testing approach compensates for not doing the structural break unit root tests.<sup>10</sup>

However, since the consequences of treating a variable as  $I(1)$  if it is  $I(0)$  are generally less grave than treating a variable as  $I(0)$  if it is  $I(1)$ , we acknowledge the conflicting evidence with respect to the integration properties of the dividend growth rate but finally decide to continue with the Johansen-based long-run structural modelling exercise in this section.<sup>11</sup> For this purpose, we now start with the determination of the lag length of the VAR of our highly stylized monetary model of the German stock market.

#### **IV.2. Determining the optimal lag length of the VAR**

In a second step, we estimate an unrestricted VAR in the two variables  $\Delta d3$  and  $ilm$  (including an intercept as a deterministic variable) for the total sample period 1974M8 to 2003M9 in order to determine the optimal lag length underlying the Johansen cointegration test. In applications of the multivariate analysis it is always worth bearing in mind that the order of the VAR,  $p$ , often plays a crucial role in the empirical analysis. When selecting it, special care must be taken to ensure that it is high enough to make sure that the disturbances  $u_t$  in the VAR model are not serially correlated and that, for a selected  $p$ , the remaining sample for estimation is large enough for the asymptotic theory to work reasonably well. This involves a quite difficult balancing act. For the VAR order selection we rely on the AIC and the Schwarz Bayesian Criterion (SBC). Applied to our data set, their use leads to slightly different choices for  $p$ , and we have to decide on the

best choice of  $p$  for the problem at hand. On the basis of the results which are available on request the AIC selects order 7 and the SBC order 4. However, if one now inspects the estimates of the individual equations in the VAR for these orders, it becomes clear that the values of the information criteria only differ by some decimal points and that the SBC displays a second local minimum at order 7. Hence, the AIC and the SBC suggest that the lag order for the cointegrating VAR to be used in the following is 7.

### **IV.3. Testing for cointegration within the Johansen framework**

As a third step, we now use a cointegrating VAR to determine the number of cointegrating relationships between the two variables  $\Delta d3$  and  $ilm$  again for the whole available sample period. To specify the number of cointegrating (or long-run) relations of our model, 0 or 1, we employ the maximum eigenvalue and the trace statistics advanced by Johansen. Hence, we finally make use of the standard Johansen system approach to test for cointegration among German dividend growth  $\Delta d3$  and the one-month money market rate  $ilm$ .

In the following we make use of Johansen's unified ML-framework for estimation and testing of cointegrating relations in the context of VAR error-correction models. For this purpose, we estimate VAR(7) models with restricted intercepts and no trends. These are referred to as Case II in Pesaran and Pesaran (1997, pp. 133 ff). Case II is relevant for variables which reveal *no clear trend* as in our case of German interest rates and stock returns (see Figure 2).

For these models *the trace and maximum eigenvalue statistics* for testing the null hypothesis of no cointegration ( $r = 0$ ) against the alternative hypothesis that there are  $r = 1$  cointegrating relationships among the variables  $\Delta d3$  and  $ilm$  are summarized in Table 4. In order to gain cross-checking results which are comparable to those derived from the

bounds test procedure, we display the results in the Tables 4 and 5 below. The respective critical values in the following tables depend on the number of endogenous and exogenous regressors and are obtained from Pesaran, Shin and Smith (1997). The econometric model behind the following cointegration analysis is given by the following *general vector error correction model*:

$$(2) \quad \Delta y_t = a_{0y} + a_{1y}t - \Pi_y z_{t-1} + \sum_{i=1}^{p-1} \Gamma_{iy} \Delta z_{t-i} + \Psi_y w_t + u_{ty},$$

with

$$z_t = \begin{pmatrix} y_t \\ x_t \end{pmatrix},$$

where  $a_{0y}$  and  $a_{1y}$  are  $m_y \times 1$  vectors;  $\Pi_y$  is of order  $m_y \times m$  (with  $m = m_x + m_y$ ) and represents the long-run multiplier matrix;  $\Gamma_{iy}$  (with dimension  $m_y \times m$ ) measures the short-term dynamics; and  $\Psi_y$  (with dimension  $m_y \times q$ ) quantifies the coefficients on the  $I(0)$  exogenous variables. We differentiate between the following types of variables:  $y_t$  as a  $m_y \times 1$  vector of jointly determined (endogenous)  $I(1)$  variables,  $x_t$  as an  $m_x \times 1$  vector of  $I(1)$  exogenous variables,  $w_t$  as a  $q \times 1$  vector of  $I(0)$  exogenous variables, and intercepts and potential linear deterministic trends. We use the following implicit VAR for the included  $I(1)$  exogenous variable (in our case, the short-term money market rate *ilm*):

$$(3) \quad \Delta x_t = a_{0x} + \sum_{i=1}^{p-1} \Gamma_{ix} \Delta z_{t-i} + \Psi_x w_t + v_{tx},$$

assuming that the  $x$  variables are not cointegrated with each other. It is important to note here that by taking into account equation (3) we allow for a *sub-system approach* in the sense that the  $m_x$  vector  $x_t$  is regarded as *structurally exogenous*. In addition, the error terms in this sub-system are uncorrelated with those in the rest of the system. We im-

pose that there are *no error-correction feedbacks* in the equations explaining  $\Delta x_t$ . In other words, we allow for contemporaneous and short-term feedbacks from  $y_t$  to  $x_t$ . However, we rule out such feedbacks in the long-run. Hence, we can interpret the vector  $x_t$  as the '*long-run forcing*' variables or the '*common stochastic trends*' (Pesaran and Pesaran, 1997, pp. 429 f., Pesaran, Shin and Smith, 1997, pp. 5 ff.).

In the following, we first test for cointegration between  $\Delta d3$  and  $ilm$  assuming as usual that both variables enter the cointegrating relation as endogenous regressors (see section IV.3.1). Second, we take up our main argument from section II and reassure that there is no problem of "reverse causation", i.e. that the short-term money market rate really is the 'forcing variable' for German stock returns. For this purpose we test for cointegration between  $\Delta d3$  and  $ilm$ , but now assuming that either  $ilm$  or  $\Delta d3$  enters the cointegrating relation as an exogenous regressor (see section IV.3.2). If cointegration is indicated in case of  $ilm$  as the exogenous regressor (and this is not the case if  $\Delta d3$  is the exogenous regressor), we think it is justified to treat  $ilm$  as the long-run 'forcing variable'. By doing so, we believe we may be able to solve the endogeneity problem emerging in discussions about monetary policies and stock returns.

#### **IV.3.1. Testing for cointegration in models with the short-term interest rate as an endogenous I(1) regressor**

The following tables display the estimated VAR(7) models with restricted intercepts and no trends (case II). For this purpose, the *trace* and *maximum eigenvalue statistics* for testing the null of no cointegration ( $r = 0$ ) against the alternative hypothesis that there are  $r = 1$  cointegrating relations among the variables  $\Delta d3$  and  $ilm$  are displayed.

- Table 4 about here -

Applying the 95% critical value, the Maximum Eigenvalue and the Trace Statistics in-

dicate that there is exactly one cointegrating vector. The null hypothesis that  $r = 0$  is strongly rejected. Both results are strictly in accordance with the results obtained in section III using the bounds testing procedure.<sup>12</sup> However, there still remains some uncertainty concerning the order of integration of the underlying variables which can often be traced back to the low power of the unit roots test in finite samples, especially with respect to the alternative of strong persistence which is relevant in case of the stock return variable. Moreover, one has to take into account the distinguishably different 'spirits' of the Johansen-ML test and the Pesaran, Shin and Smith (1996, 2001)-approach, the latter assuming 'long-run forcing' variables in the long-term German stock market relationship. From this point of view, the intuition is that the above Johansen cointegration tests only make sense as a benchmark (Pesaran, Shin and Smith, 1996, p. 18) in order to reject the possibility that there is *no* long-term relationship. However, it seems to be by far more adequate to base our analysis on certain testable *exogeneity properties* of variables which contribute to an explanation of German dividend growth, especially the often disputed exogeneity of monetary policy, i.e. the one-month money market rate *ilm*.

#### **IV.3.2. Testing for cointegration in models with one exogenous I(1) variable**

We now assume - as clearly suggested by the results in section III that *one* of the I(1) variables in the cointegrating VAR-model, namely the one-month money market rate *ilm*, is a 'long-run forcing' variable, in the sense that in the long-run the one-month money market rate is not 'caused' by the other variable(s) in the model. For this purpose, we make use of the possibility to implement  $x_t$  as an  $m_x \times 1$  vector of I(1) exogenous variables into the vector error correction model (VECM). Hence, this paper applies a generalization of the analysis of cointegrated systems in the context of a VECM advanced by Johansen (1991, 1995). As explained above, we now allow for a sub-system approach in which we may regard a subset of random variables as structurally exoge-

nous; that is, any cointegrating vectors present do not appear in the sub-system VECM for these exogenous variables and the error terms in this sub-system are uncorrelated with those in the rest of the system.

This generalization is particularly relevant in macro-econometric analysis of economies/markets where we believe it is plausible to assume that some of the forcing variables of the model which are integrated of order one (I(1)), such as monetary policy, are exogenous. For instance, this might be the case if the central bank does not react to asset price movements and decides against pre-emptive action in case of an emerging bubble. At the same time, this extension paves the way for a more efficient multivariate analysis of economic time series. In the following, we make use of an efficient conditional estimation of a VECM containing I(1) exogenous variables.

Following the analysis put forward by Pesaran, Shin, and Smith (1997) we have again begun to distinguish the five different Johansen (1995) cases based on the assumed deterministic trending behaviour of the variable levels. While Johansen (1995) treats *all* the I(1) variables in the VAR as *endogenous*, the relevant critical values for cointegration tests, if there is at least one I(1) exogenous variable in the system, are given in Pesaran, Shin and Smith (1997), Tables D1-D5 in their Appendix D. We again only concentrate on case II. The results are displayed in Table 5.

- Table 5 about here -

#### **IV.4. Long-run structural modelling versus bounds testing approach in the light of the empirical results**

Let us now turn to a brief discussion on the above unit root and cointegration test results. In interpreting our unit root test results, we closely follow Narayan and Smyth (2003, 2004, 2004a) and all others who unambiguously stress that this scenario of some variables

are indicated to be  $I(0)$  and others  $I(1)$  - is exactly the scenario in which the bounds testing approach to cointegration is applicable. We believe that its use allows one to reap the greatest benefits.<sup>13</sup> What all of these studies have in common with ours is that they tested the stationarity of the variables using the Augmented Dickey-Fuller or other unit root tests and the results in general suggest that some of the key variables are  $I(0)$ , while the other variables are  $I(1)$ . Using the bounds test appears certainly appropriate under these circumstances.<sup>14</sup> Hence, most empirical work using the ARDL bounds testing procedure totally dispenses with such kind of unit root pre-testing which is especially useful in those instances where some of the variables cannot be rejected to be  $I(1)$  and some are classified as  $I(0)$  by the unit root tests (Pesaran, Shin and Smith, 2001, p. 18).

So does the Johansen procedure lead to similar results to that of the ARDL approach in terms of indicating of how one should model the impact of monetary policy on stock returns, i.e. dividend growth rates? If the answer is yes, what are the main merits of applying the bounds testing procedure? The preceding sections came up with the result that both procedures lead to astonishingly similar results in terms of cointegration.<sup>15</sup> Hence, our claim that we have found a significant long-run relation running from monetary policy on stock returns appears to be built on a broader basis now. If exogeneity is imposed on the one-month money market rate, the existence of no cointegration vector has to be rejected. If, in turn, exogeneity is imposed on the German stock returns, the null hypothesis of no cointegration cannot be rejected any more. This clear result strongly corresponds to the results in section III which are based on the ARDL approach to cointegration. The results again highlight that the one-month money market rate can be considered as the 'forcing variable' for stock returns if defined as the annualised one-month dividend growth rate in percent ( $\Delta d$ ). In general, the results of these traditional cointegration exercises indicate that cointegration properties appear clearly in those instances where exogeneity is imposed



(solely) on the monetary policy variable. This is the important message of our paper in the light of the literature on monetary policy reaction functions and on the impact of monetary policy on asset prices,

What exactly is the value added of applying the bounds-testing procedure? It is widely known that unit root tests have low power, which is especially true in the case of the alternative that the respective time series exhibit a persistent, yet stationary pattern as often claimed for stock returns (Canova, 1994, Payne, 2003). However, the autoregressive distributed lag (ARDL) bounds testing approach set forth by Pesaran et al. (2001) fortunately does not require any assumption as to whether the time series are  $I(1)$  or  $I(0)$ . Unlike other cointegration techniques like Johansen's procedure which require certain pre-testing for unit roots and that the underlying variables to be integrated of order one, the ARDL model provides an alternative test for examining a long-run relationship regardless of whether the underlying variables are  $I(0)$ ,  $I(1)$ , or fractionally integrated (Bahmani-Oskooee and Ng, 2002, p. 150). Accordingly and deviating from the Johansen procedure, this bounds test procedure allows one to make inferences *irrespective of the absence of any knowledge concerning the actual order of integration for the series under investigation* as long as the value of the test statistic falls outside the critical bounds.

Seen on the whole, the bounds testing approach has really been worth the effort since our unit root tests deliver evidence that the integration properties of the series involved are not a priori clear. Hence, if one would have strictly adhered to the ADF-test tables according to the standard econometric rules, the Johansen-style long-run structural modelling exercise which actually delivers astonishingly similar results would not have been tackled at all. However, strictly following Islam (2004) and others who also consider a set of variables in which there is considerable uncertainty about the integration properties of the variables involved, it is generally standard in the literature to interpret the results from the modified

Johansen procedure as a successful additional robustness check of our empirical results based on the bounds testing procedure- independent of the order of integration resulting from traditional unit root tests. Given the considerable size of our sample, the Johansen multivariate estimation approach does not suffer from a small sample size problem and appears appropriate from this angle (Pattichis, 1999, p. 1062).

## **V. Conclusions**

This paper has examined the relationship between monetary policy and stock returns for Germany using time series econometric techniques. A major empirical result is a *one-way* cointegrating relationship between monetary policy and stock returns *from the first to the latter*. Monetary policy cannot be rejected to be exogenous with respect to dividend growth in Germany but not the other way round. Hence, the monetary policy variable can best be characterised as a so-called 'forcing variable' of stock returns. The main findings of interest for policy makers and investors are that: (a) the interest rate-setting by the central bank has had a significant impact on German stock returns, (b) the Bundesbank and also the ECB were in principle able to reduce stock price volatility by diminishing the uncertainty of future rate changes. Hence, volatility spillovers to other financial markets have been avoided. In this way, the monetary authorities governing Germany have delivered an important positive contribution for economic growth since they most probably reduced the option value of waiting with investment decisions.<sup>16</sup>

Following this interpretation, one would also feel inclined to conclude that the empirical results presented indicate that the monetary policy strategies followed by the Bundesbank and the ECB have been able to provide a reliable medium-term orientation for the participants in the German asset markets (Bohl, Siklos and Werner, 2003, p. 24). From a more technical perspective, this result might suggest that rising central bank rates - in response to improved investor profit expectations – triggered an increase in firms' re-

tained earnings ratios, as reinvesting corporate profits were seen as more favourable compared to the pay-out of earnings.

Since monetary policy determines inflation in the long run, it has a strong impact on asset prices via inflation expectations. That said, the stable one-way long-term relation between monetary policy and asset price movements, i.e. stock returns, is established by the fact that the Bundesbank and also the ECB seem to have followed a quite predictable and transparent monetary policy strategy. In the German case, the central bank appears to have surprisingly closely followed the view exemplified for instance by Bernanke and Gertler (2001) which says that monetary policy should remain focused on achieving the macroeconomic goal of low inflation and strive to do no more than deal with the fallout from the unwinding of potential asset price bubbles.

However, our finding that the short-term interest rate is driving dividend growth does not necessarily imply that the central bank has actually taken an “active” or “anti-cyclical” policy approach within the sample. The reason is that a systematic feedback from changes of dividend growth to monetary policy stance would have been a necessary condition for an “active” monetary policy. However, we could not find them empirically. In any case, our evidence strongly indicates, that pre-emptive action against bubbles during the upswing as emphasized for instance by Cecchetti, Genberg and Wadhvani (2002) does not seem to have taken place.

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## Data

Stock market data for Germany was taken from the Thomson Financials' data base; we made use of TOTMKBD(PI) and TOTMKBD(MV). The stock market indices cover around 80% of the stock market capitalisation in Germany.

The following stock return measures were calculated:

$h$  = holding stock returns (capital gains plus dividend returns, presented by the stock market total performance index), expressed as the annualised one-month continuously compounded stock return in percent;

$\Delta d$  = dividend growth, expressed as the annualized one-month continuously compounded stock return in percent; and

$h - \Delta d$  = holding period return minus dividend growth (as defined above).

In the text, a number behind a variable indicates the time horizon under review. For instance,  $h_{36}$  would indicate the holding period return over the coming 36-months. Averages for return measures were used as – against the backdrop of the rational valuation formula – the forecast performance of current stock prices should generally be better for long-term return measures since these make up a larger part of the stock markets' calculated equilibrium price and, moreover, should be less susceptible to one-off shocks and “peso effects” than highly volatile short-term measures.

$ilm$  = one-month-money market rate, DM until December 1998 and Euro from January 1999 (in percent).

## Figures

Figure 1 - German dividend growth and the money market rate (1974M8 to 2003M9)

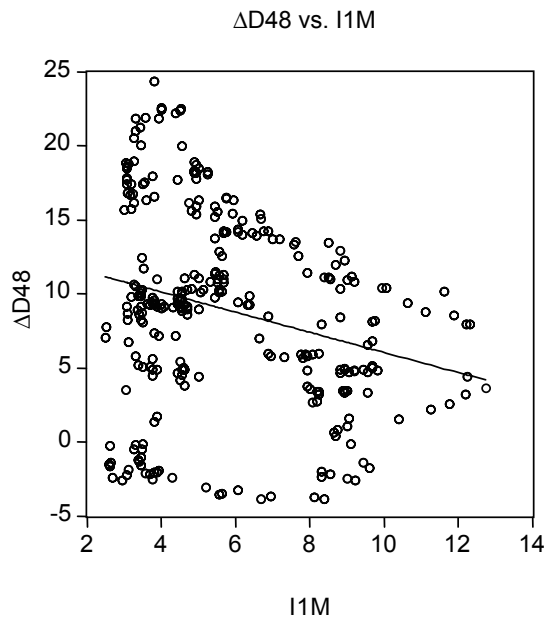
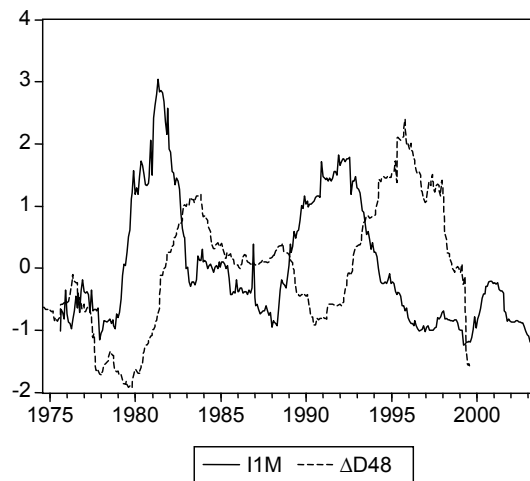


Figure 2 – German dividend growth and the money market rate over time (normalized scaling)



Source: Thomson Financials, own calculations.

## Tables

Table 1 – *F-Statistics for testing the existence of a long-run relationship between the stock return and the short-term interest rate*

<i>MA-order of h</i>	Based on regressions with the change of stock returns as dependent variable	Based on regressions with the change of the one-month money market rate as dependent variable
<i>Model 1: <math>r_i = h</math></i>		
<i>h3</i>	0.33054	0.68269
<i>h12</i>	4.1498	1.1217
<i>Model 2: <math>r_i = \Delta d</math></i>		
<i>Ad3</i>	5.7272	.34943
<i>Ad12</i>	5.7826	.30969
<i>Model 3: <math>r_i = h - \Delta d</math></i>		
<i>(h-Ad)3</i>	1.2670	.67448
<i>(h-Ad)12</i>	5.0548	1.1937
<i>F<sup>C</sup>(0.1)</i>	4.788	4.788
<i>F<sup>C</sup>(0.05)</i>	5.764	5.764
<i>F<sup>C</sup>(0.01)</i>	7.815	7.815

Notes: Lag orders:  $p = q_1 = q_2 = 12$ . Maximum sample: 1974M8 to 2003M9. Individual samples: For MA=12 months: 1975M8 to 2002M9.

Table 2 - *Unit root test results of the dividend growth rate  $\Delta d3$  (ADF Regressions with intercept but without trend)*

335 observations used in the estimation of all ADF regressions; Sample period 1975M8 to 2003M6				
	Test Statistic	LL	AIC	SBC
DF	-7.5253	-1342.3	-1344.3	-1348.2
ADF(1)	-7.6241	-1340.8	-1343.8	-1349.5
ADF(2)	-8.7944	-1332.6	-1336.6	-1344.2
ADF(3)	-5.3370	-1314.1	-1319.1	-1328.6
ADF(4)	-5.4066	-1313.5	-1319.5	-1330.9
ADF(5)	-6.5082	-1305.5	-1312.5	-1325.8
ADF(6)	-4.3336	-1289.5	-1297.5	-1312.8
ADF(7)	-4.6993	-1287.6	-1296.6	-1313.8
ADF(8)	-4.6133	-1287.6	-1297.6	-1316.7
ADF(9)	-3.5259	-1280.3	-1291.3	-1312.2
ADF(10)	-3.3452	-1280.2	-1292.2	-1315.1
ADF(11)	-3.5644	-1279.2	-1292.2	-1317.0

95% critical value for the ADF statistic = -2.8703; LL = Maximized log-likelihood; AIC = Akaike Information Criterion; SBC = Schwarz Bayesian Criterion



Table 3 - Unit root test results of the short-term interest rate *ilm*  
(ADF Regressions with intercept but without trend)

325 observations used in the estimation of all ADF regressions, Sample period 1976M9 to 2003M9				
	Test Statistic	LL	AIC	SBC
DF	-1.3072	-174.6455	-176.6455	-180.4294
ADF(1)	-1.0304	-170.7943	-173.7943	-179.4700
ADF(2)	-1.3251	-165.7947	-169.7947	-177.3623
ADF(3)	-1.6398	-161.0253	-166.0253	-175.4848
ADF(4)	-1.8382	-159.2077	-165.2077	-176.5591
ADF(5)	-2.0156	-157.9092	-164.9092	-178.1526
ADF(6)	-2.1398	-157.2902	-165.2902	-180.4255
ADF(7)	-2.3335	-156.0563	-165.0563	-182.0835
ADF(8)	-2.2496	-155.9838	-165.9838	-184.9030
ADF(9)	-2.2199	-155.9834	-166.9834	-187.7944
ADF(10)	-2.3789	-155.1127	-167.1127	-189.8157
ADF(11)	-2.0162	-151.9978	-164.9978	-189.5927
ADF(12)	-2.5952	-142.3791	-156.3791	-182.8659

95% critical value for the ADF statistic = -2.8706; LL = Maximized log-likelihood; AIC = Akaike Information Criterion; SBC = Schwarz Bayesian Criterion

Table 4 - Cointegration rank statistics applied to German dividend growth and the short-term interest rate

Case II: Restricted intercepts and no trends in the VAR				
328 observations from 1976M3 to 2003M6. Order of VAR = 7				
Variables included in the cointegrating vector:				
$\Delta d3$	<i>ilm</i>	Intercept		
Eigenvalues in descending order:				
.072373	.015420	.0000		
a) Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix				
Null	Alternative	Statistic	95% Critical Value	90% Critical Value
$r = 0$	$r = 1$	24.6411	15.8700	13.8100
$r \leq 1$	$r = 2$	5.0971	9.1600	7.5300
b) Cointegration LR Test Based on Trace of the Stochastic Matrix				
Null	Alternative	Statistic	95% Critical Value	90% Critical Value
$r = 0$	$r \geq 1$	29.7382	20.1800	17.8800
$r \leq 1$	$r = 2$	5.0971	9.1600	7.5300

Table 5 - Cointegration rank statistics in case of one  $I(1)$  exogenous variable

Case II: restricted intercepts and no trends in the VAR					
323 observations from 1976M8 to 2003M6. Order of VAR = 12.					
Variables included in the cointegrating vector: $\Delta d3$ , $i1m$ , Intercept					
<i>A) I(1) exogenous variables included in the VAR: i1m</i>					
Eigenvalues in descending order: .041800, .0000, 0.00					
a) Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix					
Null	Alternative	Statistic	95% Critical Value	90% Critical Value	
$r = 0$	$r = 1$	13.7918	12.4500	10.5000	
b) Cointegration LR Test Based on Trace of the Stochastic Matrix					
Null	Alternative	Statistic	95% Critical Value	90% Critical Value	
$r = 0$	$r = 1$	13.7918	12.4500	10.5000	
<i>B) I(1) exogenous variables included in the VAR: D3</i>					
Eigenvalues in descending order: .012583, .0000, 0.00					
a) Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix					
Null	Alternative	Statistic	95% Critical Value	90% Critical Value	
$r = 0$	$r = 1$	4.0901	12.4500	10.5000	
b) Cointegration LR Test Based on Trace of the Stochastic Matrix					
Null	Alternative	Statistic	95% Critical Value	90% Critical Value	
$r = 0$	$r = 1$	4.0901	12.4500	10.5000	

## Endnotes

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- <sup>1</sup> For this kind of reasoning see, for instance, Bernanke and Gertler (2001), Bohl, Siklos and Werner (2003), Durham (2003), European Central Bank (2002), and Rigobon and Sack (2004).
- <sup>2</sup> The regressions for dividend and profit growth are potentially subject to the omitted variables problem because, in this case, expected stock returns introduce noise. To circumvent this problem, the difference between  $h$  and  $\Delta d$ ,  $h - \Delta d$ , were also calculated and used in the bounds testing procedure.
- <sup>3</sup> See Kaul (1996), p. 284.
- <sup>4</sup> Detailed proofs can be found in Pesaran and Shin (1999) and Pesaran, Shin and Smith (1996, 2001).
- <sup>5</sup> The following estimations - like all other computations in this paper - have been carried out using the 2001 version of the Microfit 4.11 package (see Pesaran and Pesaran, 1997).
- <sup>6</sup> For instance, monetary policy could have systematically and preemptively reacted to the emergence of asset price bubbles. More generally, asset prices as predictors of the future course of the economy might have triggered some monetary policy action. See, for instance, Bean (2004), Dupor and Conley (2004), European Central Bank (2002) and Robinson and Stone (2005) for good summaries of this discussion in the literature.
- <sup>7</sup> If a sample size is small, e.g. if observations take single-digit or low double-digit values, the relevant critical values potentially deviate substantially from the critical values reported in Pesaran, Shin and Smith (2001). Therefore, exact critical value bounds have to be tailored to these sample sizes and are calculated for instance by Narayan and Smyth (2004, 2004a).  
Several previous studies have applied the ARDL approach to sample sizes smaller than ours. Pattichis (1999) uses the ARDL approach to estimate a disaggregated import demand function for Cyprus based on annual data for 1975–1994 (20 observations). Tang (2001) employs the ARDL approach to model inflation in Malaysia using annual data for 1973–1997 (25 observations). Tang (2002) uses it to estimate a money demand function for Malaysia and annual data for 1973–1998 (26 observations), whereas Tang and Nair (2002) apply the approach to estimate an import demand function for Malaysia using annual data for 1970–1998 (29 observations).
- <sup>8</sup> For instance, Narayan and Smyth (2003), p. 1651, use a sample size comparable to ours and, exactly like us, refer to the critical values tabulated in Pesaran and Pesaran (1997), p. 478, Case II.
- <sup>9</sup> For a recent application of a Sen-type unit root test that allows for a simultaneous structural break in the intercept and slope see Narayan (2005). Since the ADF-tests on the first differences in the variables throughout lead to a rejection of the null hypothesis of non-stationarity, at least the variable  $ilm$  (but also  $\Delta d$  at even higher MA-orders) cannot a priori be excluded to be integrated of order one.
- <sup>10</sup> We are grateful for this important argument to an anonymous referee.
- <sup>11</sup> However, these uncertainties that surround the integration properties of our stock return measure indicate that the Pesaran, Shin and Smith (1996, 2001) approach - which does not rely on the exact identification of the order of integration of the underlying variables and has been applied in this paper to test for cointegration between German monetary policy and stock returns - is more robust in our context.
- <sup>12</sup> This appears to be a quite important robustness check since Pesaran, Shin and Smith (1996) admit that a shortcoming of their approach is that is not appropriate in situations in which there are more than one cointegrating vectors. See also Pattichis (1999), p. 1063.
- <sup>13</sup> See among others Bahmani-Oskooee and Ng (2002), p. 150, Faria and Ledesma (2000), pp. 6f., Halicioglu (2004), p. 3, Morley (2003), p. 6, and Payne (2003), p. 1724.
- <sup>14</sup> See Narayan and Smyth (2004), p. 5: "... We tested the stationarity of the variables using the Augmented Dickey-Fuller test and the small sample unit root tests proposed by Elliot et al (1996). To save space the results are not reported, but they suggest that two of the key variables, the robbery and unemployment rates, are I(0), while the other variables are I(1). Using the bounds test is appropriate under these circumstances."
- <sup>15</sup> In contrast to our study, Islam (2004), p. 997, finds diverging results of the bounds testing and the Johansen procedure in a study on the long-run relationship between openness and government size. Hence, it is interesting to cross-check our results based on the bounds testing procedure with the Johansen approach even if there is uncertainty on the orders of integration of the variables involved. See explicitly for this line of reasoning Islam (2004), p. 996.
- <sup>16</sup> See Bean (2004), Dupor and Conley (2004), Domanski and Kremer (1998), pp. 24 and 41, and European Central Bank (2002), p. 39.

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