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The Power Law and Dividend Yields

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Non-Technical Summary

Recent research suggests that the power law is one of the most universal laws in nature and it also seems to work quite well in economics and finance. In this paper we show that the power law explains relatively well the relationship between the value of broad-based market indices and their dividends. In our analysis "power law" translates into a log-linear relationship between stock prices and dividends in levels.

The theoretical part of the paper gives a motivation for the assumption of the log-linear relationship. This relationship is then estimated for six countries: Canada, France, Germany, Japan, United Kingdom and the United States. The results show that such a relationship actually exists for the major stock markets. The only exception is Japan: there are relatively weak hints of a relationship between stock prices and dividends in the period after the crash of 1990.

The estimated relationships can be interpreted as long-run equilibrium between stock prices and dividends. Deviations from this equilibrium lead to an adjustment process in the direction of a new equilibrium. In Japan, the United Kingdom and the United States the stock price is the variable that adjusts towards the new equilibrium whereas in Canada, France and Germany dividends react.

Of particular interest is the parameter which shows the change of the stock prices related to a change of the dividends. For most of the countries, i.e., France, Germany, the United Kingdom and the United States, this parameter is significantly higher than 1. This means that, for example, an increase of the dividends of 1% is accompanied by an increase of the stock prices of more than 1%. For Canada and Japan this parameter is not significantly different from 1. A parameter higher than 1 is consistent with declining relative risk aversion whereas a parameter equal to 1 is consistent with constant relative risk aversion. Decreasing relative risk aversion means that the risk aversion decreases when wealth increases.

To sum up, the results of the empirical part show that the power law seems to have a solid economic foundation for stock prices and dividends.

The Power Law and Dividend Yields

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Abstract

Recent research suggests that the power law is one of the most universal laws in nature and it also seems to work quite fine in economics and finance. In this paper we show that the power law explains extremely well the relationship between the value of broad-based market indices and their dividends. We also show that this relationship is consistent with declining relative risk aversion of the representative investor. Hence, the power law has a solid economic foundation.

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1 Introduction

Recent research suggests that the power law is not only a very universal law in natural sciences but that it is also very important in economics and finance. For example, Zipf (1949), Okuyama et al. (1999), Axtell (2001) and Gabaix and Ioannides (2004) show that the size distribution of many entities follows a power law. Gopikrishnan et al. (1999) find evidence for the "power law distribution" of returns. Related findings are provided by Lux (1996). We mention finally the so called "power law of price impact" (see for example Gabaix et al., 2003) which states that the price impact Δp of a trade of size V scales approximately as $\Delta p \sim V^{0.5}$.[2]

In addition to these empirical findings, there is a vast literature on the predictive power of the dividend yield, see for example Lewellen (2003) and references therein. If aggregate dividends follow approximately a random walk, and there is a linear relationship between aggregate dividends and the value of the market portfolio, then market portfolio returns also follow a random walk and dividend yields have no predictive power. However, dividend yields have predictive power for future returns if the market portfolio is not a linear function of aggregate dividends. Recent theoretical work by Franke et al. (1999) and Lüders and Franke (2004) shows that if the representative investor is not constant relative risk averse then the market portfolio is not a linear function of aggregate dividends and therefore it is not governed by a geometric Brownian motion. Recent empirical findings on option prices support the hypothesis on non-constant relative risk aversion.[3]

To sum up, empirical evidence on different economic relationships leads to the conjecture that the price-dividend relationship is also governed by a power law, i.e. $P_t = KD_t^{\beta}$, where P_t is the price of the asset, D_t is the dividend and β and K are constant parameters. However, an obvious concern would be the lack of a solid economic foundation. Based on a simplified version of Lüders and Franke (2004) we show that the power law is consistent with a representative agent economy with declining relative risk aversion (DRRA). Our empirical analysis for the most important financial markets (Canada, France, Germany, Japan, UK and US) supports our hypothesis.

The paper is organized as follows. In the following section we provide a brief review of the pertinent literature. Section 3 presents the theoretical motivation of the power law. Section 4 presents the empirical results. Section 5 concludes.

[2] Instead of continuing this almost endless list we refer to Gabaix et al. (2003).

[3] See Aït-Sahalia and Lo (2000), Jackwerth (2000) and Rosenberg and Engle (2002).

2 A Brief Literature Review

Several articles have investigated the relationship between dividends and stock prices. The present value model implies a co-integration relationship between stock prices and dividends, when dividends are difference-stationary and the discount factor is constant. (e.g. Timmermann (1995)). Hence in this case we should be able to find a long run-relationship between the two variables.

Timmermann (1995) finds that even with a volatile discount rate co-integration tests are robust as long as the expected returns process is not strongly persistent.

Campbell and Shiller (1987) find quite persistent deviations from the present value model. Their model fits the data poorly, although at high levels of significance it can not be rejected statistically.

LeRoy and Porter (1981) and Shiller (1981) find that with constant discounting the volatility of the stock price is too high to be explained by movements in future dividends. This is known as the excess volatility hypothesis. Taking into account the non-stationarity of the considered time-series, e.g. Shiller (1987) and West (1988) also find that under the assumption of a constant discount factor, stock price movements cannot be explained by dividend variation alone.

Lee (1995) investigates the effect of permanent and temporary shocks to dividends on stock prices in a present value model. He finds that stock prices initially react similarly strongly to both types of shocks, and hence a large part of stock price fluctuations is due to temporary shocks.

Strauss and Yigit (2001) report, inter alia, co-integration between log-dividends and a log-stock price index for the USA for the two periods from 1926 and 1950, respectively, until 1999.

Nasseh and Strauss (2003) support the present value model using panel co-integration and estimation methods. They find that since the mid 1990s the present value model undervalues stock prices by 43%.

However, these papers do not consider the effect of declining elasticity of the pricing kernel on the relationship between the value of the market index and aggregate dividends. Recent theoretical research by Franke et. al. (1999) and Lüders and Franke (2004), however, suggests that declining elasticity of the pricing kernel can have a significant impact on the characteristics of asset prices. As will be shown in the following section, these papers suggest that in contrast to earlier empirical studies a non-linear relationship between asset prices and dividends, i.e. $P_t = KD_t^{\beta}$, with β>1, does not signify any deviation from equilibrium. We will now present a very simplified version of the model of Lüders and Franke (2004) to motivate the power law for the price dividend relationship.

3 Theoretical Motivation

Let us assume a simple efficient exchange economy with one traded risky asset. The value of an asset at time *t*, P_t, is given by

$$P_t = E_t\left(\sum_{h=1}^{\infty} D_{t+h}\frac{\Phi_{t,t+h}}{\left(1+r_f\right)^h}\right) \tag{1}$$

where D_{t+h} is the dividend payment at time *t+h*, r_f is the risk-free interest rate and $\Phi_{t,t+h}$ is the pricing kernel to value claims at time *t* which are due at time *t+h*. E_t is the conditional expectation, where the condition is with respect to all information available at time *t*. Following Lüders and Franke (2004) let us assume that the dividend payments are governed by a geometric random walk, i.e.

$$D_{t+1} = D_t \exp\left(\mu + \sigma\varepsilon_{t+1}\right)$$

with μ and σ constant coefficients, $\varepsilon_{t+1} \sim N(0,1)$ for all *t* and that the pricing kernel $\Phi_{t,t+h}$ is a deterministic function of D_{t+h}. In this case we can write equation (1) as

$$P_t = D_t E_t\left(\sum_{h=1}^{\infty}\frac{\exp\left(\mu h + \sigma\sum_{i=1}^{h}\varepsilon_{t+1}\right)}{\left(1+r_f\right)^h}\frac{\Psi\left(D_{t+h}\right)}{E_t\left(\Psi\left(D_{t+h}\right)\right)}\right) \tag{2}$$

If we further assume that the pricing kernel has constant elasticity, which is equivalent to constant relative risk aversion of the representative investor, i.e. we assume $\Psi\left(D_{t+h}\right) = \left(D_{t+h}{}^{\delta}\right)$, then equation (2) simplifies to $P_t = D_t A_t$, where A_t is a deterministic function of r_f, μ and σ since all D_{t+h} are lognormally distributed. In this case, the asset price follows a geometric random walk. This is the well known result that constant relative risk aversion is consistent with the market portfolio being governed by a geometric random walk.[4] If we assume, instead, that the pricing

[4] See Franke et al. (1999) and Lüders and Franke (2004).

kernel $\frac{\Psi(D_{t+h})}{E_t(\Psi(D_{t+h}))}$ is given by $\frac{D_{t+h}{}^{\beta-1}}{C}$ where C is a constant parameter, then we get that $P_t = KD_t^{\beta}$.

If β is equal to 1, then we have the case of constant elasticity of the pricing kernel. If the pricing kernel has declining [increasing] elasticity, then the risk premium decreases [increases] with increasing D_t and this leads to a higher [smaller] associated increase of P_t than under constant elasticity of the pricing kernel. Hence the elasticity of P_t with respect to D_t is higher [smaller] than 1, i.e. $\beta > 1[\beta < 1]$.[5] While the analysis of Lüders and Franke (2004) reveals that the relationship between asset prices and dividends in general is much more complicated than this, such a rough approximation of the relationship seems at least appropriate for a first empirical analysis.

4 Empirical Analysis

We use the Thomson Financial Datastream Total Market Indices and the corresponding dividend indices for Canada, France, Germany, Japan, the UK and the US. Hence, we consider six of the most important stock markets in the world. We use monthly data for the time period January 1973 until October 2003. All indices are price indices.

Instead of estimating $P_t = KD_t^{\beta}$ we estimate

$$\ln P_t = \alpha + \beta \ln D_t + \varepsilon_t, \tag{3}$$

where ε_t is white noise and $\alpha = \ln K$.

A precondition for estimating equation (3) is that log-stock prices and log-dividends are co-integrated. Hence, our first step is to analyze if there is a co-integrating relationship between the time series of log-prices and log-dividends. This is the case if log-stock prices as well as log-dividends are integrated of order *1* (unit root process) and if there exists a linear relationship between these two variables which is integrated of order *0*.[6]

[5] For a detailed derivation see Lüders and Franke (2004).

[6] For a discussion of the concept of co-integration see e.g. Engle and Granger (1987) or Hamilton (1994).

We use ADF-tests and KPSS-tests to test for non-stationarity. Then, we test for weak exogeneity and test for co-integration next. The last step will be to estimate equation (3) and to test for the long-run parameters α and β.

4.1 Results of the (Non-) Stationary-Tests

We use the Augmented Dickey Fuller (ADF)-test and the KPSS test to analyze whether the logarithmic variables are stationary or integrated.[7] The ADF-test takes the unit root as null hypothesis, hence, the rejection of the null states that the variable is stationary. For each of the six countries the log-dividend and log-stock index series are analyzed. Table 1 shows the results for the levels and the first differences of the time series. The number of lags has been chosen by minimizing the Akaike Information Criterion. The figures in columns 4 and 5 show the estimates for the ADF-tests in case of using either a constant or a constant and a time-trend.[8]

Table 1 gives the results for the total data sample from January 1973 until November 2003. As the Japanese stock index changed dramatically at the beginning of 1990 we also analyze the Japanese time series before and after January 1990. The results of these additional tests are shown in Table 2.

The ADF-tests show that most of the time series are non-stationary in levels and stationary in first differences. But there are some remarkable differences. The stock indices of Canada and the USA appear to be stationary in levels as well as the dividend series of France and the UK.

The KPSS-test has the null-hypothesis „stationarity". The test uses the regression of the time series to be analyzed against a constant or a constant and a time trend. An essential part of the test statistic is the consistent estimation of the variance of the residual time series. According to Hobijn et al. (1998) the automatic lag selection procedure developed by Newey and West (1994) in combination with the Quadratic spectral kernel considerably improves the performance of the test compared to the original KPSS test. Thus, we apply this lag selection procedure in our tests.

Columns 6 and 7 of Table 1 show the results of the KPSS-tests. Almost all of the time series seem to be non-stationary in levels. Only the UK dividend series appears to be I(2) as the null-hypothesis is rejected for the differenced series. This is a strong contradiction of the result of the ADF-test as this test characterizes the UK dividend series as stationary in levels. The KPSS-test also contradicts the outcomes of the ADF-tests regarding the Canadian and US stock indices and the French dividend series.

[7] The KPSS-test has been developed by Kwiatkowski et al. (1992).

[8] The results for the ADF-tests in case of using neither a constant nor a trend are not reported as these results do not change the conclusions.

The time series for Japan seem to be I(1) for the total period under consideration. For the two sub-periods, Jan. 1973 – Jan. 1990 and Febr. 1990 – Nov. 2003, the results are less clear (see Table 2). According to the ADF-test the stock index and the dividend series are I(1) in the first period. In the second period the dividends seem to be I(2) and the stock prices I(0). According to the KPSS-test the dividends (first period) and stock prices (second period) are stationary whereas the series are I(1) during the other periods.

The following co-integration test is based on the assumption that all time series are I(1). For the total time period 1973 – 2003 this assumption is at least supported by the KPSS-tests. The only exception is the UK dividend series for which the results provide no clear indication of I(1)-behavior. For this series and the Japanese series in the two sub-periods the assumption of I(1) is only tentative and the respective results on co-integration and the long-run parameters should be interpreted with caution.

4.2 Results of the Co-integration Tests

We use the bivariate ECM test of Banerjee et al. (1998). As shown by Banerjee et al. this test is superior to alternative single-equation co-integration tests, e.g. the Engle-Granger (1987)-approach. The test is based on the following equation:

$$dY_t = \alpha + \sum_{i=0}^{p1} \beta_i dX_{t-i} + \sum_{i=1}^{p2} \gamma_i dY_{t-i} + \delta Y_{t-1} + \varphi X_{t-1} + \varepsilon_t \qquad (4)$$

The null-hypothesis of no co-integration is rejected if the parameter δ is significantly negative. This parameter is equal to the adjustment coefficient and measures the speed by which the disequilibrium is reduced. The lagged differences of X_t and Y_t are included to avoid autocorrelation of the residuals. The test has been performed in both directions, i.e., Y_t is either the log-dividend series or the log-stock index.

As an additional requirement to apply the test, the X_t series has to be weakly exogenous. The result of these tests is that the null-hypothesis „not weakly exogenous“ could not be rejected in any case. This means that the dividends (X_t) in equation (4) are weakly exogenous for the stock indices (Y_t) and vice versa. As a consequence the co-integration test of Banerjee et al. (1998) can be performed.[9]

Tables 3a and 3b show the estimates for the parameter δ as well as the t-value (in brackets). We have used the simulated critical values of Banerjee et al. (1998, Table I, p. 276) for the significance test. The structure of equation (4), i.e. the lag lengths *p1* and *p2*, has been chosen using the Akaike Information Criterion.

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[9] Details on the results of the exogeneity-tests will be sent by the authors upon request.

Table 3a shows the results for the log-dividends as left-hand-side variable Y_t, whereas Table 3b shows the results for the case „Y_t = log-stock index“. The co-integration tests have been performed for the total period 1973 – 2003. For Japan also the two sub-periods, Jan. 1973 – Jan. 1990 and Febr. 1990 – Nov. 2003, have been analyzed.

Co-integration between dividends and stock indices has been found for all countries with the only exception of Japan for which co-integration seems to exist only in the second sub-period. For France and Japan the null-hypothesis of “no co-integration” could only be rejected at the 10% significance level.

In Canada, France and Germany the dividends adjust to the disequilibrium, whereas in Japan, UK and USA the stock index reacts. The adjustment parameter δ is smallest for USA and France. A full adjustment to a disequilibrium takes about 44 to 50 months, respectively. In UK and Japan the adjustment period is shortest with about 13 to 15 months. The adjustment periods for Canada (37 months) and Germany (26 months) are in-between.

4.3 Estimation and Test of the Long-Run Parameters

The parameters of the long-run equation (3) can be estimated using the following ECM equation (where Δ indicates the first difference of the variable):

$$\Delta Y_t = \sum_{i=-p}^{p} \kappa \Delta X_{t-i} + \sum_{i=1}^{p} \gamma (Y_{t-i} - \lambda X_{t-i} - \mu) + \varepsilon_t \qquad (5)$$

This equation for estimating and testing the long-run parameters has been suggested, for example, by Inder (1993). For a further discussion see also Mills (1999, chapter 7.3).

Equation (5) is estimated by nonlinear least squares. With regard to the results of the co-integration tests equation (5) is estimated for Japan, the UK and the USA using the log-stock index as Y_t and the log-dividend series as X_t. This yields the parameters of interest regarding equation (3): $\beta = \lambda$. For Canada, France and Germany Y_t is the log-dividend series and X_t is the log-stock index because the co-integration tests exhibited that the dividends adjust to a new equilibrium and not the stock prices. For these countries the parameter β of equation (3) is just the inverse of the estimated λ-parameter of equation (5).

Tables 4a and 4b show the parameters of the long-run equation (μ and λ) and the results of the test of the null-hypotheses: λ=0 as well as λ=1. The result of the latter test is given in brackets. The t-statistics have been corrected for heteroskedasticity and autocorrelation using the Newey-West (1987) approach. The lag length p of equation (5) has been chosen according to the Akaike Information Criterion.

The λ-parameter is for all countries significantly different from zero. For most of the countries this parameter is also significantly different from 1. Only for Canada and Japan the null-hypothesis λ=1 could not be rejected at usual significance levels. As for Canada, France and Germany the parameter of interest is the inverse of the estimated λ-parameter it can be concluded that β is significantly higher than 1 for France, Germany, the UK and the USA. Moreover, β being significantly higher than 1 for most of the countries under consideration strengthens the case made by, for example, Franke et al. (1999) and Lüders and Franke (2004) that the pricing kernel seems to have declining elasticity.

5 Conclusion

This paper shows that the well known power law explains the relationship between the value of market indices and aggregate dividends remarkably well. We derive this power law from a simplified version of Lüders and Franke (2004). The theoretical analysis allows us to relate the coefficients to the utility function of the representative investor. Consistent with Franke et al. (1999) and Lüders and Franke (2004) our results suggest that at least during the last 20 years the representative investor had declining relative risk aversion.

Appendix:

Table 1: Results of the ADF-Test (H_0: Non-Stationarity) and KPSS-Test (H_0: Stationarity)

		ADF Test			KPSS Test	
Country	Variable	Lags	Constant	Constant and Trend	Constant	Constant and Trend
Canada	Dividends (level)	8	0.07	-2.18	2.99***	0.251***
	(differences)	7	-7.16***	-7.19***	0.091	0.076
	Stock Index (level)	0	-0.12	-3.16*	3.16***	0.144*
	(differences)	0	-18.23***	-18.22***	0.074	0.047
France	Dividends (level)	15	-1.25	-3.31*	3.22***	0.265***
	(differences)	14	-3.56***	-3.64**	0.115	0.037
	Stock Index (level)	3	-0.47	-3.05	3.18***	0.149**
	(differences)	2	-10.03***	-10.02***	0.080	0.082
Germany	Dividends (level)	6	-0.24	-1.97	2.98***	0.500***
	(differences)	5	-7.80***	-7.81***	0.111	0.083
	Stock Index (level)	1	-0.61	-2.52	3.08***	0.143*
	(differences)	0	-17.99***	-17.97***	0.089	0.088
Japan	Dividends (level)	3	-2.11	-1.80	2.54***	0.741***
	(differences)	2	-8.60***	-8.67***	0.079	0.08
	Stock Index (level)	0	-1.32	-0.50	2.35***	0.664***
	(differences)	0	-18.01***	-18.08***	0.300	0.123*
UK	Dividends (level)	16	-2.59*	0.12	3.19***	0.750***
	(differences)	15	-3.14**	-4.16***	1.44***	0.190**
	Stock Index (level)	3	-0.94	-2.00	3.20***	0.427***
	(differences)	2	-10.50***	-14.22***	0.119	0.091
USA	Dividends (level)	9	-3.94***	-1.71	3.20***	0.663***
	(differences)	8	-4.10***	-18.34***	1.02***	0.056
	Stock Index (level)	0	0.15	-3.15*	3.21***	0.327***
	(differences)	0	-18.85***	-18.86***	0.222	0.145*

Notes: All variables in logarithms. Period: Jan. 1973 – Nov. 2003. Significance levels: *** = 1%, ** = 5%, * = 10%. ADF-Tests: Critical values of MacKinnon (1991). Lags according to the Akaike Information Criterion. KPSS-Tests (applying the Quadratic spectral kernel): Lags according to the automatic lag selection procedure of Newey/West (1994).

Table 2: Results of the ADF-And KPSS-Tests For Japan in Different Time Periods

		ADF Test			KPSS Test	
Period	Variable	Lags	Constant	Constant and Trend	Constant	Constant and Trend
1973:1 – 1990:1	Dividends (level)	0	-0.85	-2.78	2.08***	0.107
	(differences)	0	-13.54***	-13.51***	0.047	0.041
	Stock Index (level)	0	1.57	-2.74	1.94***	0.466***
	(differences)	0	-13.34***	-13.67***	0.606**	0.082
1990:2 – 2003:11	Dividends (level)	6	-2.10	-2.04	1.42***	0.144*
	(differences)	5	-2.11	-1.20	0.200	0.159**
	Stock Index (level)	0	-2.80*	-2.98	0.744***	0.111
	(differences)	0	-12.44***	-12.44***	0.098	0.071

Notes: All variables in logarithms. Significance levels: *** = 1%, ** = 5%, * = 10%. ADF-Tests: Critical values of MacKinnon (1991). Lags according to the Akaike Information Criterion. KPSS-Tests (applying the Quadratic spectral kernel): Lags according to the automatic lag selection procedure of Newey/West (1994).

Table 3a: Results of the Co-Integration Test When Y_t = Log-Dividends, X_t = Log-Stock Index in Equation (4), Estimates for parameter δ

Country	Lags *p1*, *p2*	Parameter Value	T-Statistic
Canada	0, 2	-0.027	-3.42**
France	0, 0	-0.020	-3.07*
Germany	0, 1	-0.039	-4.00***
Japan (from Febr. 1990 on)	0, 0	-0.027	-1.31
UK	0, 0	-0.0006	-0.11
USA	0, 0	-0.0015	-1.05

Notes: Period: Jan. 1973 – Nov. 2003. Significance levels: *** = 1%, ** = 5%, * = 10%. Application of the critical values of Banerjee et al. (1998, Table I, p. 276). Lag lengths *p1* and *p2* according to the Akaike Information Criterion. HAC-corrected t-statistics according to Newey/West (1987).

Table 3b: Results of the Co-Integration Test When Y_t = Log-Stock Index, X_t = Log-Dividends in Equation (4), Estimates for parameter δ

Country	Lags *p1, p2*	Parameter Value	T-Statistic
Canada	0, 0	-0.014	-0.98
France	0, 1	-0.22	-1.95
Germany	0, 0	0.004	0.34
Japan (from Febr. 1990 on)	1, 0	-0.079	-3.12*
UK	0, 1	-0.066	-4.33***
USA	0, 0	-0.023	-3.31**

Notes: Period: Jan. 1973 – Nov. 2003. Significance levels: *** = 1%, ** = 5%, * = 10%. Application of the critical values of Banerjee et al. (1998, Table I, p. 276). Lag lengths *p1* and *p2* according to the Akaike Information Criterion. HAC-corrected t-statistics according to Newey/West (1987).

Table 4a: Long-Run Parameters (Y_t = Log-Dividends, X_t = Log-Stock Index in Equation (5))

Country	Lag Length *p*	Parameter Values and Parameter Tests	
		μ	λ $(= 1/\beta)$
Canada	1	-1.16*	0.60*** (λ=1: not rejected)
France	2	-1.11**	0.69*** (λ=1: *)
Germany	2	-1.48***	0.60*** (λ=1: **)

Notes: Period: Jan. 1973 – Nov. 2003. Significance levels: *** = 1%, ** = 5%, * = 10%. Lag length according to the Akaike Information Criterion. HAC-corrected t-statistics according to Newey/West (1987).

Table 4b: Long-Run Parameters (Y_t = Log-Stock Index, X_t = Log-Dividends in Equation (5))

Country	Lag Length *p*	Parameter Values and Parameter Tests	
		μ	λ $(= \beta)$
Japan (from Febr. 1990 on)	1	4.30***	1.47*** (λ=1: not rejected)
UK	1	2.57***	1.19*** (λ=1: **)
USA	1	1.99***	1.92*** (λ=1: ***)

Notes: Period: Jan. 1973 – Nov. 2003. Significance levels: *** = 1%, ** = 5%, * = 10%. Lag length according to the Akaike Information Criterion. HAC-corrected t-statistics according to Newey/West (1987).

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