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MARKET INDICES**

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De conformidad con la base quinta de la convocatoria del Programa de Estímulo a la Investigación, este trabajo ha sido sometido a evaluación externa anónima de especialistas cualificados a fin de contrastar su nivel técnico.

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SALIENT FEATURES OF DEPENDENCE IN DAILY US STOCK MARKET INDICES*

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ABSTRACT

This paper deals with the analysis of dependence in the US stock market. We focus first on the log-values of the Dow Jones Industrial Average, Standard and Poors 500 and Nasdaq indices, daily from February, 1971 to February, 2007. The volatility processes are also examined based on the squared and absolute values of the returns series. Finally, the “day of the week” effect is also investigated by looking at the orders of integration for each day of the week. A new modelling approach to describe the dependence in this context is also presented. The results suggests that there are very similar patterns for the Dow Jones Industrial Average, Standard and Poors 500. Further, the lowest degrees of dependence are observed on Mondays and Tuesdays.

Keywords: Long range dependence; Volatility; US stock market; Day of week effect.

JEL Classification: C22; G14; G15.

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1. Introduction

There is ample financial literature devoted to analyze, model and forecast the behavior of stock markets. The finance theory suggests many stylized facts in daily stock prices. Thus, for example, mean reversion in stock market prices have been examined in many papers (e.g., Fama and French, 1988; Poterba and Summers, 1988; Lo and MacKinlay, 1988; Kim et al., 1991; Richardson, 1993; Harvey, 1995; Balvers et al., 2000; Chaudhuri and Wu, 2003 and Gropp, 2004 among many others). However, the empirical evidence on mean reversion in stock market prices is still inconclusive. For example, the seminal papers by Fama and French (1988) and Poterba and Summers (1988) documented the mean reversion in the US stock prices, while other authors such as Lo and MacKinlay (1988) presented evidence against mean reversion using weekly US data. Second, the volatility in the stock returns presents an autocorrelation function that decays slowly. Bollerslev and Wright (2000) and Ding et al. (1993) among others found that absolute and squared returns present mean reversion. Third, daily stock prices usually tend to present day of the week effect.¹ The day of the week effect is a relevant stock market anomaly which is extensively documented in the financial literature (see, for example, the initial evidence for the US case in Osborne, 1962; Cross, 1973; French, 1980 and Gibbons and Hess, 1981). This anomaly in the stock market has been recently observed in many other countries (e.g., Gultekin and Gultekin, 1983; Jaffe and Westerfield, 1985; Solnik and Bousquet, 1990; Dubois and Louvet, 1996; and Brooks and Persaud 2001 among others).

In this paper we focus on these stylized facts using long range dependence. Initially, we examine the long range dependence in three daily US stock market indices; then the volatility processes are also examined from a long memory viewpoint. Finally,

¹ Other some calendar anomalies are the January effect and the turn of the month effect among others.

the “day of the week” effect is investigated, proposing a fractional model where the long run dynamics depends exclusively on the day of the week. Long range dependence has been also extensively examined in stock markets (e.g., Lo, 1991; Granger and Ding, 1995a,b; Ding et al., 1993; Baillie et al., 1996; Lobato and Savin, 1998; Sibbertsen, 2004 and recently, Christodoulou-Volos and Siokis, 2006).² Most of the empirical evidence using long range dependence is also inconclusive. Thus, some authors find little or no evidence of long memory in stock markets (e.g., Hiemstra and Jones, 1997, and the references therein). On the other hand, Crato (1994), Barkoulas and Baum (1996), Barkoulas et al. (2000), Sadique and Silvapulle (2001), Henry (2002), and Tolvi (2003) among others, find evidence of long memory in monthly, weekly, and daily stock market returns.

The structure of the paper is as follows. Section 2 briefly describes the statistical models employed in the paper. Section 3 describes the data; In Section 4 we present the main results of the paper, while Section 5 contains some concluding comments and extensions.

2. The statistical models

The statistical models employed across this paper are all based on different versions of fractionally integrated models. This allows a greater degree of flexibility than the standard approaches based on stationarity $I(0)$ or nonstationarity $I(1)$ since the number of differences required to get $I(0)$ series may non-necessarily be an integer number but a real value.

² For example, Lobato and Savin (1998) and Sibbertsen (2004) found long range dependence in volatilities while Christodoulou-Volos and Siokis (2006) using daily data for 34 countries found evidence long memory in returns. We extend previous studies by including both long range dependence in stock returns and volatility in three US stock market indices.

Following the work of Granger (1980), Granger and Joyeux (1980), and Hosking (1981), a rapidly growing body of literature has emerged on fractionally integrated ARFIMA processes. Robinson (1994a, 2003), Beran (1994), and Baillie (1996) present excellent surveys of the literature. A process x_t is integrated of order d if,

$$(1 - L)^d x_t = u_t, \quad t = 1, 2, \dots, \quad (1)$$

with $x_t = 0, t \leq 0$, where u_t is an $I(0)$ process, defined as a covariance stationary process with spectral density function that is positive and finite, and L is the backward shift operator. In the event that d is not an integer, the series x_t requires fractional differencing in order to obtain a stationary (possibly) ARMA series. ARIMA(p, d, q) models in which d is a positive integer are special cases of the general process in (1). If $d > 0$ in (1), x_t is said to be long memory, so-named because of the strong association between observations widely separated in time.

For stock indices, the evidence in favor of long memory may be due to the effect of aggregation. In fact, aggregation is one of the main sources of long memory. The key idea is that aggregation of independent weakly dependent series can produce a strong dependent series. Robinson (1978) and Granger (1980) show that fractional integration can arise as a result of aggregation when: a) data are aggregated across heterogeneous autoregressive (AR) processes, and b) data involving heterogeneous dynamic relationships at the individual level are then aggregated to form the time series. Moreover, the existence of long memory in financial asset returns suggests that new theoretical models based on nonlinear pricing models should be elaborated. Mandelbrot (1971) notes that in the presence of long memory, martingale models of asset prices cannot be obtained from arbitrage. In addition, statistical inference concerning asset pricing models based on standard testing procedures may not be appropriate in the context of long memory processes. (See, e.g., Yajima, 1985, and Barkoulas et al., 2000).

Throughout this paper we focus on Robinson's (1994b) parametric approach, which does not require preliminary differencing; it allows us to test any real value d in (1) encompassing stationary and nonstationary hypotheses. We use the following model:

$$y_t = \beta' z_t + x_t, \quad (2)$$

where y_t is the time series we observe, β is a $(k \times 1)$ vector of unknown parameters; z_t is a $(k \times 1)$ vector of deterministic components, and x_t is given by (1), testing the null hypothesis:

$$H_0 : d = d_0, \quad (3)$$

for any real value d_0 . Thus, the null hypothesized model is:

$$y_t = \beta' z_t + x_t; \quad (1 - L)^{d_0} x_t = u_t, \quad t = 1, 2, \dots$$

Another advantage of this approach is that the limit distribution is standard normal, and this limit behaviour holds independently of the deterministic regressors used in z_t and the type of weak dependence allowed for the $I(0)$ disturbance term u_t . In the final part of the article, equation (1) will be replaced by:

$$(1 - L^5)^d y_t = u_t, \quad t = 1, 2, \dots, \quad (4)$$

to take into account the “day of the week” long memory effects. The functional forms of the test statistics employed in the paper are described in the Appendix.

3. Data description

We used daily data for the Dow Jones Industrial Average (DJIA), Standard and Poors 500 (S&P) and Nasdaq indices from February 5, 1971 to February 7, 2007 yielding 9,395 observations. All data are obtained from the Bloomberg web site (<http://www.bloomberg.com>) and they are all in natural logs.

[Insert Figure 1 and Table 1 about here]

Figure 1 displays the plots of the three US stock markets time series indices: DJIA, S&P and Nasdaq. Apparently the three time series display a very similar shape over the sample period and show a clear upward trend.

Summary statistics on the three US stock markets indices (in logs) and their corresponding returns are shown in Table 1. The Nasdaq index presents the highest volatility over the sample period, while S&P and the DJIA are more stable markets. When we analyze the log stock market prices in first differences (i.e., returns), all series present a similar mean and volatility behavior across the sample period. The skewness and kurtosis coefficients reveal departures from normality in the data, confirmed by the Jarque-Bera statistic.³

4. Salient features of dependence in US stock market indices

This section is divided into various sub-sections. In the first one we examine the degree of dependence of the three series by estimating their orders of integration. Here we employ Robinson's (1994b) parametric approach along with an estimate of d based on the Whittle function. The second sub-section is devoted to the volatility processes, and the same procedures as in the previous sub-section are applied here to the squared and absolute values of the returns, which are used as proxies for the volatility. The third sub-section deals with the "day of the week" effect.

4.1 Initial results on dependence

³ Robinson's methods employed in this article are robust against non-Gaussian disturbances.

Table 2 summarizes the results of Robinson's (1994b) parametric approach in (1) and (2) for the three log-series, assuming that u_t in (1) is white noise; since we must also specify the deterministic components of z_t in (2), we consider the three standard cases, i.e., no deterministic terms (i.e., $z_t = 0$), an intercept ($z_t \equiv 1$), and an intercept and a linear time trend ($z_t = (1, t)^T$). Table 2 shows the test results; the numbers in bold are the maximum likelihood estimates of d obtained with the Whittle function. Table 2 also shows the 95% confidence bands for the non-rejection values of d_0 using Robinson's (1994b) parametric approach.⁴

[Insert Table 2 about here]

The first thing we observe in Table 2 is that the estimated values of d are slightly above 1 in the three series and this happens for the three types of models employed, being higher for the Nasdaq than for the other two indices. Moreover, the confidence intervals for the non-rejection values of d_0 include the unit root (and also some values of d below 1) in all cases with the exception of the Nasdaq with an intercept and/or a linear trend. The fact that d is found to be in most cases equal to or higher than 1 suggests that the markets are efficient according to this simple specification. However, these significant results might be due in large part to the un-accounted-for $I(0)$ autocorrelation in u_t . Thus, we also obtained results using autoregressions for the disturbance term. Table 3 displays the results of Robinson's (1994b) statistic assuming that u_t in (1) is AR(1), and we take d_0 -values in (3) equal to 0, 0.10, ..., 1.40 and 1.50. The limit distribution of the test statistic is unaffected by the inclusion of (weakly) autocorrelated terms, and thus it is still standard $N(0,1)$. A significant feature of these results is that the

⁴ Here, we test H_0 (3) in model given by (1) and (2), using d_0 -values from 0 to 2 in 0.001 increments.

value of the test statistic does not monotonically decrease with d_0 . Such monotonicity is a characteristic that should be expected in these results given correct specification and adequate sample size since the test statistic is one-sided. Thus, for example, if H_0 (3) is rejected with $d_0 = 0.10$ against the alternative $d > 0.10$, the test statistic is significantly positive and above the critical value, and a more significant result in this direction should be expected when $d_0 = 0$ is tested.

[Insert Tables 3 and 4 about here]

We observe in Table 3 that monotonicity is not satisfied if d is small for any model and any series. This lack of monotonicity may be explained in terms of model specification as is argued for example in Gil-Alana and Robinson (1997), though it may also be a consequence of the competition between the fractional differencing parameter and the AR coefficient in describing the nonstationarity. Thus, we observe that if $d = 0$, H_0 cannot be rejected in most of the cases, and, though not reported, the AR coefficients were then very close to 1 in all cases. We finally observe that only if $d > 1$ (in case of no deterministic terms) or if $d > 0.60$ (with an intercept and/or a time trend) is monotonicity satisfied. Table 4 displays the Whittle estimates of d and the 95% confidence bands for those regions of d where monotonicity is achieved. Here, we observe that if no deterministic terms are included, the unit root null hypothesis is rejected for the three series in favour of higher orders of integration; however, if an intercept and/or a linear trend are included, the values of d are strictly smaller than 1 for the DJIA and the S&P, implying a small degree of mean reversion for these two series. On the other hand, for the Nasdaq, d is strictly above 1 in all cases.⁵

⁵ Higher AR orders were also employed and the results were completely in line with those reported here.

The following two figures display the estimates of d using samples of one complete year of observations. In doing so we want to investigate if the order of integration of the series has changed across the years. Figure 2 refers to the case of white noise disturbances while Figure 3 corresponds to the AR(1) case.

[Insert Figures 2 and 3 about here]

Starting with the white noise case we observe that for the DJIA and the S&P, the values decrease slowly across the sample from values above 1 at the beginning of the sample to values below 1 at the end of the sample. For the Nasdaq, the decrease is more pronounced. In fact, for the first half of the sample, the values are strictly above 1 and only for the last ten years there are estimates below unity. A similar pattern is observed if AR(1) disturbances are considered, though here the confidence bands include the unit root in all cases, also for the values corresponding to the Nasdaq index. In general, a slight reduction is observed in the orders of integration of the series across the sample, in some cases, below unity, implying that the market is becoming more inefficient as time goes by, especially for the DJIA and the S&P indices.

4.2 Long memory in the volatility series

We use two alternative measures of volatility: absolute returns and squared returns which have been already used in the financial literature. Absolute returns were employed by Ding et al. (1993), Granger and Ding (1996), Bollerslev and Wright (2000), Gil-Alana (2005), Cavalcante and Assaf (2004), Sibbertsen (2004) and Cotter (2005), whereas squared returns were used in Lobato and Savin (1998), Gil-Alana (2003), Cavalcante and Assaf (2004) and Cotter (2005).

Table 5 displays the confidence bands and the Whittle estimates of d for the absolute and squared returns of the three indices, assuming that the disturbances are white noise. The first thing we observe here is that the values are robust to the three cases of no regressors, an intercept and an intercept with a linear trend. In all cases the confidence bands are constrained in the interval $(0, 0.5)$ implying stationary long memory volatility processes. For the DJIA and the S&P, the values are higher with the absolute returns, while the opposite happens for the Nasdaq index. Thus, for the squared returns, d is around 0.111 for the DJIA; it is 0.126 for the S&P, and is around 0.218 for the Nasdaq. For the absolute values, the corresponding values are around 0.164, 0.166 and 0.210 respectively for the three indices.

[Insert Table 5, Figure 4 and Table 6 about here]

In Figure 4 we display the estimates (and confidence intervals) using samples of 1-year observations and we observe that the behaviour of the estimates are relatively stable across the sample for the three series and the two proxies. In Table 6 we report the average values of d in the volatility processes across years. We note that d is constrained between 0.035 and 0.150 for the DJIA, it is between 0.060 and 0.195 for the S&P, and is between 0.045 and 0.175 for the Nasdaq. Thus, in all cases d is above 0 implying long memory volatility.

4.3 “Day of the week” effect

The day of the week effect is a relevant stock market anomaly which is extensively documented in the financial literature (see, for example, the initial evidence of Osborne, 1962; Cross, 1973; French, 1980 and Gibbons and Hess, 1981). This anomaly in the

stock market has also been observed in many countries (for example, Jaffe and Westerfield, 1985 for Canada, Australia, Japan and the UK; Solnik and Bousquet, 1990 for the case of France; and Brooks and Persaud, 2001 for South Korea, Malaysia, the Philippines, Taiwan and Thailand).

Tables 7 and 8 refer to the same procedures as before, splitting the data by the day of week. Starting with the white noise case (Table 7) we see that for the DJIA and S&P, the estimated d is smaller than 1 for the five days of the week, and, moreover, for Mondays and Tuesdays, the unit root null is rejected in favour of smaller orders of integration. Thus, we obtain evidence of mean reversion of these two days of the week. On the other hand, the Nasdaq index shows strong evidence of no mean-reversion, with values of d strictly above 1 for all days of the week.

[Insert Tables 7 and 8 about here]

Allowing AR(1) disturbances the values are higher in all cases and evidence of $d < 1$ is only obtained for the DJIA index on Mondays. The estimates are also smaller than 1 for Tuesdays and Wednesdays on the Dow Jones, and for Mondays and Tuesdays in the S&P. Similarly to the white noise case, evidence of $d > 1$ is obtained for the Nasdaq across all days. In any case, even for the Nasdaq index, smaller orders of integration are obtained on Mondays and Tuesdays compared with the other days of the week.

The results presented in this subsection indicate that the degree of persistence is different depending on the day of the week, observing lower orders of integration on Mondays and Tuesdays in all series. In what follows, we consider a model where the long memory dynamic effects are supposed to be dependent on the day of the week,

while other short run dynamics are described through the $I(0)$ structure of the disturbance term. In particular, we suppose that the data follow a model of form:

$$y_t = \beta' z_t + x_t, \quad t = 1, 2, \dots, \quad (5)$$

$$(1 - L^5)^d x_t = u_t, \quad t = 1, 2, \dots, \quad (6)$$

$$u_t = \rho u_{t-1} + \varepsilon_t, \quad t = 1, 2, \dots, \quad (7)$$

where z_t in (5) again adopts the three models of no deterministic terms, an intercept, and an intercept with a linear time trend. This model implies that the present value of the series (y_t) depends in the long run on its value five periods before (y_{t-5}), and on all past observations which are backwards multiples of 5, i.e., $y_{t-10}, y_{t-15}, \dots$.⁶ Here, we employ another version of Robinson's (1994b) parametric approach, and the functional form of the test statistic (which is also asymptotically $N(0,1)$ -distributed) is presented in the Appendix. This type of model is relevant in the context of daily financial data, where the value of an asset on a given day of the week may be strongly influenced by its value on the same day of the previous week. There is in fact an extensive literature documenting the presence of calendar anomalies (such as the weekend effect, the day of the week effect, and the January effect) in financial series, both in the US and in other developed markets, dating back to Osborne (1962). Negative Monday returns were found, *inter alia*, by Cross (1973), French (1980), and Gibbons and Hess (1981), the former two analysing the S&P index, the latter the DJIA. Similar findings have been reported for other US financial markets, such as the futures, bond and Treasury Bill markets (e.g., Cornell, 1985; Dyl and Maberly, 1986), foreign exchange markets (Hsieh, 1988), and for Australian, Canadian, Japanese and UK financial markets (e.g., Jaffe and Westerfield, 1985; Jaffe et al., 1989 and Agrawal and Tandon, 1994). Effects on stock market volatility have also been documented (Kiyamaz and Berument, 2003).

Various explanations have been offered for the observed patterns. Some focus on delays between trading and settlement in stocks (Gibbons and Hess, 1981): buying on Fridays creates a two day interest free loan until settlement; hence, there are higher transaction volumes on Fridays, resulting in higher prices, which decline over the weekend as this incentive disappears. Others emphasise a shift in the broker investor balance in buying-selling decisions which occurs during weekends, when investors have more time to study the market themselves (rather than rely on brokers); this typically results in net sales on Mondays, when liquidity is low in the absence of institutional trading (Miller, 1988). It has also been suggested that the Monday effect largely reflects the fact that, when daily returns are calculated, the clustering of dividend payments around Mondays is normally ignored; alternatively, it could be a consequence of positive news typically being released during the week, and negative ones over the weekend (Fortune, 1998). Additional factors which could be relevant are serial correlation, with Monday prices being affected by Friday ones, and a negative stock performance on Fridays being given more weight (Abraham and Ikenberry, 1994); measurement errors (Keim and Stambaugh, 1984); size (Fama and French, 1992); volume (Lakonishok and Maberly, 1990).

Initially, we suppose that there are no short run components (i.e., $\rho = 0$ in (7)) and therefore all the dynamic behaviour of the series is described through the fractional differencing polynomial in (6). The results are displayed in Tables 9 and 10.

[Insert Tables 9 and 10 about here]

⁶ Note that for the three indices, if there was no value for a given day, the arithmetic mean using the previous and the following observation was computed.

Table 9 displays the values of the test statistic, testing again H_0 (3) now in the model given by (5) and (6) for d_0 -values from 0 to 1.50 with 0.10 increments. The first thing we observe in this table is that the value of the test statistic monotonically decreases with d in all cases across the three models presented. It is also observed that the only non-rejection value reported in the table takes place at $d = 1$ for the Nasdaq index in the case of no deterministic terms. However, other non-rejections may occur at values of d in the intervals across the points. Thus, it seems that some non-rejections could take place at values of d between 0.9 and 1 for the DJIA and S&P, and at values constrained between 1 and 1.1 for the Nasdaq index.⁷ Note that these are the values where the test statistic changes its sign and thus we can find values within the $N(0,1)$ confidence interval. Table 10 reports the 95% confidence band of non-rejection values along with the Whittle estimate of d . It is observed that d is below 1 for the DJIA and S&P, while it is slightly above 1 for the Nasdaq index.

[Insert Tables 11 – 13 about here]

Table 11 is similar to Table 9 but imposing an AR(1) structure for the disturbance term. Here we observe that the null hypothesis is rejected in all cases except when $d = 0$, implying that a simple AR(1) model could be an adequate specification for the series. However, though not reported, the AR coefficient was in these cases extremely close to 1, implying once more that the fractional differencing polynomial is competing with the AR parameter in describing the long run effect. Note that the polynomial $(1-L)^5$ can be decomposing into $(1-L)(1+L^2+L^3+L^4)$ implying then the existence of a unit root at the long run or zero frequency, which may compete with the

⁷ The day of the week effect may not be the same across markets. For example, Wang et al. (1997) and Chang et al. (2004), show that the weekend effect was more intense in the Nasdaq than in other US stock

AR coefficient if this is close to 1. Thus, we also employ an alternative to the AR model, which is based on Bloomfield (1973) exponential spectral model. This is a non-parametric approach of modelling the $I(0)$ disturbances, where the spectral density function is given by:

$$f(\lambda; \tau) = \frac{\sigma^2}{2\pi} \exp\left(2 \sum_{r=1}^m \tau_r \cos(\lambda r)\right), \quad (8)$$

where m is the number of parameters required to describe the short run dynamics of the series. Bloomfield (1973) showed that the logarithm of an estimated spectral density function is often found to be a fairly well-behaved function and can thus be approximated by a truncated Fourier series. He showed that the spectral density of an ARMA process can be well approximated by (8). Moreover, this model is stationary across all values of τ , and the model accommodates extremely well in the context of Robinson's (1994b) tests. The results using the model of Bloomfield (1973) (with $m = 1$) are displayed in Tables 12 and 13. Higher orders for m were also tried and the results were practically the same in terms of the non-rejection values of d . First, we observe that monotonicity is again achieved in all cases and the non-rejections take place at values of d constrained between 0.70 and 0.80 for the DJIA and S&P, and between 0.80 and 0.90 for the Nasdaq. The estimated values of d are higher in case of a linear time trend and in all cases d is between 0.74 and 0.81, implying long memory and mean reversion for the “day of the week” effect.

Next we wonder which may be the best model specifications in the context of “day of the week” effects, and use likelihood criteria along with t-tests to determine the best models. It is obtained that the best specification for the DJIA is the following:

$$y_t = 6.72541 - 0.00082t + x_t; \quad (1 - L^5)^{0.790} x_t = u_t; \quad u_t \approx \text{Bloomfield}(\tau = 1.059) \\ (463.64) \quad (-51.75)$$

with the t-values in parenthesis. For the S&P the selected model is:

$$y_t = 4.52984 - 0.00054t + x_t; \quad (1 - L^5)^{0.779} x_t = u_t; \quad u_t \approx \text{Bloomfield}(\tau = 1.052), \\ (347.80) \quad (-41.51)$$

and finally, for the Nasdaq index,

$$y_t = 4.57865 - 0.00054t + x_t; \quad (1 - L^5)^{0.805} x_t = u_t; \quad u_t \approx \text{Bloomfield}(\tau = 1.109). \\ (303.72) \quad (-30.30)$$

Thus, once more we obtain a higher degree of persistence for the Nasdaq than for the other two indices.

5. Conclusions

This paper deals with the analysis of the dependence in the US stock market by means of investigating the orders of integration of the DJIA, the S&P and the Nasdaq indices over the period February 5, 1971 to February 7, 2007. The volatility processes are also examined by looking at the degrees of integration of the absolute and squared returns, which are used as proxies for the volatility. In the final part of the paper we have examined the “day of the week” effect, first by looking at the orders of integration separately for each day of the week. A new modelling approach, based on long run dependence for the week-day effect is also considered.

The results can be summarized as follows: first, we observe very similar patterns for the DJIA and the S&P indices compared with the Nasdaq index. The Nasdaq index only includes technology stocks whereas the DJIA and the S&P indices include stocks from the industrial sector and other sectors and thus are usually good indicators of the stock market as a whole. Thus, the similarities observed in these two indices may also be translated to their degrees of persistence. Starting with the original log-series, we

observe that if all the dependence is captured through the fractional differencing polynomial, the order of integration is slightly above 1, and the unit root cannot be rejected for the DJIA and the S&P. However, allowing weak dependence (AR and Bloomfield) the orders of integration are strictly smaller than 1 (and thus showing mean reversion) for these two indices, while it is strictly above 1 for the Nasdaq. In general, we observe a slight decrease in the degree of dependence across the years in the sample. With respect to the volatility processes, there is strong evidence of some degree of stationary long memory ($0 < d < 0.5$) for the two proxies in the three series, being once more higher in the case of the Nasdaq index. The long range dependence in volatility are also found in Ding et al. (1993), Lobato and Savin (1998) and Sibbertsen (2004). If we separate the data by the day of the week, the lowest degrees of dependence are observed on Mondays and Tuesdays. Modeling the “day of the week” effect by means of a long memory model, mean reversion is obtained in practically all cases, with higher values obtained again for the Nasdaq index. Thus, the Nasdaq seems to be the closest market to efficiency while the DJIA and S&P500 seem to present a small degree of mean reversion.

Appendix

Assuming that y_t is described by equation (2), the regression errors, x_t adopt the form:

$$\rho(L; \theta) x_t = u_t, \quad t = 1, 2, \dots, \quad (\text{A1})$$

where ρ is a scalar function that depends on L and the unknown parameter θ that will adopt different forms as shown below, and u_t is $I(0)$. The function ρ is specified in such a way that all its roots should be on the unit circle in the complex plane, and therefore it includes polynomials of the form $(1-L^k)^{d+\theta}$, where k is an integer and d may be a real value. Thus, in what follows, we assume that

$$\rho(L; \theta) = (1 - L^k)^{d+\theta}. \quad (\text{A2})$$

Robinson (1994b) proposed a Lagrange Multiplier (LM) test of the null hypothesis:

$$H_0 : \theta = 0, . \quad (\text{A3})$$

in a model given by (2) and (A1 – A2). Based on H_0 given by (A3), the estimated β and residuals are:

$$\begin{aligned} \hat{u}_t &= (1 - L^k)^{d_0} y_t - \hat{\beta}' w_t, \\ w_t &= (1 - L^k)^{d_0} z_t; \quad \hat{\beta} = \left(\sum_{t=1}^T w_t w_t' \right)^{-1} \sum_{t=1}^T w_t (1 - L^k)^{d_0} y_t. \end{aligned}$$

The functional form of the test statistic is then given by:

$$\hat{r} = \frac{T^{1/2}}{\hat{\sigma}^2} \hat{A}^{-1/2} \hat{a} \quad (\text{A4})$$

where T is the sample size, and

$$\begin{aligned} \hat{A} &= \frac{2}{T} \left(\sum_{j=1}^* \psi(\lambda_j)^2 - \sum_{j=1}^* \psi(\lambda_j) \hat{\varepsilon}(\lambda_j)' \times \left(\sum_{j=1}^* \hat{\varepsilon}(\lambda_j) \hat{\varepsilon}(\lambda_j)' \right)^{-1} \times \sum_{j=1}^* \hat{\varepsilon}(\lambda_j) \psi(\lambda_j) \right) \\ \hat{a} &= \frac{-2\pi}{T} \sum_{j=1}^* \psi(\lambda_j) g(\lambda_j; \hat{\tau})^{-1} I(\lambda_j); \quad \hat{\sigma}^2 = \sigma^2(\hat{\tau}) = \frac{2\pi}{T} \sum_{j=1}^{T-1} g(\lambda_j; \hat{\tau})^{-1} I(\lambda_j); \end{aligned}$$

$$\hat{\varepsilon}(\lambda_j) = \frac{\partial}{\partial \tau} \log g(\lambda_j; \hat{\tau}); \quad \lambda_j = \frac{2\pi j}{T}; \quad \hat{\tau} = \arg \min_{\tau \in T^*} \sigma^2(\tau),$$

and the sums over $*$ in the above expressions are over $\lambda \in M$ where $M = \{\lambda: -\pi < \lambda < \pi, \lambda \notin (\rho_l - \lambda_l, \rho_l + \lambda_l), l = 1, 2, \dots, s\}$ such that $\rho_l, l = 1, 2, \dots, s < \infty$ are the distinct poles of $\psi(\lambda)$ on $(-\pi, \pi]$. Also,

$$\psi(\lambda_j) = \text{Re} \left[\log \left(\frac{\partial}{\partial \theta} \log \rho(e^{i\lambda_j}; \theta) \right) \right]_{\theta=0}, \quad (\text{A5})$$

and $I(\lambda_j)$ is the periodogram of u_t evaluated under the null. The function g above is a known function coming from the spectral density of u_t ,

$$f(\lambda; \sigma^2; \tau) = \frac{\sigma^2}{2\pi} g(\lambda; \tau), \quad -\pi < \lambda \leq \pi.$$

Note that these tests are purely parametric, and, therefore, they require specific modelling assumptions about the short memory specification of u_t . Thus, if u_t is a white noise, then $g \equiv 1$, (and thus, $\hat{\varepsilon}(\lambda_j) = 0$), and if it is an AR process of the form $\phi(L)u_t = \varepsilon_t$, $g = |\phi(e^{i\lambda})|^{-2}$, with $\sigma^2 = V(\varepsilon_t)$, so that the AR coefficients are a function of τ .

Based on H_0 (A3), Robinson (1994b) showed that under certain very mild regularity conditions:

$$\hat{r} \rightarrow_d N(0,1) \quad \text{as } T \rightarrow \infty.$$

Hence, we are in a classical large sample-testing situation: an approximate one-sided $100\alpha\%$ level test of H_0 (A4) against the alternative: $H_a: d > d_0$ ($d < d_0$) will be given by the rule: “Reject H_0 if $\hat{r} > z_\alpha$ ($\hat{r} < -z_\alpha$)”, where the probability that a standard normal variate exceeds z_α is α .

Note that given the functional form of ρ in (A2),

$$\begin{aligned}\psi(\lambda_j) &= \operatorname{Re} \left[\left(\frac{\partial}{\partial \theta} \log \rho(e^{i\lambda_j}; \theta) \right) \right]_{\theta=0} = \operatorname{Re} \left[\frac{\partial}{\partial \theta} (d + \theta) \log(1 - e^{i\lambda_j k}) \right] = \\ &\operatorname{Re} \left[\log(1 - e^{i\lambda_j k}) \right] = \operatorname{Re} [\log(1 - \cos \lambda k - i \sin \lambda k)] = \log |(1 - \cos \lambda k - i \sin \lambda k)| = \\ &\log(2 - 2 \cos \lambda k)^{0.5}.\end{aligned}$$

In some simple cases, the above formula simplifies. Thus, for example, if $k = 1$, which is the form employed in Sections 4.1 and 4.2,

$$\psi(\lambda_j) = \log \left| 2 \sin \frac{\lambda_j}{2} \right|.$$

If $k = 2$, and noting that $(1 - e^{i2\lambda}) = (1 - e^{i\lambda})(1 - e^{-i\lambda})$,

$$\psi(\lambda_j) = \log \left| 2 \sin \frac{\lambda_j}{2} \right| + \log \left(2 \cos \frac{\lambda_j}{2} \right)$$

and similarly, if $k = 4$,

$$\psi(\lambda_j) = \log \left| 2 \sin \frac{\lambda_j}{2} \right| + \log \left(2 \cos \frac{\lambda_j}{2} \right) + \log |2 \cos \lambda_j|.$$

A common feature of all these expressions is that they have a finite number (k) of poles across the spectrum, but they are all squared integrable.

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Figure 1: Original time series (in logarithm form)

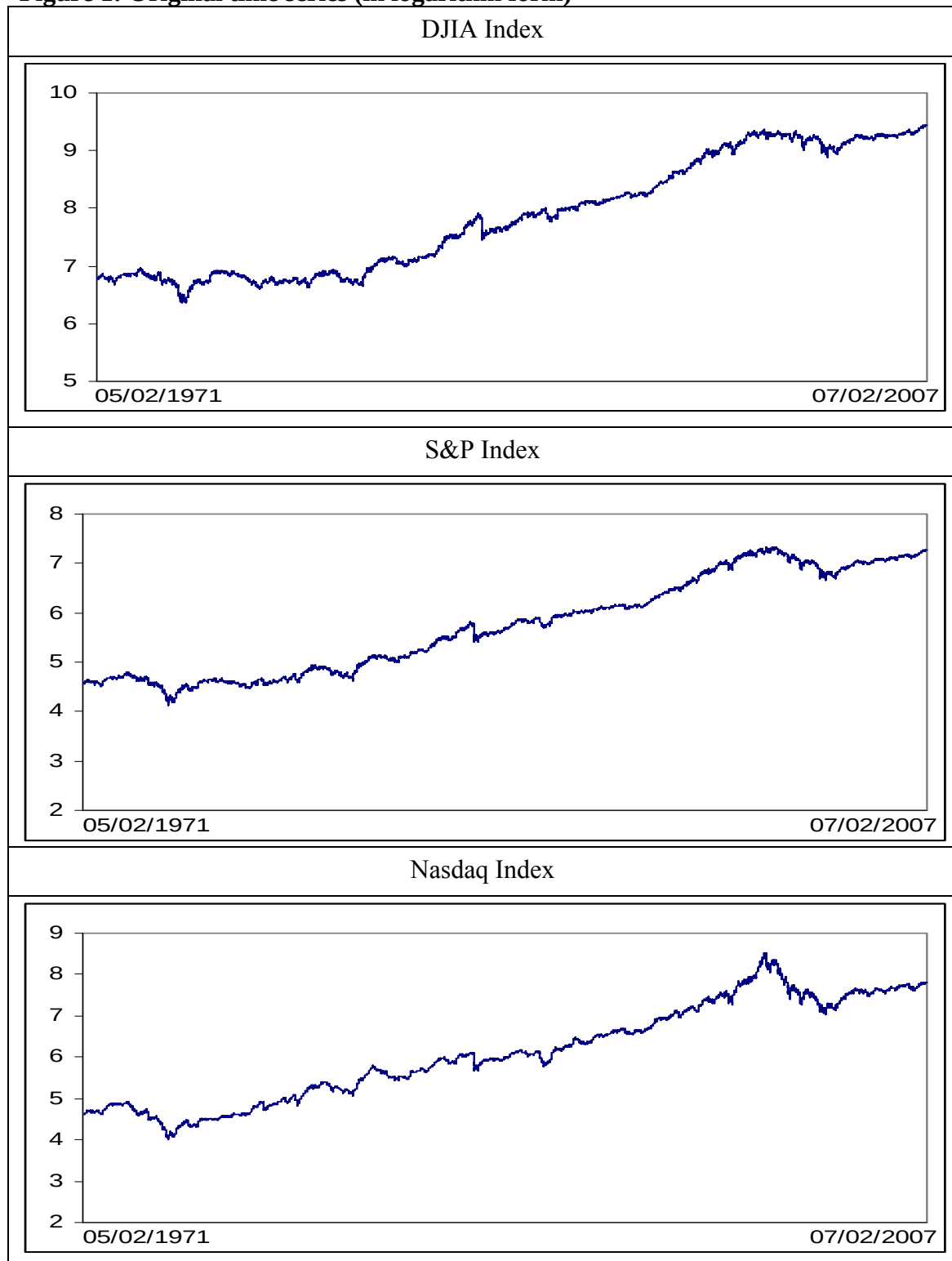


Table 1: Summary statistics on US stock market indices

	M	SD	Skew	Kurtosis	Jarque-Bera
Original series					
DJIA	7.8566	0.9885	0.24	1.51	958
S&P	5.7481	0.9807	0.16	1.54	864
NASDAQ	6.1176	1.1697	0.10	1.75	623
First differences (returns)					
DJIA	0.000284	0.0100	-1.83	53	98480
S&P	0.000288	0.0098	-1.43	38.9	510419
NASDAQ	0.00034	0.0116	-0.31	14.3	50144

Table 2: Estimates of d based on white noise disturbances

Series	No regressors (NR)	An intercept (I)	A linear trend (LT)
DJIA	1.001 [0.988, 1.014]	1.010 [0.996, 1.025]	1.010 [0.996, 1.024]
S&P	1.001 [0.989, 1.015]	1.014 [1.000, 1.029]	1.014 [1.000, 1.029]
NASDAQ	1.006 [0.993, 1.019]	1.067 [1.054, 1.081]	1.067 [1.054, 1.081]

The values in brackets refers to the 95% confidence intervals.

Table 3: Values of the test statistic (Robinson, 1994b) with AR(1) disturbances

d	DJIA			S&P			NASDAQ		
	NR	I	LT	NR	I	LT	NR	I	LT
0.00	-0.475	-0.474	-0.618	-0.315	-0.315	0.625	0.387	0.387	6.371
0.10	-17.675	-17.676	-16.708	-17.618	-17.631	-16.180	-15.317	-15.407	-15.234
0.20	-21.599	-21.152	-21.903	-21.599	-21.618	-21.749	-21.264	-21.255	-21.221
0.30	-26.123	-26.342	-27.421	-26.750	-26.993	-27.458	-26.697	-26.819	-26.897
0.40	-32.936	-33.935	-33.964	-33.769	-35.092	-33.565	-33.849	-34.724	-32.435
0.50	-42.787	-48.140	-33.910	-44.097	-48.012	-28.395	-44.368	-44.128	-27.176
0.60	-57.736	-11.500	29.698	-59.963	13.003	46.215	-60.808	5.380	24.514
0.70	-78.094	35.333	69.657	-83.578	37.456	58.182	-85.318	41.293	52.186
0.80	-104.51	19.552	28.490	-105.39	18.191	22.911	-100.07	27.285	29.951
0.90	-93.079	4.999	6.334	-63.853	4.847	5.489	-53.363	12.530	12.899
1.00	-1.491	-3.584	-3.584	-1.681	-3.494	-3.494	-2.533	1.989	1.989
1.10	19.673	-9.284	-9.280	15.125	-9.135	-9.139	12.739	-5.476	-5.476
1.20	12.173	-13.618	-13.612	10.895	-13.377	-13.370	9.214	-10.964	-10.959
1.30	4.534	-17.049	-17.052	3.511	-16.733	-16.733	2.969	-15.130	-15.130
1.40	-1.803	-19.865	-19.844	-2.466	-19.491	-19.482	-2.789	-18.407	-18.401
1.50	-7.036	-22.225	-22.201	-7.506	-21.820	-21.811	-7.720	-21.071	-21.064

In bold, significant statistics at the 5% level. NR = No deterministic terms; I = Intercept. LT = Linear trend.

Table 4: Estimates of d based on AR(1) disturbances

Series	No regressors (NR)	An intercept (I)	A linear trend (LT)
DJIA	1.370 [1.344, 1.397]	0.953 [0.935, 0.973]	0.957 [0.940, 0.975]
S&P	1.357 [1.330, 1.385]	0.953 [0.934, 0.974]	0.957 [0.940, 0.975]
NASDAQ	1.350 [1.322, 1.378]	1.024 [1.004, 1.044]	1.024 [1.004, 1.044]

The values in brackets refers to the 95% confidence intervals.

Figure 2: Estimates of d based on 1-year samples with white noise disturbances

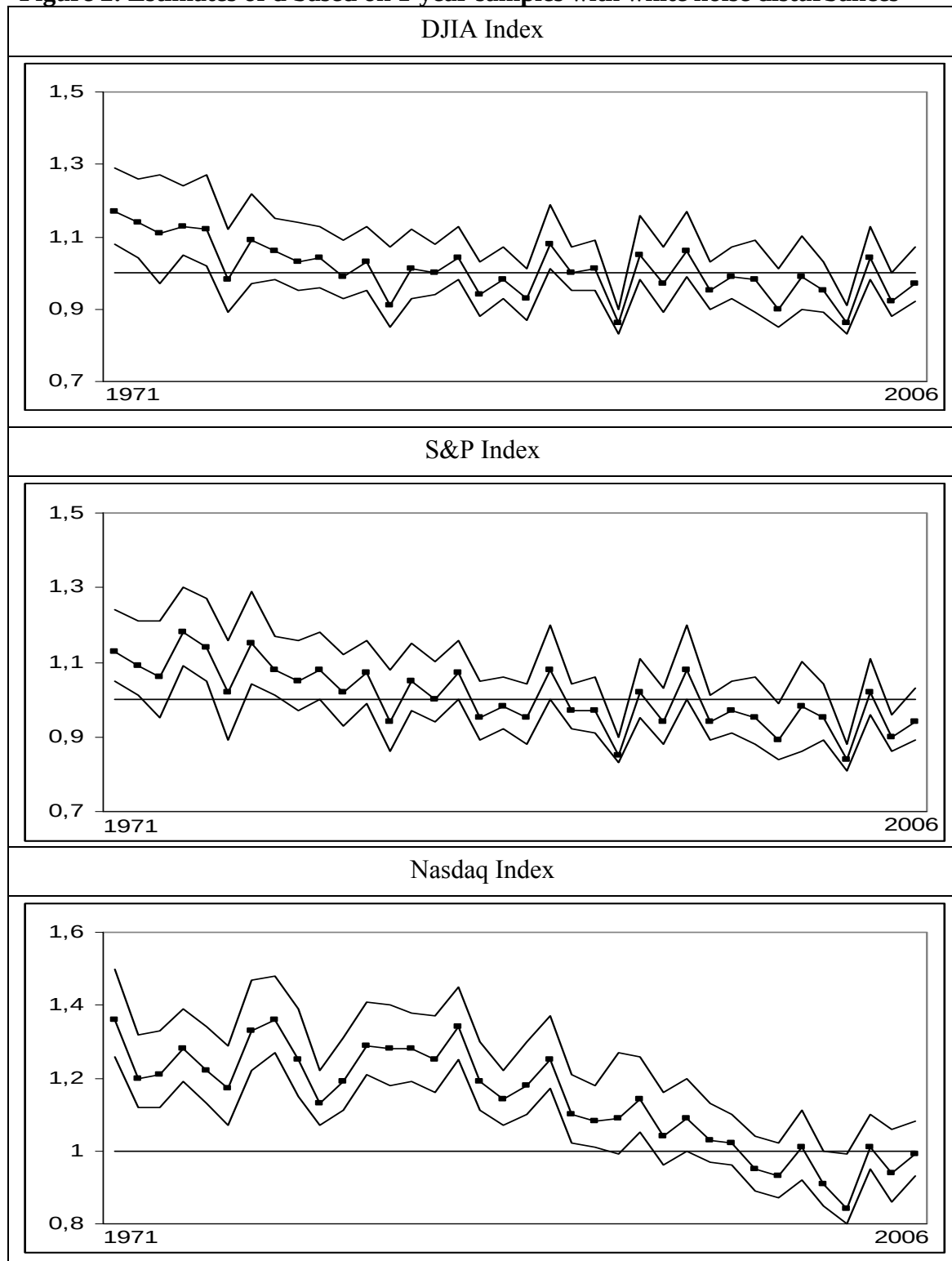


Figure 3: Estimates of d based on 1-year samples with AR(1) disturbances

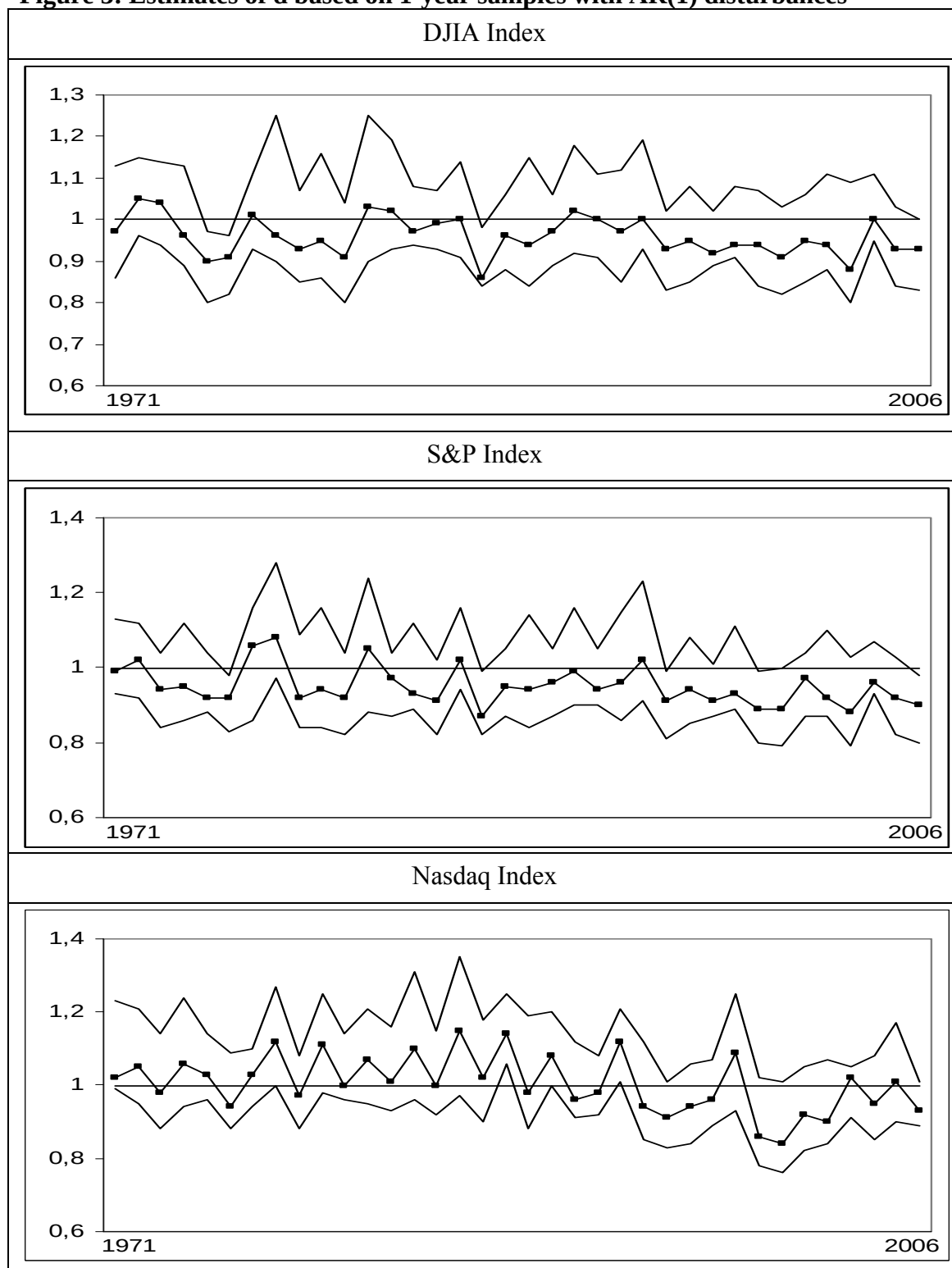


Table 5: Estimates of d for the volatility processes based on white noise disturbances

a) Squared returns			
Series	No regressors (NR)	An intercept (I)	A linear trend (LT)
DJIA	0.111 [0.100, 0.123]	0.111 [0.100, 0.213]	0.111 [0.100, 0.123]
S&P 500	0.126 [0.115, 0.138]	0.126 [0.115, 0.138]	0.126 [0.115, 0.138]
NASDAQ	0.218 [0.209, 0.227]	0.218 [0.209, 0.227]	0.216 [0.207, 0.226]
a) Absolute returns			
Series	No regressors (NR)	An intercept (I)	A linear trend (LT)
DJIA	0.164 [0.156, 0.173]	0.164 [0.156, 0.173]	0.161 [0.153, 0.169]
S&P	0.168 [0.160, 0.176]	0.168 [0.160, 0.176]	0.166 [0.158, 0.174]
NASDAQ	0.210 [0.203, 0.217]	0.210 [0.203, 0.217]	0.208 [0.201, 0.215]

The values in brackets refers to the 95% confidence intervals.

Fig. 4: Estimates of d in volatility based on 1-year samples with white noise disturbances

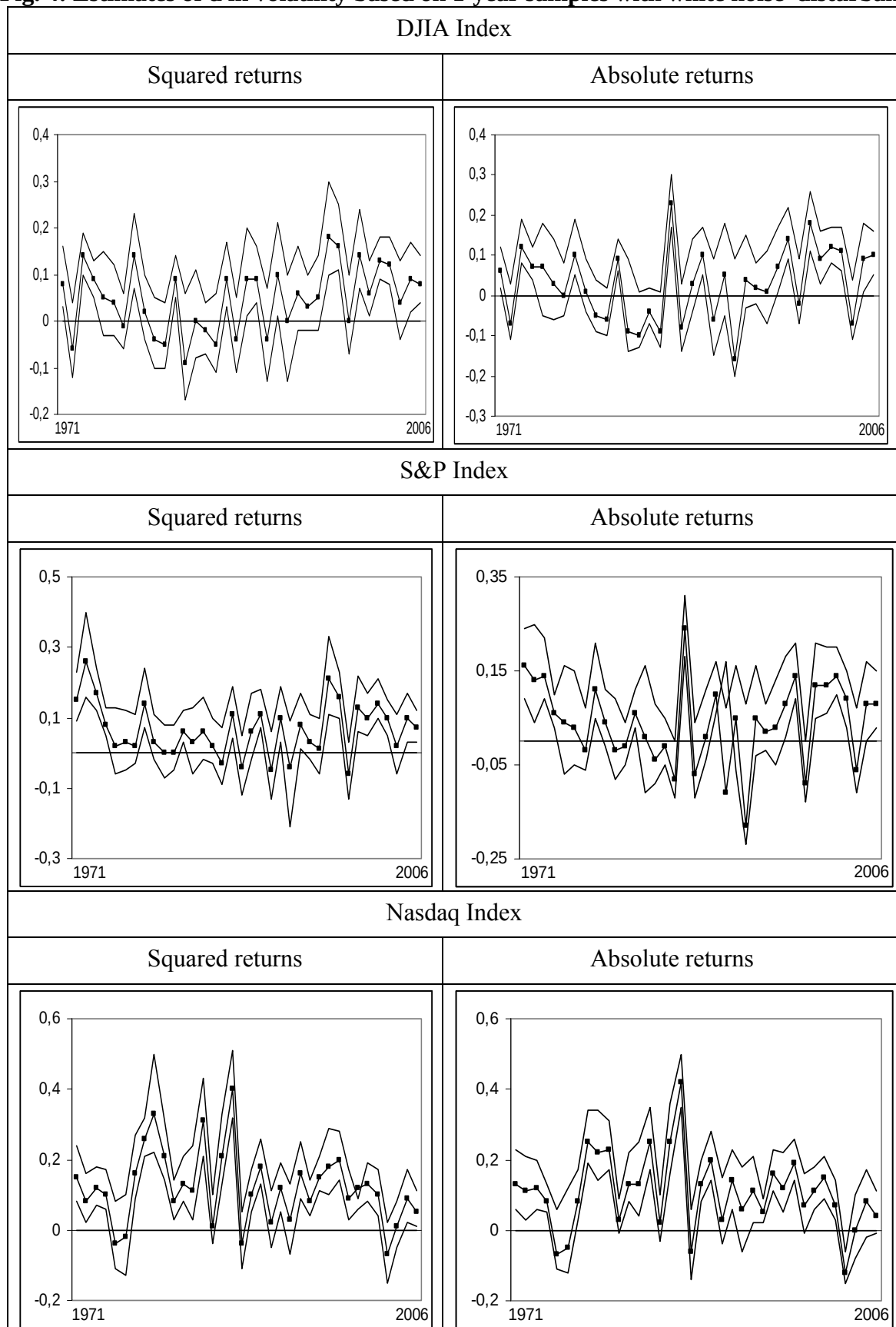


Table 6: Average values of d in the volatility processes across years

Series	Squared returns	Absolute returns
DJIA	[0.035 (0.057) 0.150]	[0.035 (0.080) 0.140]
S&P	[0.060 (0.110) 0.175]	[0.060 (0.120) 0.195]
NASDAQ	[0.045 (0.100) 0.175]	[0.025 (0.085) 0.170]

Table 7: Estimates of d for each day of the week based on white noise disturbances

	Monday	Tuesday	Wednesday	Thursday	Friday
DJIA	0.939 [0.915, 0.967]	0.959 [0.934, 0.987]	0.991 [0.964, 1.022]	0.991 [0.964, 1.021]	0.978 [0.953, 1.007]
S&P	0.940 [0.915, 0.966]	0.954 [0.930, 0.982]	0.983 [0.957, 1.013]	0.987 [0.960, 1.016]	0.981 [0.956, 1.010]
NASDAQ	1.033 [1.007, 1.063]	1.046 [1.018, 1.076]	1.065 [1.037, 1.096]	1.079 [1.050, 1.111]	1.083 [1.054, 1.115]

Table 8: Estimates of d for each day of the week based on AR(1) disturbances

	Monday	Tuesday	Wednesday	Thursday	Friday
DJIA	0.955 [0.918, 0.998]	0.979 [0.938, 1.025]	0.989 [0.945, 1.040]	1.000 [0.956, 1.053]	1.022 [0.976, 1.076]
S&P	0.967 [0.930, 1.009]	0.988 [0.949, 1.033]	1.000 [0.958, 1.050]	1.003 [0.961, 1.054]	1.022 [0.978, 1.074]
NASDAQ	1.080 [1.033, 1.133]	1.067 [1.020, 1.118]	1.084 [1.034, 1.140]	1.100 [1.044, 1.158]	1.079 [1.027, 1.137]

Table 9: Values of the test statistic for the “day of the week” with white noise disturbances

d	DJIA			S&P			NASDAQ		
	NR	I	LT	NR	I	LT	NR	I	LT
0.00	409.32	409.32	333.80	409.22	419.21	325.87	406.34	406.34	307.85
0.10	386.74	386.78	322.49	408.95	409.73	322.77	404.26	404.03	312.56
0.20	281.78	283.39	314.65	302.55	305.73	318.31	307.64	310.09	303.34
0.30	244.11	242.61	309.70	237.44	239.99	304.47	232.85	237.11	291.76
0.40	206.06	200.00	304.64	197.74	195.86	297.38	193.58	194.91	283.30
0.50	159.82	151.29	301.89	154.24	153.13	285.49	152.71	159.75	265.07
0.60	111.31	112.56	256.46	108.51	120.46	225.46	108.99	135.10	205.37
0.70	68.81	75.58	142.87	67.74	77.47	118.93	69.09	96.31	120.44
0.80	36.73	35.56	51.02	36.44	35.83	44.65	37.81	55.48	60.42
0.90	14.61	10.71	12.52	14.55	10.60	11.61	15.63	26.83	27.40
1.00	-0.23	-3.93	-3.93	-0.22	-4.29	-4.29	0.52	7.87	7.86
1.10	-10.33	-13.31	-13.36	-10.32	-13.79	-13.82	-9.81	-4.94	-4.95
1.20	-17.43	-19.82	-19.83	-17.43	-20.32	-20.32	-17.09	-13.88	-13.88
1.30	-22.63	-24.60	-24.59	-22.62	-25.06	-25.06	-22.39	-20.32	-20.31
1.40	-26.56	-28.23	-28.23	-26.55	-28.64	-28.64	-26.39	-25.08	-25.08
1.50	-29.63	-31.07	-31.07	-29.61	-31.43	-31.43	-29.50	-28.70	-28.70

Table 10: Estimates of d for the “day of the week” based on white noise disturbances

Series	No regressors (NR)	An intercept (I)	A linear trend (LT)
DJIA	0.998 [0.985, 1.011]	0.968 [0.955, 0.982]	0.969 [0.958, 0.983]
S&P	0.998 [0.985, 1.011]	0.965 [0.957, 0.978]	0.967 [0.953, 0.979]
NASDAQ	1.004 [0.992, 1.018]	1.057 [1.048, 1.072]	1.058 [1.044, 1.070]

The values in brackets refers to the 95% confidence intervals.

Table 11: Values of the test statistic for the “day of the week” with AR(1) disturbances

d	DJIA			S&P			NASDAQ		
	NR	I	LT	NR	I	LT	NR	I	LT
0.00	-0.40	-0.41	-0.38	-0.39	-0.39	-0.35	-0.42	-0.42	-0.37
0.10	-7.68	-7.68	-8.16	-7.08	-7.09	-8.24	-6.90	-8.92	-8.38
0.20	-16.54	-16.52	-15.92	-16.20	-16.17	-16.02	-16.12	-16.09	-16.18
0.30	-22.94	-22.93	-22.12	-22.89	-22.86	-22.18	-22.94	-22.91	-22.34
0.40	-27.61	-27.63	-26.71	-27.62	-27.60	-26.68	-27.68	-27.65	-26.82
0.50	-31.25	-31.11	-29.47	-31.22	-30.79	-28.97	-31.25	-30.75	-29.23
0.60	-34.03	-29.53	-28.56	-33.94	-30.42	-27.59	-33.93	-30.77	-28.99
0.70	-36.02	-31.28	-26.98	-35.90	-31.11	-27.40	-35.87	-30.13	-29.36
0.80	-37.45	-33.72	-30.32	-37.31	-31.31	-30.76	-37.26	-31.64	-31.42
0.90	-38.52	-34.44	-33.57	-38.38	-33.79	-33.71	-38.33	-33.61	-33.58
1.00	-39.38	-36.65	-35.70	-39.25	-35.80	-35.80	-39.21	-35.45	-35.45
1.10	-40.12	-38.01	-37.30	-40.01	-37.42	-37.42	-39.97	-37.07	-37.07
1.20	-40.78	-38.63	-38.63	-40.69	-38.88	-38.75	-40.66	-38.46	-38.46
1.30	-41.39	-39.75	-39.75	-41.32	-39.87	-39.87	-41.29	-39.64	-39.64
1.40	-41.95	-40.72	-40.72	-41.89	-40.07	-40.83	-41.88	-40.65	-40.65
1.50	-42.48	-42.99	-41.55	-42.43	-41.66	-41.66	-42.42	-41.53	-41.53

Table 12: Values of the test statistic for the “day of the week” with Bloomfield disturbances

d	DJIA			S&P			NASDAQ		
	NR	I	LT	NR	I	LT	NR	I	LT
0.00	184.28	184.28	147.48	185.22	185.22	143.18	187.11	187.11	137.94
0.10	166.82	166.90	146.37	181.52	180.20	143.83	179.29	180.98	135.72
0.20	115.62	115.93	135.92	125.88	125.48	133.02	126.87	129.33	124.68
0.30	93.34	92.54	122.95	89.92	92.15	121.82	89.13	89.38	115.29
0.40	72.16	69.84	116.79	68.56	66.97	113.52	67.70	66.65	108.69
0.50	48.26	44.75	113.22	45.97	45.39	104.87	46.37	48.87	97.51
0.60	26.41	25.96	87.95	24.83	29.63	74.88	24.70	35.42	65.70
0.70	7.40	9.20	36.66	6.90	9.68	27.15	7.14	17.10	27.38
0.80	-6.80	-8.38	-2.29	-6.88	-8.07	-4.84	-6.62	-0.93	1.29
0.90	-16.78	-18.69	-17.94	-16.75	-18.96	-18.54	-16.58	-12.96	-12.72
1.00	-23.50	-25.16	-25.17	-23.45	-25.45	-25.45	-23.07	-21.16	-21.16
1.10	-27.98	-29.50	-29.51	-27.94	-29.59	-29.60	-27.92	-26.75	-26.75
1.20	-31.34	-32.52	-32.52	-31.30	-32.79	-32.79	-31.32	-30.80	-30.79
1.30	-33.95	-34.90	-34.90	-33.92	-35.02	-35.02	-33.81	-33.62	-33.69
1.40	-35.83	-36.74	-36.74	-35.80	-36.85	-36.85	-35.84	-35.84	-35.84
1.50	-37.43	-38.21	-38.21	-37.40	-38.30	-38.30	-37.45	-37.63	-37.63

Table 13: Estimates of d for the “day of the week” based on Bloomfield disturbances

Series	No regressors (NR)	An intercept (I)	A linear trend (LT)
DJIA	0.747 [0.734, 0.753]	0.750 [0.738, 0.758]	0.790 [0.782, 0.796]
S&P	0.748 [0.736, 0.754]	0.748 [0.742, 0.756]	0.779 [0.773, 0.783]
NASDAQ	0.748 [0.738, 0.761]	0.796 [0.787, 0.804]	0.805 [0.797, 0.817]

The values in brackets refers to the 95% confidence intervals.

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