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CAUSALITIES AND CASUALTIES:
MEDIA ATTENTION AND TERRORISM,
1970–2010

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Causalities and casualties: Media attention and terrorism, 1970–2010

CHRISTOPH P. PFEIFFER

Zusammenfassung/ Abstract

The results of empirical research on whether media attention encourages terrorism differ according to the study period. The present contribution identifies which episodes from 1970 to 2010 are characterized by Granger-causalities. Structural breaks are systematically taken into account, and both domestic and transnational terrorism are considered. Data on terrorism are drawn from the *Global Terrorism Database* (GTD), while the media attention variable is constructed by the number of *New York Times*' (NYT) articles that contain the word 'terrorism'. A bidirectional Granger-causality between transnational terrorism and media attention was found only from 1999:09 to 2002:07, indicating a temporary effect of 9/11.

JEL-Klassifikation / JEL-Classification: C32, D74

Schlagworte / Keywords: terrorism, media, Bai-Perron test, GTD, Toda-Yamamoto procedure, NYT

Sheikh Usama knows that the media war is not less important than the military war against America. That's why Al Qaeda has many media wars. (Al Qaeda spokesman, cited in Gendron, 2007)

1 Introduction

Terrorism can be described as a communication strategy that employs violence to spread a political message (Waldmann, 2005). One important aspect of media coverage on terrorism is that it may encourage further terrorism. While Wardlaw (1989) argued that 'there is no clear evidence that publicity is responsible for significantly affecting the occurrence of terrorism' (p. 78), most scholars seem to agree that the media and terrorists live in a symbiotic relationship (Hoffman, 2006).

However, few studies address this question quantitatively. On the one hand, Nelson and Scott (1992) found that media attention did not Granger-cause terrorism for the 1968–1984 period. In a later contribution, Scott (2001) explained the absence of a causal relationship between terrorism and media attention with a congestion-effect that emerges from the public's preference for variety and the media's constraints regarding the amount of news that can be covered. With more terrorist attacks the congestion effect increases, and the probability of an attack being covered in the news declines. On the other hand, Frey and Rohner (2007) found a positive bidirectional Granger-causality between terrorism and media attention for the period January 1998 to June 2005 which implies a vicious circle of 'blood and ink'. Frey and Rohner (2007) argued that the difference between Nelson and Scott's and their own findings was likely due to the different study period, globalization and 09/11 having induced a new equilibrium of increased media attention and increased terrorism.

The present paper contributes to the findings in this area by systematically identifying which episodes between 1970 and 2010 are characterized by Granger-causalities between terrorism and media attention. The data on terrorism were extracted from the *Global Terrorism Database* (GTD). The GTD contains data on domestic and transnational terrorism, which are disentangled using the method devised by Enders et al. (2011). Contrary to previous studies that focused on the media and either transnational or domestic terrorism, both kinds of terrorism are taken into account, since the impact of domestic and transnational terrorism on the media and vice versa is likely to be different. The media attention variable was constructed by counting the number of articles published by *The New York Times* (NYT) that contain the word 'terrorism'. The NYT is commonly held the most influential newspaper in the USA. Further, the article count matches other measures of public concern about terrorism well (see Section 3).

The remainder of this paper proceeds as follows: after a brief review of the literature, a theoretical framework is presented in Section 3. Section 4 describes the relevant data. In Section 5, the empirical procedure is outlined, and results are discussed. Section 6 summarizes the findings and draws some conclusions.

2 Literature review

In addition to the loss of lives and the destruction caused by terrorism, there are further, less tangible consequences that mainly manifest through the distortion of individual choices. For instance, many individuals substituted safe air traffic by relatively dangerous road traffic after the terrorist attacks on September 11, 2001 (henceforth 9/11), which is estimated to have led to a significant increase in the number of traffic related accidents (Gigerenzer, 2004). Terrorism has been shown to negatively affect other activities such as international trade (e.g. Nitsch & Schumacher, 2004), tourism (e.g. Enders et al., 1992; Drakos & Kutan, 2003; Pizam & Smith, 2000) or net foreign direct investment (Enders & Sandler, 1996). It seems reasonable to assume that most of the indirect costs of terrorism can be attributed to its portrayal in the media (Melnick & Eldor, 2010).

A starting point for an empirical analysis on terrorism and the media is to note that the media actually cover very few terrorist incidents. Weimann and Winn (1994) estimated that newspapers cover only one-third of all transnational attacks. For TV news, the proportion falls to approximately one-sixth. This raises the question which characteristics of a terrorist attack attract media attention. There is consensus that those attacks tend to receive more attention that are successful, involve suicide, have a high number of victims, are aimed at Western countries, and include hostage taking or ransom demands (Delli Carpini & Williams, 1987; Weimann & Winn, 1994; Chermak & Gruenewald, 2006).

Nacos (2007) argued that terrorist organizations have four media-centered goals: (i) to create awareness of their organization, (ii) to spread recognition of the organization's motives, (iii) to gain sympathy of those in whose interest they claim to act, (iv) and to achieve a quasi-legitimate status. Further, Schbley (2004) found that US and European media entities were important transmitters of attack ideas to potential terrorists. Slone (2000) demonstrated that the consumption of news on terrorism (in this case, TV news) induces more anxiety than news that is unrelated to situations of national danger. Similarly, Pfeiffer et al. (2004) showed that media consumption tends to increase the perceived crime rate. Although

empirical evidence is lacking, it seems likely that a similar relationship applies to terrorism (Pfeiffer, 2012b).¹

Melnick and Eldor (2010) analyzed all 3045 articles on terrorism published in the four major Israeli newspapers (*Yedioth Aharonoth*, *Maariv*, *Haaretz* and *Globes*) in 2002. They found that a terrorist attack leads, on average, to 2.7 front page articles, 14.2 pictures and 5.7 large headlines. The more attention a terrorist attack receives, the further the decline in the *Tel Aviv Stock Exchange* (TASE) relative to *Standard & Poor's 500* Index. Calculating the opportunity costs of reporting about terrorism, they found that terrorists are provided with media coverage free of charge that is on average worth NIS 0.38 million per incident. Further, they estimated that it is not the terrorist act itself, but media coverage that causes the decline in the TASE.

Becker and Rubinstein (2011) explored how a low-probability event such as terrorism can have a considerable influence on the choices made by individuals. In contrast to the findings typically derived from a standard state-dependent utility model, they argued that low-probability events with a negative impact have a utility reducing effect during consumption when it is unclear whether or not the negative outcome will arise. Indeed, they maintained that even when it eventually becomes clear that the negative event will not occur, the utility lost during the previous period of uncertainty is irretrievable.

Most of the existing quantitative research on terrorism has focused on transnational terrorism, most likely due to the better availability of data. The GTD provides data on both domestic and transnational terrorism. However, it does not inherently distinguish between the two kinds of terrorism. Enders et al. (2011) devised a method to separate transnational from domestic terrorism in the GTD and outlined how some of the inconsistencies in the GTD coding procedure can be corrected. In their subsequent analysis, they found that domestic terrorism tends to spill-over to transnational terrorism, but not vice versa.

The analysis of breakpoints in terrorism related time series has received increasing attention in recent years. While Enders and Sandler (1993) and Yaya (2009) determined breakpoints exogenously in their respective empirical studies on transnational terrorism, data driven methods that can detect multiple structural breaks endogenously have gained in popularity. Enders and Sandler (2005) and Lee et al. (2009) employed the Bai-Perron (2003) method to endogenously detect structural breaks in different time series of terrorist attacks. Santifort et al. (2012) used a Bayesian Reversible Jump Markov chain Monte Carlo (RJCMCMC) changepoint analysis in order to detect changes in attack patterns of terrorist organizations.

¹ For an extensive review of the economics of terrorism and counterterrorism, see Schneider et al. (2010a, 2010b). Walsh (2010) provides a review on the media attention–terrorism relation.

3 Theoretical framework

The model presented in this section is drawn from Pfeiffer (2012a). Two hypotheses are set out at the end of this section.

A terrorist organization has a political objective. It derives utility from non-violent political activity O and from media attention to terrorism T :

$$U_T = \alpha(pT)^\beta + O. \quad (1)$$

$\alpha > 0$ defines the relative preference for terrorism, $0 < p < 1$ represents the estimated probability that the terrorist attacks are covered in the news, and $0 < \beta < 1$ determines the marginal utility of media attention to terrorism. With the resource constraint $O + T = 1$, the terrorist organization maximizes its utility by choosing T so that

$$T^*(p) = (p(\alpha\beta)^{1/\beta})^{\beta/(1-\beta)}. \quad (2)$$

The terrorists' reaction function is thus concavely increasing in the media's reporting propensity for $\beta < 0.5$. For $\beta > 0.5$ equation (2) increases convexly in p . The media firm's quasi linear utility function in is given by

$$U_M = \gamma(pT)^\delta + F \quad (3)$$

with $0 < \gamma < 1$ representing relative preference for news about terrorism, $0 < \delta < 1$ the marginal utility from news about terrorism, and F the news unrelated to terrorism. With the media's space restriction equal to $1 = pT + F$ optimization in p leads to

$$p^*(T) = \frac{(\gamma\delta)^{\frac{1}{1-\delta}}}{T}. \quad (4)$$

As can be seen in equation (4), the media will *ceteris paribus* react to increased terrorism with a lower reporting propensity.² For any level of T , the media will chose p so that the overall amount of news about terrorism is equal to $pT = (\gamma\delta)^{\frac{1}{1-\delta}}$. Therefore, with $T > 0$ media coverage of terrorism will change if and only if there is a change in the media's preferences. Thus, assuming constant preferences, the following two hypotheses are set out:

H1: Media coverage about terrorism does not lead to further terrorism.

H2: More terrorism does not lead to higher levels of media attention.

² Assuming common knowledge of rationality and simultaneous moves in a single interaction of the two players without the possibility of previous commitment, a Nash equilibrium can be calculated in $T = \alpha\beta(\gamma\delta)^{\beta/(1-\delta)}$; $p = \frac{(\gamma\delta)^{(1-\beta)/(1-\delta)}}{\alpha\beta}$.

4 Data

Data on terrorism were drawn from the GTD maintained by the National Consortium for the Study of Terrorism and Responses to Terrorism (START, 2011). Contrary to other databases such as the *International Terrorism: Attributes of Terrorist Events* (ITERATE) or the *National Memorial Institute of the Prevention of Terrorism* (MIPT) which provide data only on transnational terrorism, the GTD provides data on both domestic and transnational terrorism. Unfortunately, the GTD does not include a variable that distinguishes between the two kinds of events.

Since the mutual effects from media attention on transnational and domestic terrorism are likely to differ, I use the five step procedure developed by Enders et al. (2011) to separate transnational from domestic terrorism. Events are, hence, classified as transnational when

1. the nationality of a victim is different from the attack venue,
2. the target is a diplomatic entity or an NGO,
3. the attack is directed against a US entity outside the USA,
4. there are US victims, hostages or injuries or there are ransom demands to or payments from a US source when the incident is outside the USA,
5. the attack concludes in a different country relative to where it started.

Step 3 had to be omitted, because START decided to remove the necessary information from its database.³ In addition, events that cannot be definitely classified as terrorism are removed (see Enders et al., 2011). Owing to inconsistencies in the classification process, the GTD underestimates terrorism for the 1970:01–1977:06 period and overestimates terrorism for the 1991:06–1997:12 period relative the ITERATE database that used a consistent coding procedure throughout the entire period. Following Enders et al. (2011) the first period was inflated by a factor of 2.11 and the second period was deflated by a factor of 1.52. The values for the year 1993 had to be interpolated because the data were physically lost and could not be fully reconstructed (ibid.; START, 2011).⁴

A concern in empirical analyses on terrorism and particularly on terrorism and the media is that databases on terrorism have to rely on the media to obtain data. Drakos (2007) pointed out that since a “terrorist event is recorded if, and only if, it is reported by the media [...] this creates the possibility of observed recorded terrorist activity being different from actual activity to the extent that the propensity for reporting is lower than 1” (p. 909). This is particularly the case in countries that restrict press freedom. A possible solution would be inflating the data on terrorism in countries that do not enjoy full press freedom. However,

³ I would like to thank T. Sandler for clarifying this point.

⁴ All data manipulation, analyses, and visualizations were conducted in ‘R’ 2.15.2 (R Development Core Team, 2012).

even though Drakos (2007) provided an estimate of the size of the under-reporting bias, the extent of under-reporting may vary over time—especially after regime changes or following the introduction of new technologies⁵. Therefore, in order to correct for the under-reporting bias, its size would have to be quantified dynamically, which is beyond the scope of this paper.

Two time series were extracted from the database: the number of fatalities due to transnational terrorism (NFAT_trans) and the number of fatalities due to domestic terrorism (NFAT_dom). Using the number of fatalities allows capturing the intensity of terrorism better than the number of incidents. In addition, events that involve fatalities are usually recorded more accurately. For instance, ITERATE and GTD track each other much more closely for casualties than for overall incidents (Enders et al., 2011). Further, (lamentably) fatalities from terrorist attacks tend to be a more reliable predictor of media attention than the number of incidents, which would include unsuccessful attacks or hoaxes. The media attention variable was constructed by counting the number of articles in the NYT that contain the word ‘terrorism’. The NYT is generally held to be the most influential newspaper in the United States (Cision, 2011) and also enjoys high international reputation. Coverage of terrorism in the NYT thus seems to be a good proxy for a terrorist organization’s ability to gain attention internationally.

The number of articles that contain the word ‘terrorism’ is arguably a rough measure of media attention to terrorism. Details, such as whether the article appeared on the front page, involved pictures (Melnick & Eldor, 2010) or column inches (Scott, 2001) are not taken into account. Further, an article that contains the word ‘terrorism’ does not have to be directly related to a certain attack. As Hoffman et al. (2010) pointed out, a large part of the newspaper content about terrorism is related to counter-terrorism measures or other political discussions about terrorism. However, since counter-terrorism originates back to actual terrorism, it can be argued that terrorist attacks were still causal for the ensuing article. Further, the NYT article count seems to match variations in the public interest in terrorism quite well. When compared to the worldwide number of Google searches⁶ of the word ‘terrorism’, the two time

⁵ Zeitzoff (2011), for example, used data from Twitter and other social media sources to investigate the micro-dynamics of the 2008-2009 Gaza Conflict. Social media content may enable databases to lessen the degree of under-reporting in countries with restrictions on press freedom.

⁶ The number of Google searches was retrieved from Google’s Search for Insights. Using data generated by social-networks or search-engines is becoming common practice in social sciences. For example, the number of Google searches has been used to monitor the spread of influenza (Ginsberg et al., 2009) or to estimate public interest in science (Baram-Tsabari & Segev, 2009). Schietle (2012) shows that estimates of public sentiments based on the use of search engines closely correlate with the results from questionnaires. The Google variable is calculated as follows: In a first step, the absolute number of searches for a certain term within a geographical unit is counted. Since the number of searches depends on the number of people in the geographical unit, the data is normalized in a second step by dividing the number of searches by the number of overall searches within the geographical unit. This rate is then compared to all other searches within the same geographical unit. The rate

series are highly correlated (see figure 1). The number of Google searches takes into account a large and diverse population. Analyzing the number of queries in an internet search engine has the methodological advantage of measuring a populations’ actual interest rather than the editor’s choices which are not necessarily aligned (Pfeiffer, 2012b). However, the number of Google searches can only be retrieved back to 2004, whereas NYT articles are available from 1851 onwards (New York Times, 2012).

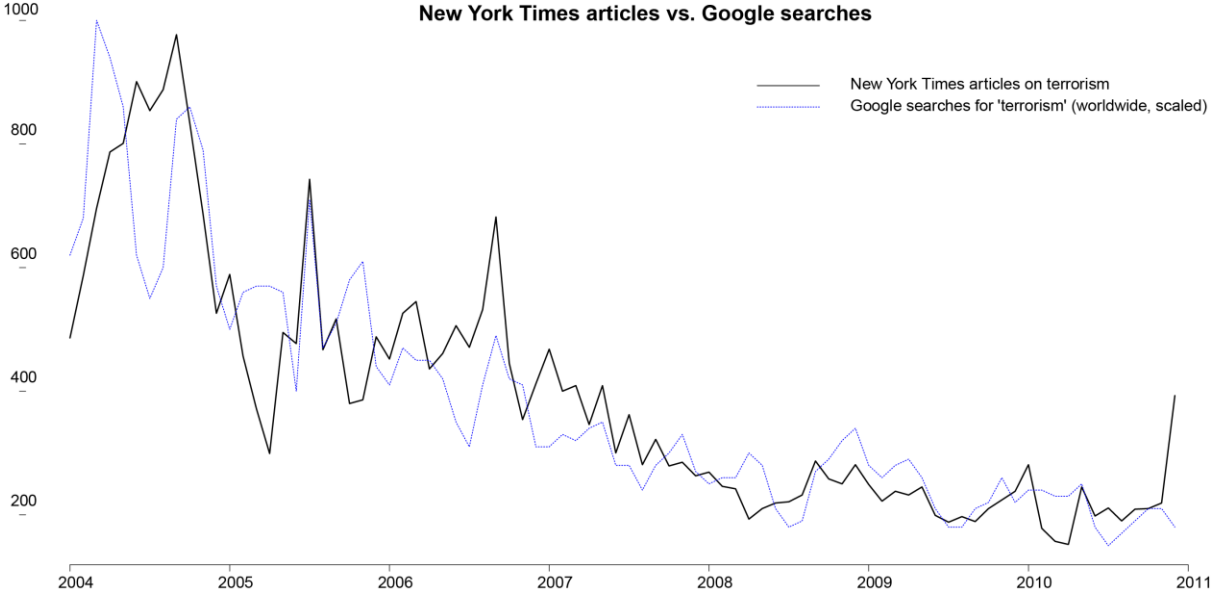


Figure 1: Number of Google searches for ‘terrorism’ compared to the number of articles in the NYT that contain the word ‘terrorism’. The Pearson correlation coefficient between the two time series is 0.853.

All three time series (NYT, NFAT_trans, NFAT_dom) were aggregated on monthly values since further aggregation would imply few observations in the segmented periods after the estimation of breakpoints. Aggregation on quarterly values was used in some instances to improve visualization and to avoid too many zero observations when a series was relatively thin.

Although quantitative evidence is lacking, there seems to be consensus that transnational terrorism is more successful in attracting media attention than domestic terrorism (Blomberg et al. 2011; Li, 2005; Arce & Sandler, 2010). Transnational terrorism involves higher risks for the terrorist organization. Blomberg et al. (2011) found that a low proportion of transnational attacks increases the survival rate of terrorist organizations.

with the highest frequency is assigned a value of 100, all other search terms are assigned their relative corresponding value (Schietle, 2012).

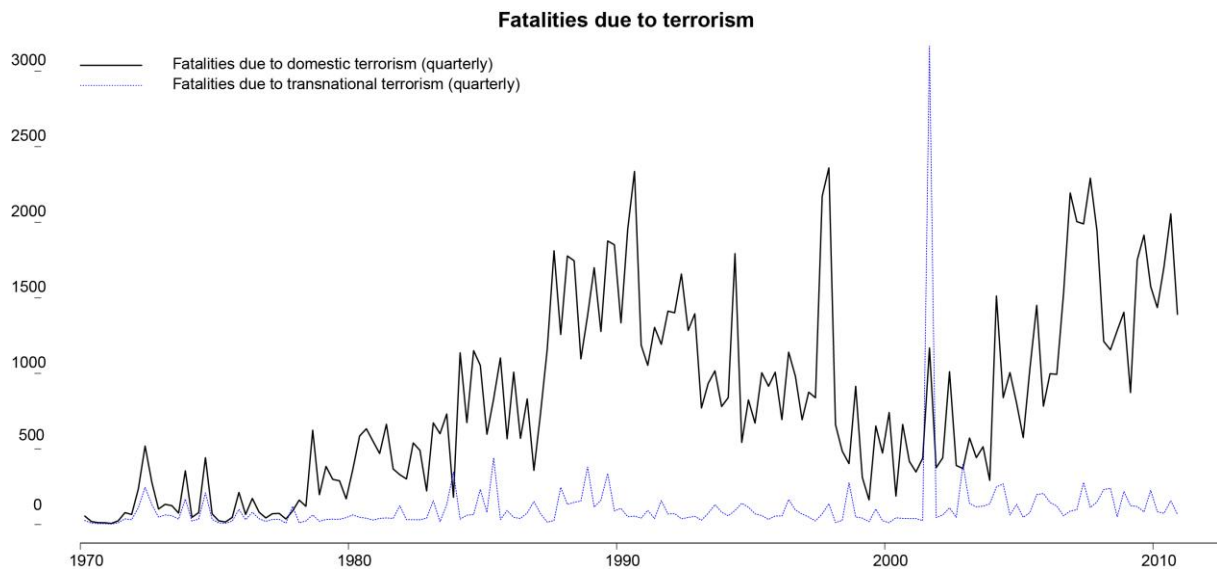


Figure 2: Domestic and transnational terrorism in the GTD after decomposition.

In the present dataset, the number of fatalities due to domestic terrorism surpasses the number of fatalities due to transnational terrorism most of the time (see figure 2). There are only 4 quarters in which the fatalities from transnational terrorism exceed those from domestic terrorism (1977:IV, 1983:IV, 2001:III, 2002:IV). While there are 211 months with more than 250 fatalities due to domestic terrorism, this number falls to only 6 when considering transnational terrorism.

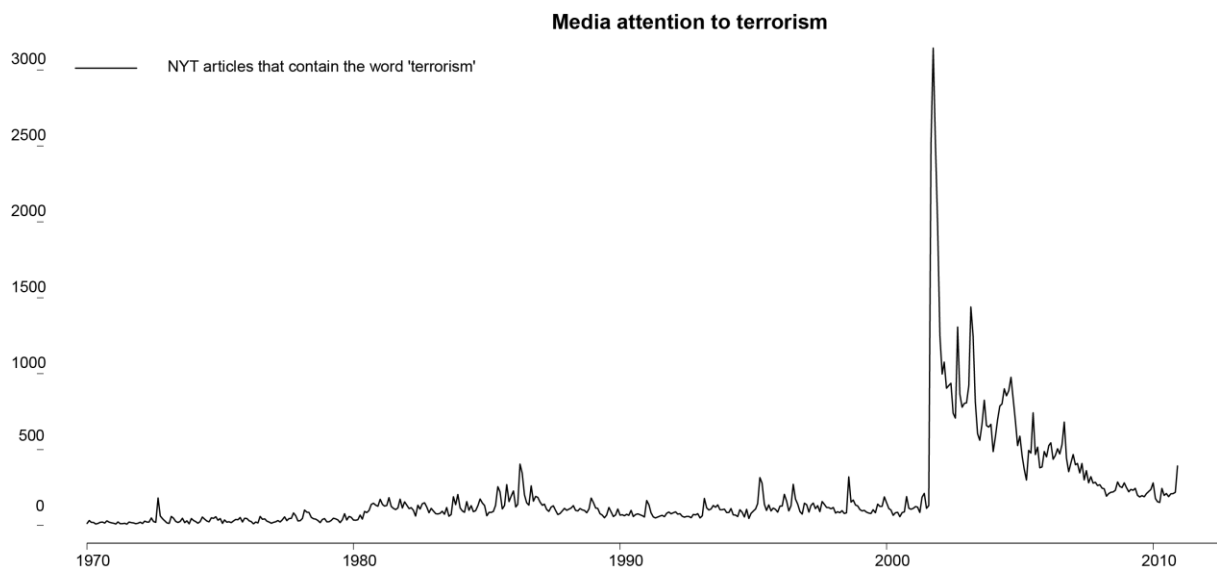


Figure 3: Media attention to terrorism measured by the number of articles in the NYT that contain the word 'terrorism'.

9/11 clearly stands out owing to the unprecedented number of victims and media attention (see figures 2 and 3); the number of casualties on this day approximately equals the total number of fatalities from transnational terrorism between 1988 and 2000 (Sandler, 2003). Similarly, the number of articles published in September and October 2001 in the *NYT* that were related to terrorism is equal to the sum of all articles from October 1997 to August 2001.

5 Methodology and results

The aim in this section is to test whether and when there is a causal relationship between terrorism and media attention. In order to test for causalities between two time series, the concept of Granger-causality has proved to be a popular method. Granger (1969) suggested that time series X_t causes time series Y_t if the predictability of Y_t improves by using both time series compared with using only the information available for Y_t . *Feedback* or a *bidirectional Granger-causality* is thus said to exist if both time series Granger-cause one another. Granger-causality is not equivalent to causality in the full Aristotelian sense, but it allows a better estimate of causation than a mere test of correlation (Geweke, 1984). However, there has been wide criticism of the use of Granger-causalities mainly due to its sensitivity to model specification (e.g. Lütkepohl, 1982; Chowdhury, 1987).

For the present study, there are three main, inter-related concerns when performing the test for Granger-causality. Firstly, testing for Granger-causality using F -statistics when one or both time series are non-stationary can lead to spurious causality (He & Maekawa, 1999). Secondly, the presence of structural breaks may lead to a false rejection of the null hypothesis of no Granger-causality (Ventosa-Santaulària & Vera-Valdés, 2008). Thirdly, the frequently used Dickey-Fuller test may be biased towards the non-rejection of the null hypothesis of a unit-root in the presence of a structural break (Perron, 1989).

Consider the following AR(n) model:

$$\Delta y_t = \alpha_0 + \gamma y_{t-1} + \sum_{k=2}^n \alpha_k y_{t-k+1} \Delta y_{t-i+1} + \varepsilon_{1t}, \quad (5)$$

where y_t represents the number of fatalities due to domestic or transnational terrorism or the number of NYT articles on terrorism in time period t . A slope component was not included as no clear evidence of a deterministic trend was found. Further, theoretical considerations in Section 2 tend to oppose a deterministic trend.⁷

Equation (5) is hence estimated beginning with a lag of $n = 48$. The lag order was reduced by 1 until the t -statistic for α_n was different from 0 at a 5% significance level. Then, the Augmented Dicky-Fuller test (ADF) was performed. In all three cases, it did not reject the null hypothesis of non-stationarity (see Appendix A: table A1), which may be due to the presence of structural breaks (Perron, 1989). Therefore, the Zivot-Andrews (ZA) (1992) test was performed on the time series. The ZA test allows for one endogenously determined break and may therefore help to ascertain whether a unit root is present or whether the rejection of the null hypothesis by the ADF test was due to the presence of a structural break. The ZA test only rejected the null hypothesis of a unit-root for NFAT_dom; therefore, both NYT and

⁷ Causality tests with a (piecewise) deterministic trend led to similar results.

NFAT_trans seem to be integrated (table A3). The Toda-Yamamoto (TY) (1995) procedure will thus be used to test for non-causality as it does not require stationarity and avoids a potential pretest bias since prior testing for cointegration is not necessary.

Because media attention, domestic terrorism and transnational terrorism might all mutually affect each other, a near-VAR model is set up. Lag lengths are allowed to differ across equations⁸:

$$d_t = \alpha_1 + \sum_{k=1}^{n_1} \alpha_{11k} m_{t-k} + \sum_{k=1}^{n_1} \alpha_{12k} r_{t-k} + \sum_{k=1}^{n_1} \alpha_{13k} d_{t-k} + \varepsilon_{1t}, \quad (6a)$$

$$r_t = \alpha_2 + \sum_{k=1}^{n_2} \alpha_{21k} m_{t-k} + \sum_{k=1}^{n_2} \alpha_{22k} r_{t-k} + \sum_{k=1}^{n_2} \alpha_{23k} d_{t-k} + \varepsilon_{2t}, \quad (6b)$$

$$m_t = \alpha_3 + \sum_{k=1}^{n_3} \alpha_{31k} m_{t-k} + \sum_{k=1}^{n_3} \alpha_{32k} r_{t-k} + \sum_{k=1}^{n_3} \alpha_{33k} d_{t-k} + \varepsilon_{3t}. \quad (6c)$$

m_t represents the number of NYT articles during time period t , r_t is the number of fatalities due to transnational terrorism, and d_t the number of fatalities due to domestic terrorism. 9/11 is a large outlier both for media attention and transnational terrorism and likely to be a structural break. However, in order to avoid imposing preconceptions and ignoring previous or subsequent structural breaks—which might hence lead to spurious causality (Ventosa-Santaulària & Vera-Valdés, 2008)—the Bai-Perron (2003) procedure was used which can endogenously detect an unknown number of breaks by minimizing the sum of the squared residuals in the potential segments.

In order to apply the Bai-Perron method to each equation, the maximum number of breaks, the minimum segment size, and the number of lags had to be specified.⁹ If the minimum segment size is specified to be too large, structural breaks may be overlooked. If it is chosen too small, there may remain too few observations in the resulting segments to draw any conclusions. Thus the maximum number of breaks was set to 6 and the minimum segment size to 48 months, which is short enough to capture periods of disturbance, but yields sufficient observations in each segment. As per the Schwarz criterion (SC), the number of lags was chosen to be 4 for equation (6a), 1 for equation (6b), and 7 for equation (6c). Residual serial correlation and parameter stability were at acceptable levels for all three estimates (see table B1 and figure B1).¹⁰ A ‘partial change model’ could have been employed in which only one coefficient is allowed to change for each regime (e.g. Enders & Sandler,

⁸ Error terms are assumed to be independent. Thus, OLS estimation is used for each equation separately.

⁹ The ‘strucchange’ package (Zeileis et al., 2003) was used to implement the Bai-Perron method.

¹⁰ The Akaike Information Criterion (AIC) was also considered, but recommended large number of lags that would have left few degrees of freedom.

2005). However, since enough observations were available for the post-9/11 period, a ‘pure structural change model’ was used in which all parameters were separately estimated for each regime.

Once the lag order was determined, the time series were tested for causality according to the TY procedure which does not require stationarity and can handle multivariate models. The TY method is based on estimating an augmented $VAR(n + d_{max})$ model where n is the lag order from the original model and d_{max} represents the maximum order integration of the time series involved. In order to test for causality, a Wald test for zero restrictions is conducted on the first n coefficients of variables in question in the augmented model. The test asymptotically follows a χ^2 -distribution with n degrees of freedom. Since first order differencing led to the rejection of the null hypothesis of a unit root in all cases (tables A2 and A4), the maximum order of integration was set as $d_{max} = I(1)$.

In line with the stability tests (Appendix B: table B1, figure B1), no breakpoints were found for the regression on the number of fatalities due to domestic terrorism. Thus, two tests were sufficient for the entire period.

Table 1: Toda-Yamamoto causality tests, endogenous variable: domestic terrorism (1970:01–2010:12).

Null hypothesis	Lag (n)	χ^2 –statistic	p -value
NYT does not cause NFAT_dom	4	0.72	0.95
NFAT_trans does not cause NFAT_dom	4	0.62	0.96

Neither media attention nor transnational terrorism was found to be causal for domestic terrorism (see table 1). An explanation for these findings could be that terrorist organizations in remote regions are not concerned about media attention in Western newspapers. However, the picture does not change much even when only fatalities from domestic terrorism in the USA, Western Europe and Eastern Europe are considered (see table 2).

Table 2: Toda-Yamamoto causality tests for domestic terrorism in the USA, Western and Eastern Europe, period: 1970:I-2010:IV. In order to avoid too many 0 observations the series were aggregated on quarterly values.

Null hypothesis	Lag (n)	χ^2 –statistic	p -value
NYT does not cause NFAT_dom	3	2.7	0.45
NFAT_trans does not cause FAT_dom	3	2.9	0.41

Note: The models resulting from different lag specifications according to the AIC and SC were tested for residual correlation and stability, lags were chosen accordingly.

It seems that domestic terrorism is unaffected by media attention. In addition, no evidence of a (reverse) spill-over effect from transnational terrorism on domestic terrorism was found.

For the number of fatalities due to transnational terrorism (equation 6b), two breakpoints were calculated, 1999:09 and 2003:09. The data were thus divided into three segments which, essentially, insulated the period around 9/11.¹¹

Table 3: Toda-Yamamoto causality tests; endogenous variable: transnational terrorism.

Regime	Null hypothesis	Lags	χ^2 -statistic	p-value	
I	1970:01–	NYT does not cause NFAT_trans	1	0.027	0.87
	1999:08	NFAT_dom does not cause NFAT_trans	1	7.3	0.0068
II	1999:09–	NYT does not cause NFAT_trans	2	22.8	1.8e-06
	2003:08	NFAT_dom does not cause NFAT_trans	2	23.0	1e-05
III	2003:09–	NYT does not cause NFAT_trans	1	0.12	0.73
	2010:12	NFAT_dom does not cause NFAT_trans	1	0.063	0.8

Note: The models resulting from the lag specifications were tested residual correlation and stability. Changes in the number of lags to the original model were made when necessary (results are reported in Appendix B).

For segments I and III, the null hypothesis that media attention does not cause transnational terrorism could not be rejected. In segment II, however, media attention to terrorism appears to be causal for the number of fatalities due to transnational terrorism as the null hypothesis of non-causality is strongly rejected. Thus the hypothesis, that media attention does not cause further terrorism (H1) was rejected for segment II, but confirmed for segments I and III. Consistent with Enders et al. (2011), domestic terrorism was found to spill over to transnational terrorism during segment I. However, the relationship seems to have disappeared in segments II and III. The results also confirm the findings of Nelson and Scott (1991) who did not find a causal relationship between media attention and terrorism from 1968–1984.

For the regression on media attention (equation 6c), two breakpoints were found: 1998:08 and 2002:08. The breakpoints for media attention and transnational terrorism are thus fairly close to each other which is in line with the stability analysis (Appendix B: figure B1). The results indicate that domestic terrorism is not causal for NYT articles on terrorism for any of the segments analyzed. In segments I and III, transnational terrorism was not found to be causal for media attention, either (see table 4). For segment II, however, the null hypothesis that transnational terrorism does not cause media attention could be rejected. Hence, similar to the results for transnational terrorism the hypothesis (H2) that terrorism does not cause media attention could be confirmed for segments I and III, but not for segment II.

¹¹ A dummy variable was introduced for 2001:09 for equations (6b) and (6c).

Table 4: Toda-Yamamoto causality tests; endogenous variable: articles on terrorism in the NYT.

Regime	Null hypothesis	Lags (<i>n</i>)	χ^2 -statistic	<i>p</i> -value	
I	1970:01–	NFAT_trans does not cause NYT	7	2.0	0.96
	1998:07	NFAT_dom does not cause NYT	7	5.7	0.57
II	1998:08–	NFAT_trans does not cause NYT	5	25.2	0.00013
	2002:07	NFAT_dom does not cause NYT	5	4.9	0.43
III	2002:08–	NFAT_trans does not cause NYT	7	10.5	0.16
	2010:12	NFAT_dom does not cause NYT	7	6.0	0.54

Note: The models resulting from the lag specifications were tested residual correlation and stability. Changes in the number of lags to the original model were made when necessary (results are reported in Appendix B).

Segment II in table 4 (1998:08–2002:07) is quite close to segment II in table 3 (1999:09–2003:08). Thus, during the overlap of the two segments (1999:09–2002:07), there is a bilateral Granger-causality between media attention and transnational terrorism. This finding is similar to the results of Frey and Rohner (2007) that identified 1998:01–2006:06 as an episode of a bilateral Granger-causality. According to the results of the present study the bilateral Granger-causality ended a few years after the events of 9/11. Therefore, it seems that 9/11 intensified the relationship between terrorism and media attention, similar to the strengthening of connections between financial markets in times of increased economic volatility (Longin & Slonik, 1995).

Apparently, 9/11 induced a temporary shift in preferences towards an increased interest in media coverage of terrorism that was large enough to be recognized by terrorists as such. The mean number of NYT articles per casualty surged from 5.5 articles/casualty (1970:I–1999:II) to 27.7 articles/casualty (1999:III–2003:III) (see figure 4). Thus, the emergence of a Granger-causality from 1999:III–2003:III can be seen as a cost-benefit reevaluation of terrorists after the payoff for terrorist attacks had increased substantially. After 2003:III, the mean ratio dropped to 11.07 articles/casualty and the causalities between transnational terrorism and media attention in both directions had disappeared. Although interest in terrorism remained at an elevated level compared to the pre-9/11 period, the level of interest in terrorism from 2001:09 until the end of 2002 was not sustained. While the amount of articles about terrorism surged from 131 in 2001:08 to 2514 in 2001:09 and even 3145 in 2001:10—which is the highest number of articles that contained the word ‘terrorism’ in the sample—the number of NYT-articles on terrorism has been steadily declining since,

despite other spectacular transnational attacks such as the Bali (2002:10), the Madrid (2004:03) or the London (2005:07) bombings.

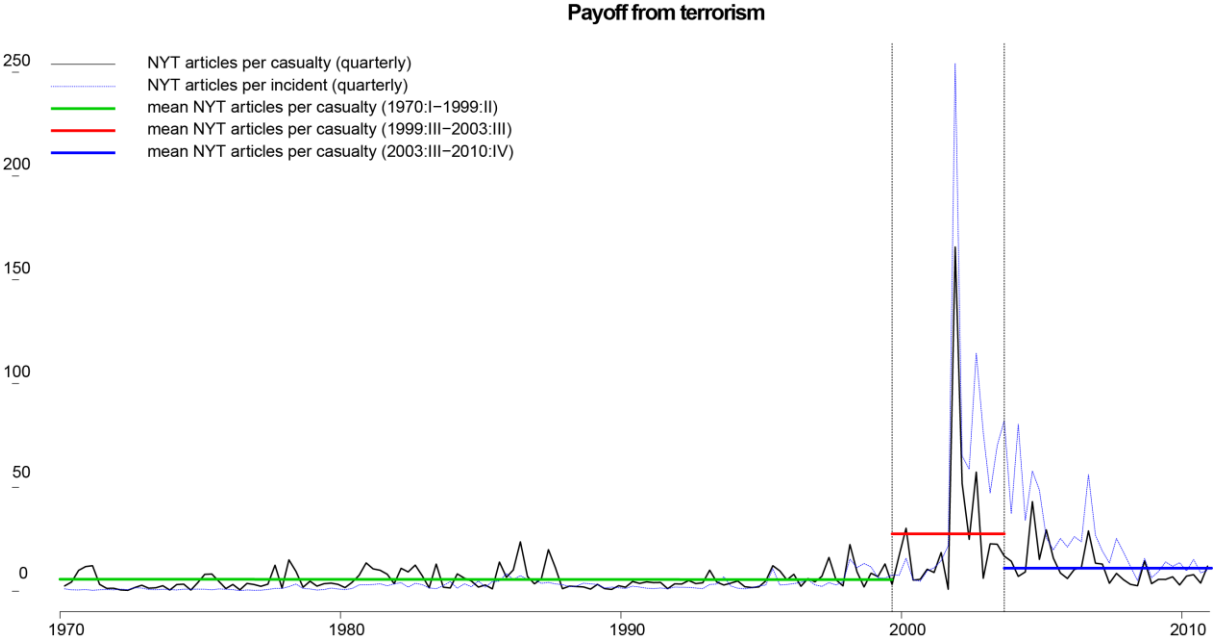


Figure 4: The media attention-terrorism is a proxy for the media’s reporting propensity on terrorism. The vertical lines indicate the segment during which a Granger-causality from media attention on terrorism was found.

The ‘war on terror’ following 9/11 led to a massive reduction in the resources available to terrorist organizations and substantially limited their ability to conduct logistically complex attacks (Enders & Sandler, 2005). Terrorist organizations thus began to increasingly rely on simple, but deadly modes of attacks such as bombing (Santifort et al., 2012) which, however, do not lead to the same levels of media attention as spectacular transnational attacks. The short circle of attention and violence from 1999:09 to 2002:07 thus seems to have been brought to an end by (i) the ‘war on terror’ and (ii) the unsustainability of extreme levels of attention to terrorism. However, while transnational terrorism almost returned to pre-9/11 level, fatalities due to domestic terrorism experienced a rise mainly due to the increases in terrorism in Iraq and Afghanistan (Enders et al., 2011).¹²

6 Conclusions

Different to previous contributions that examined causal relationships between terrorism and media attention, the present article used a relatively large study period and systematically tested for structural breaks. Further, both domestic and transnational terrorism were taken into account. In order to avoid spurious causalities, the Toda-Yamamoto procedure was used as unit root tests indicated nonstationarity for two of the three time series.

¹² Opposing to the ITERATE database, the GTD does not exclude terrorist attacks against combatants in Iraq and Afghanistan (Enders et al., 2011).

Stability tests for the regressions on transnational terrorism and media attention suggested that structural breaks occurred at roughly the same time. This was confirmed by the Bai-Perron test that estimated two breakpoints respectively to be at 1999:09 and 2003:09 for transnational terrorism and 1998:08 and 2002:08 for media attention. No structural breaks were found for the regression on domestic terrorism. Consistent with Enders et al. (2001), domestic terrorism was found to have spillover effects on transnational terrorism. However, this effect could only be confirmed for the 1970:01–1999:09 period. Domestic terrorism itself seemed to be unaffected by changes in the media coverage of terrorism and by transnational terrorism. This result held when only domestic terrorism in the USA and Europe was considered.

The segment-specific analysis indicated a Granger-causality from media attention on transnational terrorism only from 1999:09–2003:08. Similarly, a causal effect from transnational terrorism on media attention was only found from 1998:08–2002:07. Thus, according to these estimates, a bilateral Granger-causality between transnational terrorism and media attention existed between 1999:09–2002:07. This is consistent with Nelson and Scott (1991) that did not find a causal relationship between media attention and terrorism from 1968–1984. The results also confirm the findings of Frey and Rohner (2007) who identified 1998:01–2006:06 as an episode of a bilateral Granger causality.¹³ Their assertion that the reason for the difference in the findings of Nelson and Scott (1991) and their own results was mainly due to the different time period analyzed is also supported by the present analysis. However, since the bidirectional Granger-causality is closely centered around 9/11, it seems that the emergence of the bilateral Granger-causality is not caused by fundamental, globalization-induced changes such as the increased importance of the media but rather due to the singular nature of 9/11. The unprecedented amount of casualties and the ensuing enormous media attention temporarily triggered an increased interest in terrorism. It can be argued that this changed the cost-benefit relation for terrorists and hence lead to a temporary intensification of the relationship between terrorism and media attention, similar to the increases in the connection between financial markets in times of increased volatility (Longin & Slonik, 1995). Although no causal relationship between terrorism and media attention could be found during the majority of the study period, this does not imply that terrorists were not seeking media attention previous to 9/11. Rather, the result could be interpreted as rational terrorists not updating their beliefs about the media's reporting propensity after apparent changes in the level of news on terrorism.

¹³ Frey and Rohner (2007) used a different dataset (MIPT) and a slightly different media attention variable that counted the number of times the word 'terrorism' appeared in the NYT in a certain timeframe (vs. the number of articles that contain the word 'terrorism').

While previous efforts to censor news about terrorism have proven largely unsuccessful and affect the very fundamentals of democracy (Wilkinson, 1997), policymakers should keep in mind that extreme levels of media attention to terrorism may lead to a closer connection between the two. The fostering of financially independent news coverage that may have lower incentives to cover terrorism in ways that suit terrorists (Pfeiffer, 2012b) might be a promising strategy. While consumption of TV channels that financially depend on advertising has been shown to lead to an overestimation of crime rates (Pfeiffer et al., 2004), it seems likely that a similar relationship exists between media consumption and the perception of the terrorist threat. A closer analysis on the number of Google searches related to terrorism may lead to interesting results.

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Appendix

A Unit root tests

Table A1: Augmented Dickey–Fuller test against stationarity.

Null hypothesis	Lag order	Test statistic	<i>p</i> -value
NYT has a unit root	14	-2.3857	0.1689
NFAT_dom has a unit root	8	-2.635	0.08917
NFAT_trans has a unit root	13	-2.2415	0.2228

Table A2: Augmented Dickey–Fuller test against stationarity after first order differencing.

Null hypothesis	Lag order	Test statistic	<i>p</i> -value
NYT has a unit root	14	-8.4345	<0.01
NFAT_dom has a unit root	8	-11.4526	<0.01
NFAT_trans has a unit root	13	-10.3225	<0.01

Table A3: Zivot-Andrews test against stationarity.

Null hypothesis	Lag order	Test statistic	<i>p</i> -value	Potential breakpoint
NYT has a unit root	14	-4.2774	>0.1	2001:07
NFAT_dom has a unit root	8	-5.2289	>0.05	1997:12
NFAT_trans has a unit root	13	-6.348	<0.01	2001:08

Table A4: Zivot-Andrews test against stationarity after first order differencing.

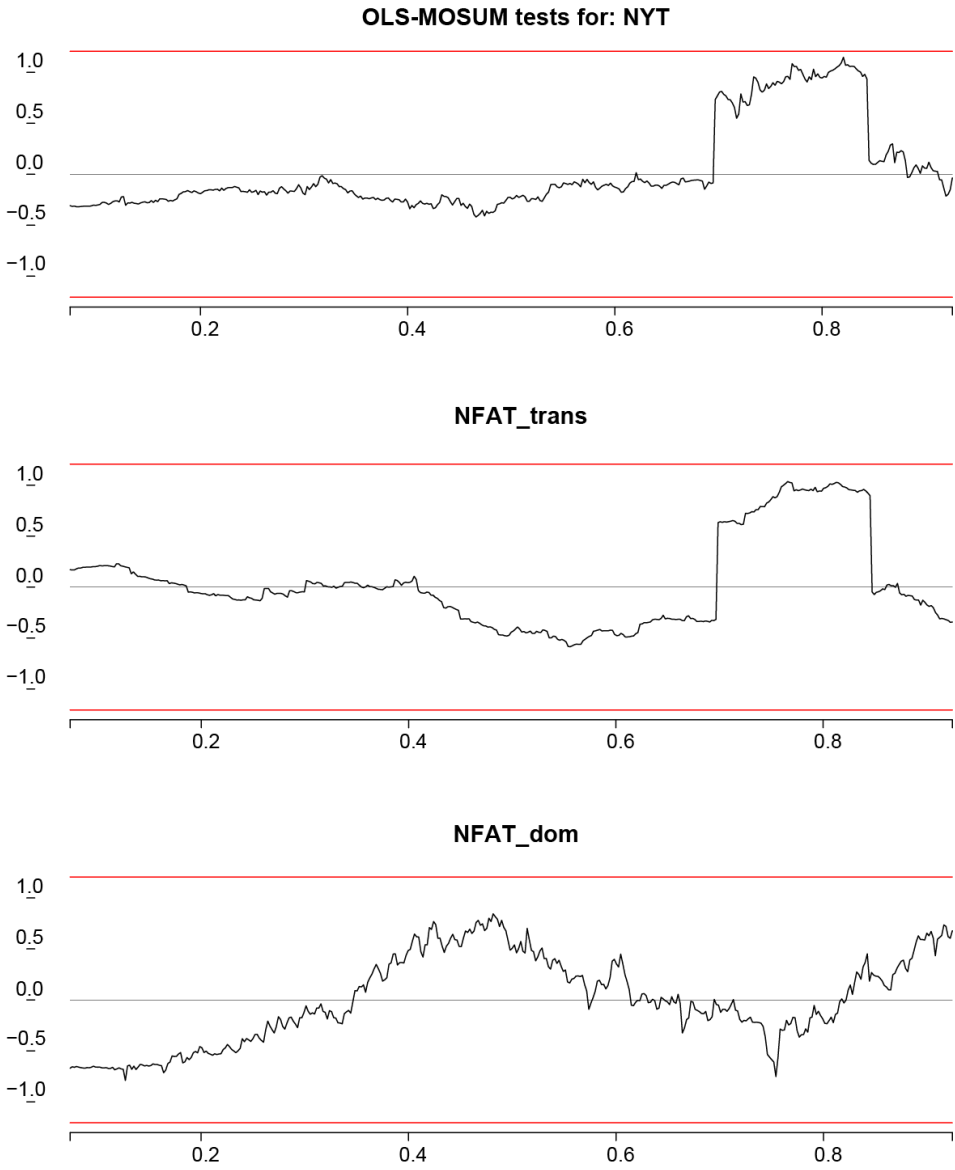
Null hypothesis	Lag order	Test statistic	<i>p</i> -value	Potential breakpoint
NYT has a unit root	14	-8.9451	<0.01	2002:07
NFAT_dom has a unit root	8	-11.6725	<0.01	1990:06
NFAT_trans has a unit root	13	-10.611	<0.01	2001:10

B Specification tests

Table B1: Specification tests for regressions in equations (6a), (6b), and (6c), p-values displayed.

Endogenous series	Lags	Ljung-Box Q	OLS-MOSUM ¹⁴
NFAT_dom	4	0.9193	0.3223
NFAT_dom_q (US&EUR)	3	0.9841	0.3502
NFAT_trans	1	0.8049	0.1691
NYT	7	0.8827	0.0845

Figure B1: Visualization of the OLS-MOSUM stability tests. In terms of parameter stability, the process for the NFAT_trans and NYT seem to be closely related.



¹⁴ The OLS-MOSUM test works with multiple structural breaks that are not specified in advance. While the CUSUM test can also handle multiple unknown breaks, it can be less sensitive to instability when the study period is relatively large (Chu et al., 1995).

Table B2: Specification tests for segmented periods, NFAT_trans, p-values displayed. The emboldened values indicate the chosen number of lags.

Segment	obs.	lags	Ljung-Box Q	OLS-MOSUM
1970:01–1999:08	356	1	0.9374	0.06968
	356	2	0.9445	0.04567
	356	3	0.9981	0.07298
	356	4	0.9797	0.07562
	356	5	0.947	0.08499
	356	6	0.947	0.1329
	356	7	0.9806	0.132
	356	8	0.7188	0.2024
1999:09–2003:08	48	1	0.5709	0.3343
	48	2	0.8419	0.3331
	48	3	0.6142	0.434
	48	4	0.6427	0.4857
	48	5	0.0658	0.5889
	48	6	0.01503	0.6781
2003:09–2010:12	88	1	0.9712	0.5251
	88	2	0.9068	0.4433
	88	3	0.835	0.4788
	88	4	0.6623	0.4596
	88	5	0.5327	0.4814
	88	6	0.8937	0.4736

Table B3: Specification tests for segmented periods, NYT, p-values displayed. The emboldened values indicate the chosen number of lags.

Segment	obs.	lags	Ljung-Box Q	OLS-MOSUM
1970:01–1998:07	343	6	0.187	0.2884
	343	7	0.9613	0.2843
	343	8	0.07237	0.2877
	343	9	0.05375	0.2909
1998:08–2002:07	48	1	0.7796	0.03515
	48	2	0.921	0.05759
	48	3	0.4519	0.1593
	48	4	0.00499	0.2606
	48	5	0.3204	0.2416
	48	6	0.187	0.2884
	48	7	0.2297	0.2843
2002:08–2010:12	101	1	0.2006	0.4263
	101	2	0.3185	1.759e-07
	101	3	0.6899	0.01858
	101	4	0.9021	0.02382
	101	5	0.4947	0.03104
	101	6	0.2006	0.4263
	101	7	0.8276	0.1835
	101	8	0.5471	0.1375

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