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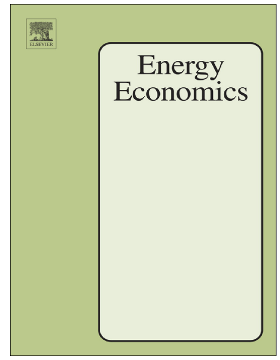

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Abstract

Rare earth elements (REEs) have gained increasing attention recently for several key reasons: 1) they are vital to many strategic industries, 2) they are relatively scarce, 3) they frequently exhibit high price fluctuations, 4) China holds a quasi-monopoly on their mining, and 5) China's REE policy, which was overly restrictive and led to a formal complaint from the U.S., Japan, and the EU at the World Trade Organization (WTO) in 2012. This paper investigates whether the announcement of a WTO dispute resolution case has the power to fundamentally change market dynamics. We find empirical support for this notion because REE prices exhibit a structural break around the announcement of the WTO dispute, and show lower variance ratios for all tested REEs afterward. This indicates a tendency toward efficiency, although REE prices still do not follow a random walk. Similarly, we find that stock price informativeness of companies in the REE industry increases after the announcement, reflecting more firm-specific than marketwide information and less governmental influence. Finally, we show that model uncertainty for option pricing models decreases, which we measure by the lower pricing differences among them.

JEL Classification: F13, G14, Q02, Q38

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1. Introduction

The rare earth elements (REE) market has gained increasing attention in recent years because REEs are of paramount and strategic importance for a variety of green- and high-technology products, such as hybrid and electric cars, wind energy turbines, photovoltaic cells, mobile phones, hard and CD drives, and permanent magnets (Van Gosen et al., 2014).[1] However, well above 80% of global REE mine production and more than 50% of worldwide REE reserves are located in China. The recent bankruptcy filing of Molycorp on June 25, 2015, which was one of the largest REE mining corporations outside China, may further increase the worldwide dependence on REEs mined in China in the future (see McCarty and Casey, 2015; Miller and Zheng, 2015; and Reuters, 2015). Thus, China essentially has global dominance over and control of the REE market, and it only intensified with the Chinese Ministry of Commerce's 2006 announcement of export restrictions in the form of export quotas (the so-called MOFCOM announcements) (see Fig. 1, step 1). These export quotas, and particularly a 40% reduction in production in 2010, were presumably the cause of the dramatic price increases seen for foreign REE prices (or so-called FOB – free on board – prices) during the second half of 2011. Domestic REE prices (or so-called China prices) were substantially lower, placing foreign competitors in the REE market at a disadvantage.

– Please insert Figure 1 about here –

In fact, the export restrictions on REEs and their overall pricing system were the catalyst for the U.S., Japan, and the EU to lodge formal complaints against China at the World Trade Organization (WTO) on March 13, 2012 (hereafter, the WTO event).[2] This has led to a series of

[1] For extensive summaries, see, e.g., Massachusetts Institute of Technology's Energy Initiative/APS Panel on Public Affairs/The Materials Research Society (2010), APS Panel on Public Affairs/The Materials Research Society (2011), Goonan (2011), and Binnemans et al. (2013).
[2] See Online Appendix A.9 for a summary of the main events of the WTO dispute resolution case.

reports concerning how to better secure access to REEs by the U.S. Congress (see Fig. 1, step 2, as well as the European Commission, 2012; Morrison and Tang, 2012; Bailey Grasso, 2013; and Humphries, 2013).

The WTO generally focuses on negotiating new agreements for reducing trade burdens among its member states. Furthermore, it is also heavily involved in evaluating complaints and issuing sanctions in case of violations of agreements in place (see Liebman and Tomlin, 2008). However, besides the shareholder value effects for firms affected by the consequences of WTO disputes and its adjudicates, we know very little about how the WTO's actions affect the underlying markets (see, e.g., Lenway et al., 1990; Lenway et al., 1996; and Liebman and Tomlin, 2007, 2008, for research on shareholder behavior).[3] While the majority of studies analyze the effect of WTO *rulings* on share prices of companies in the industries affected by the respective ruling, Desai and Hines Jr. (2008) were the first to investigate the effects of the *announcement* of the filing of a complaint at the WTO.

However, the focus so far has been on examining stock price reactions using event studies. In particular, there have been only limited efforts to ascertain the WTO's effectiveness, and whether the announcement of a dispute resolution case has the power to fundamentally change market dynamics. We believe this is an important question, because an effective enforcement mechanism to resolve disputes is essential to promoting compliance with the WTO's rules. Moreover, changes in the underlying market dynamics are presumably more long-lived than shareholder behavior (Davey, 2003).

We address this research question empirically by analyzing the announcement effects of the WTO event on the REE market. If the complaint is credible, and the WTO has effective enforcement power, we expect to find that the Chinese government reconsidered, and adapted

[3] See Liebman and Tomlin (2007) for an excellent literature review of GATT/WTO disputes.

their REE policy to somewhat reduce their influence on REEs (see Fig. 1, step 3). We observe direct evidence for this behavior in the Chinese government's statement at the end of 2014 that they would stop using export quotas for REEs because of the WTO's ruling that the practice is at odds with the General Agreement on Tariffs and Trade (GATT) (see *WSJ*, 2015).[4] However, if credible actions by the Chinese government had already been undertaken after the initiation of the WTO dispute resolution case in 2012, we anticipate finding noticeable effects on the price dynamics of REEs. This would provide evidence that even the launch of a WTO trial has the power to spur policy changes and add to the findings of Desai and Hines Jr. (2008). Accordingly, the launch of the dispute resolution case on March 13, 2012 serves as a natural experiment.[5]

First, we expect to observe a structural break in the time series of REE prices around the announcement date, which would indicate a change in their dynamics. Using Bai and Perron's (1998, 2003a, 2003b) procedure for multiple structural changes of a time series, we find support for this notion for most REEs, regardless of whether FOB or China REE prices are used. Next, we calculate Lo and MacKinlay (1988) and Wright (2000) pre- and post-announcement variance ratios in order to infer whether the market for REEs is efficient. We find that pre-announcement REE prices do not follow a random walk, and thus price formation is not efficient. However, after the initiation of the WTO trial, we find considerably lower variance ratios for all tested REEs. This indicates a general tendency toward efficiency, but prices still do not follow a random walk.

[4] See TMR (2014) and WTO (2014a, 2014b, 2014c, 2014d, 2014e, 2014f) for more information on the dispute resolution case.

[5] We use the announcement date of the WTO dispute resolution case as our event date because this is, in our view, the "cleanest" and earliest event date one could use in our "one-case event study." While there were rumors of a potential launch of a WTO dispute resolution case against China, it became a certainty on March 13, 2012, when the U.S., Europe, and Japan launched their case. Note that any information processing due to rumors or speculation before the event date would have happened during our "pre-WTO" period. This anticipation would make it harder to find statistically significant results in our "post-WTO" period. Therefore, we interpret our derived results as conservative (see Online Appendix A.9 for a detailed overview). Other event dates during the dispute resolution case could also be used to determine the importance of other announcements, but it would result in confounding events and would make it virtually impossible to isolate the "importance" of a particular event during the dispute resolution case.

Second, if the Chinese government reassesses its REE policy and ultimately interferes less in the market, this policy change should be reflected in the stock price informativeness of companies in the REE market. The idea is that less informative stock prices convey less firm-specific information. Their stock price variation can thus be explained to a larger extent by marketwide factors (see Gul et al., 2010).

If this conjecture holds, and the Chinese government changes its REE policy as a result of the announcement of the WTO dispute resolution case, we would expect the informativeness of REE stock prices to increase. This is because firm-specific information tends to take precedence over marketwide factors such as REE prices. We show that the stock price informativeness of companies in the REE industry increases statistically significantly after the WTO event compared to other industries. This result is consistent with the view that less intervention by the Chinese government after the announcement of the WTO trial led to REE company stock prices that were more reflective of firm-specific than marketwide information.

Third, we analyze model uncertainty for the three option pricing models of Black-Scholes (1973), Duan (1995), and Heston and Nandi (2000) by comparing their pre- and post-WTO event price differences. Similarly to the previous argument, that less intervention in the REE market would reduce exogenously driven erratic price behavior of REEs, the post-WTO event pricing differences among the option pricing models should be lower if the Chinese government has reduced its interference. Our results strongly support this notion. The discrepancies between the models are not only statistically significantly lower, but, with average reductions of more than 70%, economically meaningful as well. Note that the high model uncertainty prior to the WTO event presumably contributed to the blocking of a proposed REE derivatives market. In line with this reasoning, after the WTO event, the plan to open an REE (derivatives) exchange in China gained increasing attention. In March 2014, the first REE exchange, the Baotou Rare Earth

Products Exchange, was established, but it does not offer any REE derivatives products (see Bloomberg News, 2014). However, the Shanghai Futures Exchange announced in mid-2014 that they would enhance their product portfolio by launching REE derivatives (Shen, 2014). The introduction of liquid REEs and derivatives markets are necessary for suppliers and for user firms to, e.g., reduce earnings fluctuations, which will decrease bankruptcy risk and delivery bottlenecks (see Fig. 1, step 4).

Fourth, as a whole, our study contributes to the extant and growing literature on the effects of trade disputes and WTO rulings in several ways. While current literature concentrates primarily on analyzing shareholder wealth transfers of actual WTO *rulings*, we provide a unique perspective by analyzing how even the *announcement* of a WTO dispute can induce governmental changes in existing policies (see Lenway et al., 1990; Lenway et al., 1996; and Liebman and Tomlin, 2007, 2008). We extend and complement these research findings. We provide evidence that not only does the price-generating process of REEs change after the WTO event, but the stock price informativeness for REE companies increases. In contrast to Müller et al. (2016), we do not analyze effects in response to MOFCOM export quota announcements using event studies. Instead, we go one step further and focus on the commencement of the dispute resolution case at the WTO, which may to some extent be a consequence of the quotas.

In summary, we interpret our results as strong support for the notion that governments accused of violating GATT react to the announcement with policy changes. These actions, however, are not simply "window dressing," but rather have statistically significantly and economically meaningful marketwide effects and implications.

The remainder of this paper is structured as follows. Section 2 provides an overview of the methodology used for our analyses, while section 3 describes our data gathering process. Section 4 presents the empirical results of our study. Section 5 concludes.

2. Methodology

2.1. Variance Ratio Tests

Variance ratio tests have been used to test for the random behavior of price movements since the 1960s. One of the first applications was in Alexander (1961), who calculated variance ratios for the time series used by Kendall (1953), and concluded that the price series followed a random walk. Since then, time series ranging from financial market data such as stock returns (see, for example, Lo and MacKinlay, 1988; Poterba and Summers, 1988; and Ayadi and Pyun, 1994) and exchange rates (Liu and He, 1991) to economic time series such as GDP (see, e.g., Cochrane, 1988) have been analyzed with the help of variance ratio tests. Charles and Darné (2009) provide a solid overview.[6]

One of the most common variance ratio tests is the methodology proposed by Lo and MacKinlay (1988). If the time series at hand is indeed a random process, the variance of the increments of the process should be proportional to the sampling horizon because it is linear in its sampling interval (see Cochrane, 1988; Lo and MacKinlay, 1988; and Poterba and Summers, 1988). In other words, q times the variance of a time series' first differences should equal the variance of a time series' q-differences (see Liu and He, 1991; Ayadi and Pyun, 1994; Wright, 2000; and Charles and Darné, 2009). Accordingly, if a time series follows a random walk process, the variance of two-month first differences should equal twice the variance of one-month first differences. Comparing the variances for different sample intervals then allows us to test whether the time series at hand is governed by a random process (because the variance ratios should equal 1).

Formally, we calculate the variance ratios as follows:

[6] There are several different variance ratio tests. See, for example, Chow and Denning (1993) and Whang and Kim (2003) for multiple variance ratio tests, and Wright (2000) for a non-parametric variance ratio test. However, most of these tests are based on the main idea originally put forward in Lo and MacKinlay (1988).

$$M_r(q) \simeq \frac{2(q-1)}{q}\hat{\rho}(1) + \frac{2(q-2)}{q}\hat{\rho}(2) + \cdots + \frac{2}{q}\hat{\rho}(q-1), \tag{1}$$

where $\hat{\rho}(k)$ represents the estimator for the autocorrelation coefficient of order k for the first differences of the time series at hand (see Lo and MacKinlay, 1988). To allow for deviations of the time series from normality and time-varying volatility, Lo and MacKinlay (1988) develop a test statistic that allows for heteroscedasticity:

$$z^*(q) = \sqrt{nq}\bar{M}_r(q)/\sqrt{\hat{\theta}}, \tag{2}$$

where θ denotes the asymptotic variance of $\bar{M}_r(q)$.

However, conventional variance ratio tests, such as the procedure in Lo and MacKinlay (1988), suffer from size distortions and lack power in case of non-normal data (Wright, 2000; Charles and Darné, 2009). Hence, Wright (2000) proposes a non-parametric rank-based variance ratio test that is more precise and more powerful than alternatives such as fractional integration: Let y_t be a series of asset returns and $r(y_t)$ be the rank of y_t among $y_1, y_2, \ldots, y_T$. We can define two series, $r_{1t} = \left(r(y_t) - \frac{T+1}{2}\right)/\sqrt{\frac{(T-1)(T+1)}{12}}$ and $r_{2t} = \Phi^{-1}\left(\frac{r(y_t)}{T+1}\right)$, where the former is a linear transformation of the ranks so that it has mean 0 and variance 1, while the latter is the inverse normal, or van der Waerden score, and has mean 0 and variance ~1. Φ is the standard normal cumulative distribution function, and r_{1t} and r_{2t} are then substituted for the returns in the test statistic.[7]

[7] The variance ratio tests are based on log prices (levels), which could potentially be affected by different price levels. Urquhart and McGroarty (2016) provide evidence on this issue. They fail to find any consistent pattern of market conditions (e.g., bull or bear markets) driving the variance ratio test results. Instead, they find the results are dependent on individual market conditions, because different markets react differently to varying market conditions.

2.2. Multiple Structural Change Tests

Besides using variance ratio tests to gauge whether the commencement of a WTO trial has an efficiency-enhancing effect, we analyze whether we can identify structural changes in the dynamics of the time series of REE prices surrounding that date. Because the exact number of structural breaks in a time series is often unknown ex ante, Bai and Perron (1998, 2003a, 2003b) propose a series of tests to determine the number of structural changes in a time series endogenously. In particular, Bai and Perron (1998, 2003a, 2003b) develop a sup$F_T(\ell+1|\ell)$ test of the null hypothesis of ℓ breaks versus the alternative hypothesis of $\ell+1$ breaks. In principle, this method applies $\ell+1$ tests of the null hypothesis of no structural break versus the alternative hypothesis of one structural break, which is then applied to each segment containing the observations $\hat{T}_{i-1}$ to $\hat{T}_i$ $(i=1,\ldots,\ell+1)$ where $\hat{T}_0=0$ and $\hat{T}_{\ell+1}=T$. An additional break is included as long as the overall minimum value of the sum of squared residuals is sufficiently smaller than the sum of squared residuals from the model with ℓ breaks. Formally, the test statistic is given as:

$$F_T(\ell|\ell+1)=\left\{S_T\left(\hat{T}_1,\ldots,\hat{T}_\ell\right)-\min_{1\le i\le \ell+1}\inf_{\tau\in\Lambda_{i,\eta}} S_T(\hat{T}_1,\ldots,\hat{T}_{i-1},\tau,\hat{T}_i,\ldots,\hat{T}_\ell)\right\}/\hat{\sigma}^2. \tag{3}$$

This procedure uses a general-to-specific modeling strategy, and is recommended by Bai and Perron (1998, 2003a, 2003b) for determining the number of breaks.

2.3. Stock Price Informativeness

2.3.1. Measuring Stock Price Informativeness

Our goal is to measure the impact of the WTO trial on the stock price informativeness of REE companies. We thus closely follow the methodology of Morck et al. (2000), Jin and Myers (2006), Gul et al. (2010), and Tan et al. (2015). The underlying idea is to calculate the stock price informativeness of companies in the REE industry, and to compare it for each company i before and after the WTO event. The measure of choice is stock price synchronicity as used by Gul et al. (2010), which is defined as:

$$SYNCH_{i,\tau} = log\left(\frac{R_{i,\tau}^2}{1-R_{i,\tau}^2}\right), \tag{4}$$

where $SYNCH_{i,\tau}$ is the annual stock price synchronicity of company i in year (period) τ, log refers to the natural logarithm, $\tau = 0$ refers to the WTO event, and $\tau = -1$ ($\tau = Pre$) to the year (period) before the WTO event.[8] $R_{i,\tau}^2$ is obtained from the regression in Equation (5):[9]

$$R_{i,t} = \alpha_i + \beta_1 \cdot MR_t + \beta_2 \cdot MR_{t-1} + \beta_3 \cdot IR_t^{(j)} + \beta_4 \cdot IR_{t-1}^{(j)} + \varepsilon_{i,t}, \tag{5}$$

where $R_{i,t}$ is the daily stock return of company i on day t, and MR_t is equal to the market index return on day t. The proxy for the market index is the daily aggregated market return index with cash dividends reinvested (volume-weighted) for all A shares (listed on either the Shanghai or Shenzhen stock exchange) from CSMAR.[10] $IR_t^{(j)}$ represents the industry index return on day t for the respective industry j, based on the six industry classifications from *Industry Code A.*

[8] The pre-WTO period ($\tau = Pre$) covers August 20, 2009 through March 12, 2012 (2.5 years), and the post-WTO period ($\tau = Post$) covers March 14, 2012 through September 30, 2014 (2.5 years). We set the beginning (end) date of the pre-WTO period (post-WTO period) to have an equal number of trading days in both periods. This is done to obtain a balanced sample even if the calendar days suggest otherwise because of public holidays and stock exchange closures.

[9] In unreported results, we tested the robustness of our results for a reduced market model with the following form: $R_{i,t} = \alpha_i + \beta_1 \cdot MR_t + \varepsilon_{i,t}$. The results are highly similar in terms of magnitude and statistical significance and are available from the authors upon request.

[10] In unreported results, we used the equally weighted index instead, and our results are virtually identical to those obtained using the value-weighted index. Tables are available from the authors upon request.

We consider only three of the six industries simply because REE industry companies are only present in *Utilities*, *Conglomerates,* and *Industry*. A company is included in the respective industry if 1) it is active during the 2.5 years before and after the beginning of the WTO trial on March 13, 2012, 2) we observe minimum liquidity, meaning that daily return data is available for at least 200 trading days for the two years before and after the WTO event,[11] and 3) the equally weighted industry index is unique for every company i, because all companies within the same industry (satisfying conditions (1) and (2)) are considered except company i. Our final sample is comprised of 834 companies (with A shares) in the three industries *Utilities* (111), *Conglomerates* (34), and *Industry* (689) that satisfy these conditions.

The idea is as follows: High R^2s in the market model regression (see Equation (5)) mean that most of a company's price fluctuations are explained by marketwide and/or industrywide information (high stock price synchronicity), and fewer are explained by firm-specific information (low stock price informativeness). Put differently, the higher the R^2, the lower the measure for stock price informativeness, and thus a company's stock price will contain less company-specific information (see, e.g., Chen et al., 2007, for a more detailed discussion). In line with Morck et al. (2000) and Gul et al. (2010), we apply a logistic transformation to circumvent the bounded nature of R^2 within [0, 1] and calculate $SYNCH_{i,\tau}$ (see Equation (4)). Thus, a higher R^2 leads to a higher stock price synchronicity measure $SYNCH_{i,\tau}$, and implies greater co-movements with market- and/or industrywide information and fewer with company-specific information. It follows that developed markets such as the U.S. have lower R^2 and lower (usually

[11] For the two half-year periods (from 2 to 2.5 years) before and after the WTO event, we require at least daily available return data for 100 trading days.

negative) $SYNCH_{i,\tau}$ than emerging markets with less investor protection, which could discourage informed investors from trading (see, e.g., Morck et al., 2000).[12]

Finally, and analogously to our previous approach, we run the following regression to implicitly identify companies in the REE industry with a statistical method. We again calculate the R^2 for all 834 companies that satisfy the three conditions above:

$$R_{i,t} = \alpha_i + \beta_1 \cdot REER_t + \varepsilon_{i,t}, \tag{6}$$

where $REER_t$ is the respective "REE index" return on day t. Our rationale is that higher R^2s mean a higher proportion of companies' stock price variations can be explained by REE price movements. Therefore, companies have a higher exposure to the REE industry.

However, there is no single best way to identify companies in the REE industry based on this search strategy, so we tried a few alternatives for $REER_t$. We first construct four REE indices containing the four most important elements, cerium (Ce), lanthanum (La), neodymium (Nd), and yttrium (Y), which make up about 90% of total REE usage (see Goonan, 2011). We then calculate a usage-weighted index and an equally weighted index based on FOB (foreign) and China (domestic) prices (see Morrison and Tang, 2012; and Bailey Grasso, 2013).

In the next step, we select the 10, 25, and 50 companies with the highest R^2 in Equation (6) in the following selection periods: year 1 prior to the WTO event ($\tau = -1$), year 2 prior to the WTO event ($\tau = -2$), and the pre-WTO event period ($\tau = Pre$) (which covers the 2.5-year period before the event). This leaves us with thirty-six possible combinations: 3 (10, 25, and 50 companies) × 3 ($\tau = -1, -2$, and pre-selection period) × 2 (REE Index based on FOB and China prices) × 2 (equally and usage-weighted REE Indices) = 36. We manually checked five companies identified as having the highest correlation with the REE indices, and found that all

[12] Note that stock price synchronicity is not affected by REE price levels, which change substantially during the pre-WTO period. This is because R^2 is based on stock returns, and REE prices are used only to identify companies in the REE space. As a further check, the company descriptions in Table A.1 denote whether the identified companies are operating in the REE industry.

operated directly in the REE industry or held stakes in companies in the REE industry (see Table A.1 in the appendix). Therefore, we are convinced that this procedure is suitable for identifying companies with REE exposure.

In unreported results, we complemented this implicit approach for REE company identification with an explicit approach. Similarly to Müller et al. (2016), we used only publicly listed companies on the Shenzhen or Shanghai stock exchange that are mentioned in MOFCOM's REE export quota announcements. This approach has a clear advantage in that those companies have been determined to be in the REE industry. However, that advantage comes with certain disadvantages, because we capture only REE mining companies, or companies that are relatively early in the REE supply chain (refining). Chinese high-tech companies, for example, are also arguably affected by REE price development. Furthermore, we are left with a maximum of fourteen companies (not all were listed for the entire pre-/post-WTO period), which most likely significantly underestimates the industry. Nevertheless, the results of this approach are similar in terms of magnitude to the implicit approach, but with a lower statistical significance. This is attributable at least to some extent to the smaller sample size.

2.3.2. Testing for REE Stock Price Informativeness Differences around the WTO Event

To test whether our prediction of increasing REE stock price informativeness after the WTO event is correct, we first provide univariate evidence by comparing REE stock price synchronicity before and after the event. We then provide multivariate evidence, while controlling for the reference group and firm characteristics. The univariate test is as follows:

$$\Delta SYNCH_{i,[\tau_1;\tau_2]} = SYNCH_{i,\tau_2} - SYNCH_{i,\tau_1}, \tag{7}$$

where $SYNCH_{i,\tau}$ corresponds to the calculation from Equation (4), τ_1 stands for the respective period prior to the WTO event, and τ_2 stands for the period afterward. The $[\tau_1;\tau_2]$ periods we compare in the univariate tests are $[\tau_{-1};\tau_{+1}]$, $[\tau_{-1};\tau_{+2}]$, and $[\tau_{Pre};\tau_{Post}]$, where $[\tau_{-1};\tau_{+1}]$ is the one-year period pre- and post-WTO event, and $[\tau_{Post}; \tau_{Pre}]$ is the symmetrical 2.5-year period pre- and post-WTO event.

If we observe an increase in stock price informativeness (corresponding to a decrease in stock price synchronicity) following the WTO event, we would expect the R^2_{i,τ_2} from Equation (5) (and thus the resulting $SYNCH_{i,\tau_2}$) to be lower afterward than before (τ_1). Consequently, we expect the $\Delta SPI_{i,[\tau_1;\tau_2]}$ for the REE industry companies to be statistically significantly lower than zero. The following example illustrates this prediction using Hubei Biocause Pharmaceutical Co. Ltd., which Equation (5) identifies as being among the ten companies with the highest R^2 in the pre-WTO period. For this company, which has major holdings in China Minmetals Corporation (a mining company that also operates in the REE industry), we observe a -0.12 stock price synchronicity one year before the WTO event ($SYNCH_{i,-1}$) (corresponding to an $R^2_{i,-1}$ of 0.47). In the year after the WTO event, the stock price synchronicity decreased to -0.49 (corresponding to an $R^2_{i,+1}$ of 0.38). This resulted in a -0.37 change in stock price synchronicity ($\Delta SYNCH_{i,[+1;-1]}$), which means that stock price informativeness increased as hypothesized in response to the WTO event.

To test for a relationship between the WTO event and the change in stock price informativeness in an OLS regression framework, we estimate the following basic structure of the regression model. This allows us to control for further potentially explanatory factors:

$$SYNCH_\tau = \alpha + \beta_1 \cdot PreWTO(\tau_1) \times REEC + \beta_2 \cdot PostWTO(\tau_2) \times REEC$$
$$+ \sum_j \gamma_j \cdot Controls_{j,\tau-1} + \xi_\tau + \varepsilon_\tau, \quad (8)$$

where the dependent variable $SYNCH_\tau$ is the annual stock price synchronicity calculated as in Equation (4), $PreWTO(\tau_1)$ is a dummy variable equal to 1 for the pre-WTO period, and 0 otherwise, $PostWTO(\tau_2)$ is a dummy variable equal to 1 for the post-WTO period, and 0 otherwise, and $REEC$ is also a dummy variable equal to 1 if the company is identified as a REE company (hereafter, $REEC$) according to Equation (6), and 0 otherwise. Our control variables are government holdings ($Top\,Gov$), trading volume ($Volume$), firm size ($Size$), debt ratio ($Leverage$), standard deviation of the return on assets ($Std(RoA)$), market-to-book ratio (M/B), number of companies in the industry (Ind_Num), and industry size (Ind_Size) (see Table A.2 in the appendix for variable descriptions and calculations). ξ_τ are year fixed effects to control for potential macroeconomic trends in China. Given our previous argumentation, we do not show $PreWTO(\tau_1)$ or $PostWTO(\tau_2)$ explicitly in the regression tables. We use robust standard errors and omit firm-level notations for clarity in presentation of Equation (8).

The spirit of the regression is similar to a difference-in-differences (DiD) approach. The WTO event serves as an exogenous event that is expected to have an effect on $REEC$s but not on other companies. Therefore, the $REECs$ serve as a so-called treatment group, while all other firms serve as the control group. Within our regressions, we divide the sample into two periods, before ($PreWTO(\tau_1)$) and after ($PostWTO(\tau_2)$) the announcement of the WTO trial. We are predominantly interested in the DiD coefficients β_1 and β_2. If the *REECs* exhibit higher stock price synchronicity (lower informativeness) before the WTO event than the control group, we would expect β_1 to be positive. After the exogenous event, we expect the stock price synchronicity measure $SYNCH$ to drop compared to the pre-WTO period, thus implying $\beta_1 > \beta_2$. This would indicate that the $SYNCH$ measure of *REECs* decreased relative to the unaffected control group companies after the WTO event. A change from a positive sign for β_1 to a negative one for β_2 is arguably the strongest effect, meaning that *REECs* exhibited higher values of $SYNCH$ than the control group before the WTO event, and lower values afterward. If the WTO event did not affect the $SYNCH$ measure of *REECs*, then we expect β_2 to be insignificant.

2.4. Derivatives Pricing

Our goal is to determine whether market participants face less *option pricing model uncertainty* after the WTO event than beforehand. Our proxy measures are the pricing differences among three option pricing models: Black-Scholes (1973), Duan (1995), and Heston and Nandi (2000). We interpret larger differences as higher option pricing model uncertainty, and, if our intuition is correct, we expect the differences to decrease after the WTO event. As we mentioned earlier, there are no option exchanges for REEs yet. Thus, we are unable to compare the option pricing model results with actual option prices observed in a market. Given that, we choose three option pricing models exogenously and calculate prices under the assumption of how market participants would use them. We therefore aim not to identify the "best" option pricing model for REEs, but rather to establish that the pricing differences among models become smaller after the WTO event.

For the first "static" analyses, our research design is similar to a DiD approach, using the WTO event as an exogenous shock and comparing option price differences before and after the event. We use the term "static" because we compare option prices at two different points in time – at the WTO event date, and thirty months later (September 15, 2014). To elaborate, we compare option prices at the WTO event date ($s = 0$), as well as on September 15, 2014, considering respectively the stream of information from REE prices for the previous thirty months. To derive the option prices at the WTO event date, we use the REE price developments before the WTO trial announcement ($PreWTO$ period), and, for September 15, 2014, the REE price developments afterward ($PostWTO$ period) (see Fig. 2 for a visual representation). The comparison is based on the absolute option price differences between 1) Black-Scholes (1973) and Duan (1995), 2) Black-Scholes (1973) and Heston and Nandi (2000), and 3) Duan (1995) and Heston and Nandi (2000).[13]

We add the three option pricing differences (PD) for puts and calls for 35 [= 7 *times to maturity* × 5 *strike prices*] distinct option price difference combinations of *time to maturity* and *strike price*. The calculation is as follows:

$$PD_{PreWTO}(REE_k|\mathcal{F}_s)_k = \sum_{i=1}^{7}\sum_{j=1}^{5}\left[\left|BS_{ij} - NGARCH_{ij}\right| + \left|BS_{ij} - HN_{ij}\right| + \left|NGARCH_{ij} - HN_{ij}\right|\right], \qquad (9a)$$

$$PD_{PostWTO}(REE_k|\mathcal{F}_t)_k = \sum_{i=1}^{7}\sum_{j=1}^{5}\left[\left|BS_{ij} - NGARCH_{ij}\right| + \left|BS_{ij} - HN_{ij}\right| + \left|NGARCH_{ij} - HN_{ij}\right|\right], \qquad (9b)$$

$$DiD_k = PD_{PostWTO}(REE_k|\mathcal{F}_t)_k - PD_{PreWTO}(REE_k|\mathcal{F}_s)_k, \qquad (10)$$

where REE_k are the REEs ($k =$cerium, lanthanum, neodymium, and yttrium), and BS_{ij}, $NGARCH_{ij}$, and HN_{ij} represent the Black-Scholes (1973), Duan (1995), and Heston and Nandi (2000) option prices for either put or call options with seven *strike prices* [$i = 85\%, 90\%, \ldots, 115\%$] and five *times to maturity* as measured in months [$j = 1, 3, 6, 9, 12$]. PD_{PreWTO} refers to the sum of the differences in option prices based on the different models, and based on information thirty months prior to the WTO event date ($PreWTO$ period). $PD_{PostWTO}$ is the sum of the differences in option prices based on the different models, and based on information thirty months after the WTO event date ($PostWTO$ period). Filtration $\mathcal{F}$, i.e., $\mathcal{F}_s$, is all information available beginning thirty months prior to the WTO event until the WTO event date (-30 months $<$ WTO event date (s)), and $\mathcal{F}_t$ is all information available beginning at the WTO event date until thirty months afterward (WTO event date (s) $\leq$ +30 months (t)).

We obtain Black-Scholes (1973) option prices by using a block bootstrap approach to incorporate the possible time-dependent structure in REE returns (see appendix A.8 for more

[13] We consider GARCH option pricing because we find that the REE returns exhibit GARCH effects (see Fig A.1 in the appendix).

details). The two GARCH option prices – Duan (1995) and Heston and Nandi (2000) – are derived as follows:

a) Duan's (1995) NGARCH model:

Motivated by the success of GARCH models in estimating and forecasting volatility, Duan (1995) first proposed an NGARCH(1,1) option pricing model. Given the asset price stream, we assume the dynamics are:

$$ln\frac{S_{t+1}}{S_t} = r_f + \lambda\sqrt{h_{t+1}} - \frac{1}{2}h_{t+1} + \sqrt{h_{t+1}}\epsilon_{t+1}, \tag{11}$$

$$h_{t+1} = \beta_0 + \beta_1 h_t + \beta_2 h_t(\epsilon_t - \gamma)^2, \tag{12}$$

while the risk-neutral process is:

$$ln\frac{S_{t+1}}{S_t} = r_f - \frac{1}{2}h_{t+1} + \sqrt{h_{t+1}}v_{t+1}, \tag{13}$$

$$h_{t+1} = \beta_0 + \beta_1 h_t + \beta_2 h_t(v_t - \omega)^2, \tag{14}$$

where $\omega = \gamma + \lambda$.

b) Heston and Nandi's (2000) model:

Given the asset price stream, we assume the dynamics are:

$$ln\frac{S_{t+1}}{S_t} = r_f + \left(\lambda - \frac{1}{2}\right)h_{t+1} + \sqrt{h_{t+1}}\epsilon_{t+1}, \tag{15}$$

$$h_{t+1} = \beta_0 + \beta_1 h_t + \beta_2\left(\epsilon_t - \gamma\sqrt{h_t}\right)^2, \tag{16}$$

while the risk-neutral process is:

$$ln\frac{S_{t+1}}{S_t} = r_f - \frac{1}{2}h_{t+1} + \sqrt{h_{t+1}}v_{t+1}, \tag{17}$$

$$h_{t+1} = \beta_0 + \beta_1 h_t + \beta_2\left(v_t - \omega\sqrt{h_t}\right)^2, \tag{18}$$

where $\omega = \gamma + \lambda$.

Duan (1995) assumes a locally risk-neutral valuation relationship (LRNVR) to measure the return process by adjusting asset-specific drift terms under the risk-neutral distribution.

Heston and Nandi (2000) generally follow the LRNVR concept, and formulate an affine GARCH model that yields a closed-form solution. Note that such risk neutralization is conducted by means of linear pricing kernels and is not unique. For both GARCH models, we optimize the following likelihood to get the GARCH parameters:

$$ln\, L \propto -\frac{1}{2}\sum_{t=1}^{T}\left\{ln\, h_t + \left(R_t - r - \left(\mu - \frac{1}{2}\right)h_t\right)^2/h_t\right\}. \tag{19}$$

The GARCH option prices are then generated by Monte Carlo simulations,[14] with

$$C_t = e^{-rT}E^Q[max(S_T - K, 0)]. \tag{20}$$

This *static* approach has the clear advantage of being intuitive. However, it has the disadvantage that we cannot rule out alternative explanations such as non-normally distributed REE returns, or results that are sensitive to choosing only two points in time for the option price calculations. Furthermore, we can only evaluate the economic significance of the results, not the statistical significance.

To address these concerns, we apply a *consecutive* "triple-difference" (DDD) approach, and construct three subperiods of equal length: sixty months prior to thirty months prior to the WTO event ($PrePreWTO$ period), thirty months prior to the WTO event to the WTO event ($PreWTO$ period), and from the WTO event to thirty months afterward ($PostWTO$ period) (see Fig. 2, panel B, for a graphic representation). The spirit of the DDD is that it identifies whether the differences in option prices between the first two "control" subsamples ($PrePreWTO$ and $PreWTO$ period) are consistently lower than those between the latter two "treatment" subsamples ($PreWTO$ and $PostWTO$ period), which would indicate a clear influence of the WTO event.

[14] One of the major advantages of Monte Carlo option pricing is its flexibility. For examples, see, e.g., Ibáñez and Zapatero (2004).

To see whether results are robust for different *times to maturity*, *strike prices*, and option pricing models, we provide evidence separately and split the pricing differences. The following illustrates the formulas for call options based on pricing differences between the Black-Scholes (1973) and Duan (1995) models ($|BS - NGARCH|$) with a *time to maturity* of one month ($y = 1m$). This results in 210 option prices [210 = 7 *strikes* and 30 *consecutive* months] per model. The calculations for the DDDs are as follows:

$$PD_{PrePreWTO}(REE_k|\mathcal{F}_r)_k(y = 1m; |BS - NGARCH|) = \sum_{m=-48}^{-25} \sum_{i=0.85}^{1.15} |BS_{mi} - NGARCH_{mi}|, \quad (21a)$$

$$PD_{PreWTO}(REE_k|\mathcal{F}_s)_k(y = 1m; |BS - NGARCH|) = \sum_{m=-24}^{-1} \sum_{i=0.85}^{1.15} |BS_{mi} - NGARCH_{mi}|, \quad (21b)$$

$$PD_{PostWTO}(REE_k|\mathcal{F}_t)_k(y = 1m; |BS - NGARCH|) = \sum_{m=0}^{+24} \sum_{i=0.85}^{1.15} |BS_{mi} - NGARCH_{mi}|, \quad (21c)$$

$$\overline{DiD}_k(PreWTO - PrePreWTO)_{(y=1m;\,|BS-NGARCH|)} = \frac{1}{210}(PD_{PreWTO}(REE_k|\mathcal{F}_s)_k - PD_{PrePreWTO}(REE_k|\mathcal{F}_r)_k) \quad (22a)$$

$$\overline{DiD}_k\big(PreWTO - PostWTO\big)_{(y=1m;\,|BS-NGARCH|)} = \frac{1}{210}(PD_{PreWTO}(REE_k|\mathcal{F}_s)_k - PD_{PostWTO}(REE_k|\mathcal{F}_t)_k) \quad (22b)$$

$$DDD_{k_{(y=1m;\,|BS-NGARCH|)}} = \overline{DiD}_k\big(PreWTO - PostWTO\big) - \overline{DiD}_k\big(PreWTO - PrePreWTO\big) \quad (22c)$$

To obtain the Black-Scholes (1973) option prices, we construct a *consecutive* monthly return path that estimates the volatility. In other words, for the option prices calculated, e.g., 59 months before the WTO event date, we use only REE price information from the previous month (-60). For the following *consecutive* month (-58), we use only information from the previous month (-59). We thus refer to this approach as *consecutive*, because we always use the prior month's information to price the options.

For each sample period, we estimate the Heston and Nandi (2000) and NGARCH models with maximum likelihood estimation based on the REE price information available in the respective subperiod. We further simulate the GARCH option prices by using Monte Carlo methods based on filtered end-of-month volatilities. Despite the analytical solutions from both the Heston and Nandi (2000) and Black-Scholes (1973) models, we use simulations to price the options in order to ensure a fair comparison.

Finally, we test whether the DiD mean values from the "control" period ($\overline{D\iota D}_k(PreWTO - PrePreWTO)$) are lower than those from the "treatment" period ($\overline{D\iota D}_k(PreWTO - PostWTO)$), which would mean that DDD_k is statistically significantly greater than zero. Note that the DDD approach cancels out any potential price level effect on the option price.

– Please insert Figure 2 about here –

3. Data

Our data span January 2004 through September 2014, and come from the following databases:

- The REE prices for La, Ce, Nd, and Y come from the Asian Metal database. Because of the dual pricing system of REEs, we consider both FOB (export) and China (domestic) prices, which are in USD/kg. If USD prices are not available, we use the official CNY/USD exchange rate available from the Federal Reserve System. Note that we do not use metal prices because oxides have much higher trading volumes (see Jackson and Christiansen, 1993).
- To study market informativeness, we calculate usage- and equally weighted REE indices based on both domestic and FOB prices. See Appendix A.3 for a robustness check for China prices.
- Chinese companies in the REE industry must be listed on either the Shenzhen or the Shanghai stock exchange. Firm characteristics (government holdings ($Top\ Gov$), trading

volume ($Volume$), firm size ($Size$), debt ratio ($Leverage$), standard deviation of return on assets ($Std(RoA)$), market-to-book ratio (M/B), number of companies in the industry (Ind_Num), industry size (Ind_Size)), and stock price data for all Chinese companies come from or are calculated by using CSMAR. We exclude initial listing day returns due to extreme returns. If a company's stock is not traded, we set the return to zero. Finally, we calculate the market index as the value-weighted average index of all listed companies on the Shenzhen or Shanghai stock exchange, and similarly an industry index as the value-weighted average index for the respective industry.

- We use the announcement date of the WTO dispute resolution case as our event date, because it is, in our view, the "cleanest" and earliest event date one could use in our "one-case event study." Although there had been rumors of a potential WTO dispute resolution case against China, they became a certainty on March 13, 2012, with the case brought by the U.S., Europe, and Japan. Note that we consider any information processing due to rumors prior to the event date as occurring during our "pre-WTO" period. Such speculation would make it more difficult to obtain statistically significant results in our "post-WTO" period. Therefore, we interpret our derived results as conservative (see Online Appendix A.9 for a detailed overview). Other event dates during the dispute resolution case could be used to determine the importance of announcements. However, this could result in confounding events and make it virtually impossible to isolate the "importance" of any particular event during the dispute resolution case.

4. Results

4.1 Descriptive Statistics

Because the REEs are not of equal economic importance (Long et al., 2010), we focus in this article on the four elements with the highest consumption: Cerium oxide (Ce), lanthanum oxide (La), neodymium oxide (Nd), and yttrium oxide (Y). 2008 consumption statistics show that these four elements jointly account for about 90% of total REE consumption: Ce at 42,220 metric tons/32.94%, La at 38,665 metric tons/30.16%, Nd at 22,868 metric tons/17.84%, and Y at 11,610 metric tons/9.06% (see Goonan, 2011).

Table 1 shows the descriptive statistics for the time series of REE oxide prices; Table A.3 in the appendix shows the descriptive statistics of the log returns.[15] Panel A ($PreWTO$) presents the descriptive statistics for the 2.5-year period prior to the launch of the WTO dispute resolution case on March 13, 2012; panel B ($PostWTO$) presents the descriptive statistics for the 2.5-year period after the WTO event. Examining the mean of the USD/kg prices, it seems obvious that the dual pricing system leads to higher export prices (FOB prices) for the same elements than the China domestic price. This is further underlined by the fact that the median FOB price is higher than the domestic price for all elements, which holds for both the period before and the period after the launch of the WTO trial on March 13, 2012. However, in the post-WTO period, both FOB and China prices substantially decrease. Furthermore, the dispersion of prices as measured by their standard deviations decreases as well. More importantly, the differences between the domestic and export prices decreased after the initiation of the dispute resolution case against China.

Regarding the descriptive statistics, all the time series of REE prices exhibit positive skewness, except the FOB price of yttrium oxide in the pre-WTO period. Moreover, all the time

[15] All REE log returns exhibit positive skewness except the FOB price of yttrium oxide. Regarding the fourth moment, the log return series are generally leptokurtic.

series of prices are platykurtic before the WTO event, but the FOB prices of neodymium and yttrium oxides and the China price of neodymium oxide become leptokurtic in the post-WTO period.

– Please insert Table 1 about here –

4.2 Results of the Variance Ratio Tests

The goal in this subsection is to provide the first evidence about the effect of the WTO event on the price behavior of the four most important REEs (cerium, lanthanum, neodymium, and yttrium). We calculate variance ratio tests for the series of REE prices based on weekly prices, as in Lo and MacKinlay (1988) and Liu and He (1991) (see panel A, Table 2).[16] We report the calculated variance ratios $\overline{M}_r(q) + 1$ in the main rows, and the test statistics $z^*(q)$ in parentheses for FOB and China prices.[17] We then compare the results from the pre-WTO period (thirty months before the announcement) with those from the post-WTO period (thirty months after the announcement).

We find that, in the pre-WTO period, the random walk hypothesis must be rejected for all REE oxide prices for all aggregation levels at least at a 5% significance level. Because of using the heteroscedasticity-robust test statistic, our results are robust to time-varying variances. As Equation (1) implies, the variance ratio for an aggregate value of $q = 2$ in the first column of Table 2 should be equal to $\rho(1) + 1$, i.e., the first-order autocorrelation coefficient of the one-week returns plus 1. Accordingly, the value of $\overline{M}_r(q) + 1 = 1.401$ for FOB prices of lanthanum prior to the WTO event indicates that the autocorrelation of order 1 for weekly returns is about 40%. This effect becomes more pronounced when increasing the aggregation value q.

[16] We also report results based on monthly REE prices as a robustness check at the end of subsection 4.2 (see, e.g., Poterba and Summers, 1988, for different levels of data granularity).
[17] Besides using Lo and MacKinlay's (1988) heteroscedasticity-robust test statistic, we reran the analysis using Kim's (2006) wild bootstrap with 10,000 replications. The results are virtually identical and are available from the authors upon request.

Interestingly, and in line with our intuition, the variance ratios for almost all REE prices and aggregation levels decrease considerably after the launch of the WTO dispute resolution case (except for FOB prices for cerium and neodymium oxide at an aggregation value of $q = 2$). For higher aggregation values, the variance ratios for all REEs dramatically decrease, to about 67% (see "Rel. Dif." in panel A of Table 2). However, despite this substantial drop, it is not sufficient for the price behavior to be regarded as a random walk (expect for China prices for lanthanum). Because we allow the volatility to differ between the two subsamples, and because we use the heteroscedasticity-robust test statistic, these results are valid for periods of both increasing and decreasing prices, i.e., during both bull and bear markets.

To strengthen our argument, we use the non-parametric rank-based variance ratio test suggested by Wright (2000). The results are in panel B of Table 2 and underline our previous findings: After the announcement of the commencement of the WTO dispute resolution case by the U.S., the EU, and Japan, the variance ratios dropped by as much as 75%. Nevertheless, we note that most series are not martingale differences, so the market for REEs remains inefficient in that sense.

In summary, our results are consistent with the notion that, after the WTO event, REE price behavior ultimately changed, resulting in lower variance ratios. This is indicative of changes that enhanced the pricing efficiency of REEs. As a robustness check, we also used a monthly sampling interval for the variance ratio tests.[18] The results were virtually identical to the weekly sampling interval. However, the results are indicative only, and we complement them with structural break tests in the following subsection.

– Please insert Table 2 about here –

[18] Lo and MacKinlay (1988) use a four-week sampling period for their sensitivity analysis. Similarly, Poterba and Summers (1988) and Ayadi and Pyun (1994) use monthly data for their variance ratio tests. Tables showing monthly time series results are available from the authors upon request.

4.3 Results of the Structural Break Tests

In this subsection, we aim to determine whether the WTO event is related to a change in REE prices, without exogenously setting the date as we did in the previous analyses. If the WTO event has an impact on the price behavior of REEs, we expect to locate a structural break around the event, which occurred on March 13, 2012 (calendar week 11 in 2012).

We follow Awokuse et al. (2009) and use Bai and Perron's (1998, 2003a, 2003b) test for estimating the breaks in REE prices. Based on the sequential procedure, the dates for the first break are generally located in the second half of 2010/first half of 2011 (the dates for the first breakpoint span calendar week 29-2010 to calendar week 10-2011).[19] More interestingly, we find breakpoints that are relatively close to the announcement of the WTO event (the dates for the second breakpoint span calendar week 9-2012 to calendar week 31-2012), except for the China price of lanthanum and the FOB price of yttrium (see Table 3). Since it seems unrealistic that REE price behavior would change immediately, we believe the break dates are within a reasonable post-announcement date time frame. We interpret this as further support for the notion that the launch of the WTO dispute resolution case against China is related to changes in REE price behavior.

– Please insert Table 3 about here –

In summary, we find multiple streams of evidence that the price-generating process of REEs changed in response to the launch of the WTO trial, and experienced a structural break in such a way that the variance ratios of REE prices shrank thereafter. However, we have not yet determined whether the changes are economically meaningful. We address this in the next two subsections by analyzing the impact on stock price informativeness and on derivatives pricing.

[19] Note that the break dates identified by Bai and Perron's (2003a, 2003b) sup$F(\ell + 1|\ell)$ test in Table 3 are not in chronological order but are as identified by the testing procedure.

4.4 Results for Stock Price Informativeness

This subsection provides evidence about the differences in stock price informativeness before and after the WTO event. First, we present univariate evidence only for companies in the REE industry. Second, we use a multivariate DiD setting to illustrate the change in stock price informativeness of REE companies compared to other companies in the same industry. Finally, we conduct several robustness checks of the results for alternative estimation and REE industry identification strategies.

The average stock price synchronicity measure for all firms in the sample is -0.369 (see Table 4), which is similar to the -0.232 reported by Gul et al. (2010), but much larger than the -1.742 reported by Piotroski and Roulstone (2004) for U.S. firms.[20] The rather small difference from Gul et al. (2010), who also analyze Chinese companies, presumably stems from the deviation in time periods, and the restriction to only the three industries possibly related to REEs in our sample. We attribute the difference from Piotroski and Roulstone's (2004) results to differences in stock price synchronicity in emerging versus mature stock markets.

– Please insert Table 4 about here –

In a univariate setting, we compare the stock price synchronicity measure of REE companies before and after the WTO event, as calculated in Equation (7). We test whether it significantly decreased after the WTO event by using *t*-tests. A decrease in stock price synchronicity implies that the REE companies co-move to a lesser extent with market- and industrywide information, and therefore exhibit more firm-specific information. This notion is

[20] Comparing the means for the other firm-specific characteristics in our sample with the means reported in Gul et al. (2010), we find that only $Volume$ and M/B appear to differ. The mean $Volume$ in our sample of 3.560 is higher than the 1.245 found by Gul et al. (2010), which can be explained by the fact that trading volume in Chinese stocks increased over the years. The M/B in our sample is also higher, which could be attributable to the use of different sample industries and different time periods in the analysis. Furthermore, Gul et al. (2010) divide by total net assets instead of book value of equity to calculate M/B.

statistically supported by the evidence in Table 5, which shows that the stock price synchronicity of REE companies decreased after the WTO event.

We observe the strongest support in terms of statistical significance when we compare the 2.5 years before the WTO event with the 2.5 years afterward, where the *t*-statistics range from -1.402 to -9.743 (see column [Pre; Post] in Table 5). The “weakest” support is provided for the shortest time period (one year) after the WTO event (see column [-1; +1] in Table 5). This makes sense intuitively, because we cannot expect stock price synchronicity to change with a similar speed as stock returns, which can incorporate new information instantaneously. Stock price synchronicity is measured only once a year in a regression framework, which means that, if synchronicity is changing with a certain delay to the WTO event, the change will be less pronounced in the subsequent year. This will be reflected by lower *t*-statistics (in absolute terms). Thus, the change in synchronicity is expected to be slow or creeping. This view is supported by the fact that the *t*-statistics increase almost monotonically the more distant the post-WTO period is from the WTO event (compare columns [-1; +1], [-1; +2], and [Pre; Post] from top to down in Table 5).

Note that these results are less pronounced when using the REE indices based on China prices instead of FOB prices to identify companies in the REE industry. They are also statistically insignificant for the examination one year after the WTO event (see column [-1; +1] in Table 5). We argue that this may be attributable to a less precise selection of companies in the REE industry, because, for Chinese companies, FOB prices are more relevant for stock price movements. Consider the following explanatory examples: As FOB prices are always higher than China prices, an increase in FOB prices would be even more important for REE mining companies that export REEs and could thus generate additional revenue. Moreover, REE user companies can purchase REEs at cheaper China prices than outside competitors, which gives their end-products a competitive advantage. Both effects, however, are captured less than completely in the selection process, because the REE indices are based on China prices.

In summary, we interpret our univariate results as strong support for the notion that stock price synchronicity (informativeness) decreased (increased) significantly after the WTO event. The results are arguably stronger when FOB-based REE indices for REE company identification are used, and when the comparison period after the WTO event is further from the event to allow the synchronicity measure to better unfold.

– Please insert Table 5 about here –

Despite this univariate evidence, we have not yet controlled for other alternative explanations, such as a general regressive tendency of stock price synchronicity in the Chinese market due, e.g., to higher transparency or enforcement, or for firm characteristics such as company size. In the multivariate DiD analyses, we aim to isolate the effect of changes in stock price synchronicity for *REECs* in response to the WTO event, and to then control for other potentially influencing factors.

Table 7 shows the regression results for OLS regressions that explain stock price synchronicity (including firm-specific controls and time fixed effects).[21] We are primarily interested in the coefficients for the variables $PostWTO(\tau_2) \times REEC$ and $PreWTO(\tau_1) \times REEC$ (the DiD coefficients).[22] The interaction between the pre-WTO dummy variable and $REEC$ is statistically significantly positive for all specifications at the 5% and 1% levels (*t*-statistics range from 2.17 to 4.51, for 10, 15, and 50 $REECs$, equally or usage-weighted REE index based on FOB prices). This means that stock price synchronicity for *REECs* before the WTO event was

[21] Table 6 shows the correlation matrix for all variables used in the multivariate analyses.

[22] The multivariate results correspond to the univariate evidence. To clarify, the corresponding univariate results are framed with dotted lines in Table 5. Specifications (1)-(3) are the multivariate pendants of the univariate results from Table 5 in panel (A), Equally Weighted FOB REE Index, for a selection period of two years before the WTO event and a [Pre; Post] comparison. Similarly, specifications (4)-(6) are the multivariate pendants of the univariate results from Table 5 in panel (C), Usage-Weighted FOB REE Index, for a selection period of two years before the WTO event and a [Pre; Post] comparison.

statistically higher than for non-*REECs* in the same industries. This effect could be attributed to the previously discussed aggressive governmental regulation that *REECs* face, for example, the setting of export quotas with the so-called MOFCOM announcements. However, the coefficient on the interaction $PostWTO(\tau_2) \times REEC$ is negative and statistically significant at least on a 10% level in all specifications (t-statistics range from -1.88 to -3.33). This suggests that the WTO event caused a significant drop (increase) in stock price synchronicity (informativeness) for $REECs$ relative to non-*REECs* in the same industries.

The signs of our controlling variables are in line with those reported in related works. We find a statistically significantly positive relationship between firm size ($Size$) and stock price synchronicity on a 1% level for all specifications, which means that larger companies in China are tied more to market and industry developments than comparable smaller companies (see Chan and Hameed, 2006; and Gul et al., 2010). Furthermore, M/B is statistically negative at a 10% level for all specifications, implying that companies with more growth options convey more firm-specific information and thus have higher stock price informativeness (see Gul et al., 2010).

Finally, our control variable for government ownership ($Top\,Gov$) reveals that government control is negatively correlated with proper governance and with weaker shareholder protection (managerial entrenchment), which leads to higher stock price synchronicity (see Shleifer and Vishny, 1994; and Gul et al., 2010). This effect is also present in our results by a positive and statistically significant coefficient at a 1% level in all specifications.

– Please insert Tables 6 and 7 about here –

We conclude this section with a few robustness checks. First, we apply alternative estimation strategies to test for robustness of the OLS regressions in Table 7. One concern with using OLS regressions is that the stock price synchronicity measure is rather sticky over time and can thus exhibit autocorrelation. To address this concern, we apply two-stage least squares regressions with Newey-West (1987) standard errors and lag(1) (see panel A of Table 8, which replicates our main model from Table 7). Note that the signs and the *t*-statistics for both DiD coefficients are comparable to those in Table 7, suggesting that potential autocorrelation in the stock price synchronicity measure does not affect our results.[23]

Next, we use a panel regression random effects estimator with robust standard errors (see panel B of Table 8). The results are highly similar in terms of magnitude to Table 7 and to panel A of Table 8, although the power of our tests decreases. We again find negative coefficients for the interaction $PostWTO(\tau_2) \times REEC$ and positive coefficients for $PreWTO(\tau_1) \times REEC$, but the latter is no longer statistically significant when we restrict the number of companies in the REE industry to ten. However, the negative coefficient on $PostWTO(\tau_2) \times REEC$ remains statistically significant at least at the 10% level (except for ten *REECs* and when the usage-weighted REE index based on FOB prices serves as the selection criteria). Our basic statement, that $\beta_1 > \beta_2$, therefore remains unchanged, and we are confident that the alternative estimation strategies do not alter our previous evidence that stock price synchronicity of *REECs* declined significantly after the WTO event.[24]

– Please insert Table 8 about here –

Second, we test the robustness of our results with regard to the selection criteria to identify *REECs* by using an REE index based on China prices rather than FOB prices (see Table A.4 in the appendix). The univariate evidence already shows that the results from using China prices are much weaker, presumably following the previous argument that FOB prices are more relevant for stock price formation. Therefore, we expect this property to be reflected in the multivariate results as well. However, we again find a statistically significant negative coefficient

[23] In unreported results, we changed the lags to two and three, and our results are virtually the same. The corresponding table is available from the authors upon request.

[24] Note that we do not apply Arellano-Bond dynamic panel estimators because it would be in opposition to the DiD approach. See, for example, Jacob et al. (2016) for a similar argumentation in a related context.

(at least at a 5% level) for the interaction $PostWTO(\tau_2) \times REEC$ identifying fifty *REECs*, whether we apply OLS, Newey-West standard errors, or a random effects estimator. If we consider only portfolios of 25 or 10 companies, both DiD coefficients are largely insignificant, but their predicted signs remain.

Despite the insignificant results for the smaller *REEC* portfolios, we interpret this robustness check as further support for our prediction that the stock price synchronicity of *REECs* declined after the WTO event for three reasons: 1) We argued that China REE prices are less precise than FOB prices, a fact that we expect to translate into decreased power for our test statistics, 2) We consider using only 10 *REECs* as the strictest test. This is because some companies that should obviously be classified as *REECs* have presumably been attributed to the group of non-*REECs* (as Müller et al., 2016, for example, already identified 14 REE mining and refining companies). This leads to reason 3), that, with an increase in *REECs*, the *t*-statistics for the interaction $PostWTO(\tau_2) \times REEC$ increase almost monotonically, supporting the idea that more than 10 or 25 companies are active in the REE industry.

Third, synchronicity assumes equal effects on the upside and the downside. Thus, there could be an effect from, e.g., a distinct bull or bear market. To address this concern, we already included year fixed effects in the regression that capture potentially asymmetric effects in a particular year. In unreported results, we also narrowed it to include “half-year” fixed effects, because bull and bear trends do not necessarily span an entire year. Our results remain virtually unchanged.

Fourth, there could be further control variables that are related to stock price synchronicity for which we have not controlled. We extend the set of control variables by the following: a) auditors (see Chen et al., 2013), b) political connections (see Xu et al., 2015; Wu et al., 2016; Schweizer et al., 2017; and Schweizer et al., 2018), c) split share structure reform (see

Cumming and Hou, 2014; and Cumming et al., 2016), d) financial analysts (see Chen et al., 2016), and e) corporate governance and ownership (see Bai et al., 2004; Chen et al., 2006; Liu and Lu, 2007; and Schweizer et al., 2017, 2018).

For media coverage, we use the following indirect measures, which are strongly correlated with actual media coverage (see Cumming et al., 2016): measured size, percentage change in sales growth over the last fiscal year, ownership concentration in the last fiscal year, and executive ownership.[25] After re-running the same regression including the additional control variables, we find that the magnitude of our main coefficient of interest, $PostWTO(\tau_2) \times REEC$, does not change much in the various specifications (see Tables A.5 and A.6). In absolute terms, the t-statistics are higher for some specifications and lower for others, but there is no pattern of consistent reduction in the t-statistics or the magnitude of the coefficient for our main variables due to the inclusion of the new variables. However, the inclusion of such a large number of control variables increased the Variance Inflation Factors, because some of the new variables are correlated with previous and newly added control variables, which was not the case before and could be interpreted as a sign of multicollinearity. Furthermore, we find that the inclusion of the new control variables reduces the adjusted R2 by about 1 percentage point. In summary, we find that our main results remain stable and robust after the inclusion of the new control variables.

In summary, we present robust evidence that the stock price informativeness of *REECs* increased compared to non-*REECs* after the March 13, 2012, announcement of the U.S., EU, and Japan bringing a dispute resolution case against China to the World Trade Organization (WTO). We interpret this result to mean that, after the beginning of the WTO trial, the Chinese government's influence on the REE market (by, e.g., MOFCOM announcements) decreased. The

[25] A direct measure for media coverage, such as that used in Cumming et al. (2016), is the cleanest way to control for this potential influence. Given that the indirect measures had no impact on the main results or the explanatory power as to R-square, it seems unlikely that a direct measure of media coverage would have an impact on our primary results.

MOFCOM announcements were typically used biannually to set the maximum amount of REEs for export, and they were generally thought to increase the value of Chinese stocks. The export quotas are generally not known to shareholders; thus, stock prices react in response to the announcement. This provides an opportunity for company insiders to trade on a private information advantage, especially for companies with high levels of state ownership and close ties to the government. Consequently, this should be reflected in higher (lower) stock price synchronicity (informativeness) of REE companies. It seems that, after the beginning of the WTO trial, the Chinese government was more reluctant to set “unexpected” export quotas, and did tend to intervene less in the REE market. We consider this as an explanation for the increase in stock price informativeness of REE companies after the WTO event.

4.5 Derivatives Pricing

This subsection provides empirical evidence for changes in the option price differences among various option pricing models before and after the WTO event. If the WTO event caused less intervention by the Chinese government, such as the removal of REE export quotas, we expect that the price-generating process for REEs will become less “erratic,” which will be echoed by lower pricing differences between models.[26]

For the “static” comparison in Table 9, we observe that the absolute pricing differences (PD) for all REEs between all option pricing models before the WTO event ($PreWTO$) are considerably higher than after ($PostWTO$) the event (except for one case, for Ce between the

[26] Commodity options are commonly based on related futures contracts instead of on spot prices, such as the underlying. The dynamics of futures prices are determined by, not only the spot price, but also by the convenience yield (or “availability” premium) and the interest rate. The stochastic process modeling of the convenience yield has an impact on the option value if the convenience yield is modeled with, e.g., constant correlation (see, for example, Gibson and Schwartz, 1990), non-constant correlation to the spot price (e.g., Routledge et al., 2000), or the spot price and the interest rate (Casassus and Collin-Dufresne, 2005). In the following, we base our calculations on only the spot price dynamics, for practical reasons. As mentioned in the introduction, no futures contracts on REEs are yet available, which would allow us to specify the dynamics of the convenience yield and the correlation with the spot price and the interest rate. Therefore, any exogenous and presumably inaccurate assumptions will most likely not improve the quality of our results. Furthermore, our “difference-in-differences” approach can cope with this quality if the convenience yield process does not change over time.

Black-Scholes (1973) and Duan (1995) models). For example, the aggregate absolute pricing difference for La for call options between the Black-Scholes (1973) and Duan (1995) models, given all *strike price* and *time to maturity* combinations, is 106.24 before the WTO event, and 26.39 afterward. This represents a 79.84 decrease, or, in relative terms, an approximately 75.16% decrease (see column (2) of Tables 9 and 10 for detailed calculations corresponding to the values framed by the dotted lines and shaded in grey for lanthanum in Table 9). The total absolute pricing difference for all option pricing models for La call options is 230.76 before the WTO event, and 55.11 afterward. This represents a 175.65 decrease, or a similar relative decrease of 76.12% (see column (1) of Tables 9 and 10). The decrease for the total absolute pricing difference for call options for cerium, neodymium, and yttrium range from 63% to 79% (see column (1) of Table 9). For put options, the effect is stronger for all REEs.

In Fig. 3, we present the above-described evidence from a graphical representation point of view for call options, based on lanthanum. Panel A presents the absolute call option price differences based on the Black-Scholes (1973) and Duan (1995) models for various strike price and time to maturity combinations. Column (1) shows the call option price differences before the WTO event ($PreWTO$), and column (2) shows them afterward ($PostWTO$). Both figures illustrate similar behavior, namely, that the absolute call option price differences are larger for longer maturities and for at-the-money prices (e.g., the strike price is equal to the current lanthanum price). However, absolute pricing differences are lower in the $PostWTO$ period compared to the period beforehand. This is reflected by the negative differences between absolute pricing differences between the $Post$- and $PreWTO$ periods (see column (3) in Fig. 3). We can interpret this result similarly as the previously mentioned 175.65 pricing decrease for lanthanum in Table 9, which indicated a reduction in pricing differences after the WTO event.

Qualitatively similar figures are presented for the absolute call option price differences between Black-Scholes (1973) and Heston and Nandi (2000) in panel B. The results in panels A and B reveal that absolute pricing differences are consistently lower after the WTO event. However, for panel C, where the absolute call option price differences between Duan (1995) and Heston and Nandi (2000) are compared, we do not find a similar relationship between the pricing difference and different times to maturity and strike prices. From a visual inspection, it seems that the pricing differences after the WTO event are smaller than those for $PreWTO$ and are of lower magnitude, which can also be seen in column (4) in Table 9.

– Please insert Tables 9 and 10 and Figure 3 about here –

From the "static" DiD comparison, we find strong support for the notion that the pricing differences among the different option pricing models are substantially lower after the WTO event than before it ($PreWTO$ period). For the "static" DiD between the three option pricing models we find that pricing differences for call options for the four REEs decrease between 63.14% (neodymium) and 79.56% (yttrium) and for put options between 83.59% (neodymium) and 91.95% (yttrium). This decrease between the option pricing models is quite large and economically significant, especially for put options where the pricing difference can be marginalized for some REEs after the WTO dispute resolution case commences. This underpins our argumentation that the model risk for pricing derivatives on REEs decreased in response to the WTO event. However, as outlined in the methodology section, the static approach is not suitable from which to draw conclusions about statistical inference. It also focuses solely on option price calculations for two particular days.

To overcome these issues, we use a *consecutive* "triple-difference" (DDD) methodology that can rule out constant unobserved differences among REE pricing behavior during the Pre- and $PostWTO$ periods that may bias our results. The intuition is similar to the static DiD approach used previously. If the WTO event had a *positive* impact on the price behavior of REEs so that it reduced option pricing model uncertainty, we would expect to find a larger drop in option price differences among models during the "treatment" and "control" periods, implying $\overline{DiD}_k(PreWTO - PostWTO) > \overline{DiD}_k(PreWTO - PrePreWTO)$, and resulting in a positive DDD on average. This is consistently observed for all possible *strike prices*, *times to maturity*, and option model comparisons for lanthanum, neodymium, and yttrium (see Table 11).[27] These differences result in highly statistically significant positive DDDs, with t-statistics ranging from 2.84 to 20.62. However, similarly to the "static" DiD, we fail to find a consistent decrease in option price differences among all models after the WTO event for cerium. We can summarize the results for our comparison of price differences for the Black-Scholes (1973) and Duan (1995) models as inconsistent. But the comparisons among Black-Scholes (1973), Duan (1995), and Heston and Nandi (2000) indicate that the pricing differences decreased significantly after the WTO event (see again Table 11). For example, for lanthanum, we find that the option model pricing difference drops on average between 30% and 42% during the event period relative to the period before the WTO event. This drop is dependent on the respective comparison of the option pricing models, which is arguably economically meaningful.

In summary, we find compelling evidence from the *static* DiD and from the *consecutive* DDD approach that pricing differences among different option pricing models decreased significantly after the WTO event, not only statistically but also economically. We interpret this to mean that the influence of the WTO on the Chinese government triggered somewhat fewer interventions on the REE market. This is reflected afterward in less erratic price behavior of REEs, which we expect to facilitate the introduction of derivatives exchanges on REEs.

– Please insert Table 11 about here –

[27] For the sake of clarity, the following presents results only for call options, because results for put options are highly similar. The table is available from the authors upon request.

5. Conclusion

Our paper contributes to the growing literature on the economic implications of dispute resolution cases brought to the WTO. The literature thus far has focused mainly on the stock price reactions of affected companies in a classical event study setting (see Liebman, 2006; Liebman and Tomlin, 2007, 2008, 2015; and Desai and Hines Jr., 2008). However, we choose a different approach, and use the commencement of the WTO dispute resolution case on March 13, 2012 as a natural experiment in order to highlight its effects from four different angles.

First, we use variance ratio tests to show that price formation on the REE market is not efficient, in the sense that the prices are random processes. However, after the initiation of the WTO dispute resolution case, the magnitude of the variance ratios is considerably lower. Hence, this means that even the *announcement* of a WTO dispute resolution case seems to have "efficiency-enhancing effects" with respect to the price formation of REEs. This is in line with the notion of using variance ratio tests to study the potential enhancement of market efficiency in other contexts. Examples include the liberalization of investment restrictions on B shares in the Chinese stock market (see Hung, 2009), the opening of Asian stock markets and the unprecedented inflow of funds into these markets after the Asian financial crisis in 1997 (see Hoque et al., 2007), and the analysis of crude oil market efficiency over time (see Tabak and Cajueiro, 2007).

Second, we use structural change tests and find that the price processes of REEs exhibit a significant change in dynamics around the announcement date of the dispute resolution case. However, there may have been some foreshadowing, or anticipation, of the launch of a WTO dispute resolution case about REEs from prior WTO rulings against China in other commodity-related trials in 2009 (see Bond and Trachtman, 2016). Nevertheless, we find that a WTO trial announcement has economically meaningful effects for REEs and related public companies. If investors had foreseen the start of the WTO resolution case against China, logically, we would have expected them to rationally adjust their expectations accordingly. Thus, we would not have been able to document significant effects. This is also supported by the Google Trends Index for REEs, which had shown a consistent decline in interest in REEs prior to 2010, followed by a huge spike at the time of the WTO trial announcement (see Fig. A.2 in the appendix and Choi and Varian, 2009, 2012, for further insights into Google Trends).

Third, we study the impact of the WTO event on the information content of stock prices and find that stock price informativeness significantly increases for companies that are active in the REE industry relative to firms from other industries. Because the Chinese government ultimately presumably ceased intervening in the REE market at least to a certain extent due to the start of the WTO trial, firm-specific information became more important than marketwide information.

Moreover, as new trading venues for spot and futures markets emerge, such as the Baotou Rare Earth Products Exchange, and the potential trading of derivatives on these elements on the Shanghai Futures Exchange gears up (Shen, 2014), investors and regulators need to understand the price processes of REEs more deeply. Accordingly, fourth, we compare the results of three different option valuation models, i.e., Black-Scholes (1973), Duan (1995), and Heston and Nandi (2000). Note that we are not interested in determining the best option pricing model in order to minimize pricing errors (recall that the introduction of REE derivatives has not materialized yet). Rather, we compare pricing differences among the three models before and after the launch of the WTO case. We show that the differences among prices calculated from the three frameworks are much larger during the pre-WTO period than after the launch of the WTO trial. We interpret this finding as another indication that the WTO event triggered important changes to China's REE policy.

Overall, we document compelling evidence that the announcement of a WTO dispute trial can influence the stock prices of potentially affected companies (Desai and Hines Jr., 2008), as well as induce governmental changes in existing policies. While previous research found the WTO's enforcement mechanism to be somewhat ineffective (see Liebman and Tomlin, 2015),[28] we show that even the announcement of a trial can trigger economically significant changes in the REE market.

Thus, we provide first evidence for the notion that governments accused of violating GATT react to the announcement with "measurable" policy changes even before an official WTO ruling. This implies that the launch of a WTO case itself helps to resolve uncertainty, while in other legal cases, such as class action lawsuits, uncertainty tends to increase after the commencement of a case. This has important implications for investors and policy makers alike: The WTO is far from being a paper tiger. In fact, faced with the potential suspension of concessions and retaliatory threats, the compliance rate is well above 80% (Bekhar, 2010). Thus, governments take into account the costs of dispute settlement, and hence take the rulings of the WTO seriously (see the theoretical bargaining game, as well as the empirical evidence presented in Sattler et al., 2014). Accordingly, investors of affected companies and industries should carefully monitor WTO trials in order to ensure they are positioned correctly.

Our study also serves as an impetus for further research. Once the planned introduction of options on REEs is realized, one could use step 4 in our Fig. 1 to analyze the impact of a derivatives market on the spot market in regard to volatility, bid-ask spread, and liquidity (see Conrad, 1989; Skinner, 1989; Detemple and Jorion, 1990; Kumar et al., 1998; and Danielsen et

[28] Note that even in the case of Byrd subsidies, as analyzed by Liebman and Tomlin (2015), the policy in question (Continued Dumping and Subsidy Offset Act, CDSOA), which was criticized by the WTO, was eventually ended. Moreover, firms in industries with higher export orientations exhibit positive share price reactions with respect to the WTO's ruling that the CDSOA policy is illegal. Likewise, the U.S. firms facing retaliation from U.S. trade partners exhibit negative share price reactions. Accordingly, the claim that the WTO's enforcement mechanism is ineffective seems somewhat misguided.

al., 2007). However, as derivatives are unavailable so far, policy makers should consider building strategic stockpiles as a way to secure access to rare earth elements and shield user industries from price peaks and volatility (Bailey Grasso, 2013; and Humphries, 2013). Research from other high-risk commodity markets such as copper provides some guidance in that respect. Athanasiou et al. (2008), for example, show how to analytically derive a solution for the computation of the storage capacity as well as for the cost of a potential buffer stock program.

Moreover, the bankruptcy of Molycorp highlights how important it is for the REE industry to have a fuller understanding of what drives REE prices. At its peak, approximately four years ago, Molycorp's stock was trading at USD $79.16. Before filing for Chapter 11 bankruptcy protection, it was down to about USD $0.36 (see McCarty and Casey, 2015; and Reuters, 2015). Therefore, future research could explore the impact of macroeconomic factors such as exchange rates, interest rates, and the business cycle on REE prices and on *REECs* (see Faff and Chan, 1998; Tufano, 1998; Boyer and Filion, 2007; Baur, 2014; and Gil-Alana and Yaya, 2014, for related research on the gold and oil markets).

Due to the often-extreme price movements of REEs, future research may also want to concentrate on volatility modeling and prediction (see Batten et al., 2010; and Haugom et al., 2014, for the oil and metal markets) as well as risk management approaches. Extreme Value Theory (EVT)-based procedures successfully employed in other commodity markets might serve as a starting point for risk management applications (see, for example, Marimoutou et al., 2009; Aloui et al., 2014; Koch, 2014; and Herrera et al., 2017). It is well known that failing to account for the fact that distributions of asset returns can substantially deviate from the normal due to fat tails can greatly underestimate risk measures such as value at risk (VaR). Risk measures that focus on the tails, such as expected shortfall, may hence be more suitable for REEs (see the descriptive statistics of REE returns in Table A.3).

The evidence concerning the WTO's enforcement power is far from being conclusive (see, for example, the evidence presented in Liebman and Tomlin, 2015). In order to foster our understanding of the economic importance of requests for consultations at the WTO, the evidence presented in this paper needs to be broadened. Accordingly, the effects on the claimant as well as the accused countries need to be investigated using a cross-section of cases. It may be true that the effects depend on which country files the complaint and which country is accused.

For example, while the U.S. is the dominant economy in terms of GDP, the export sector is far less important for its industry: In 1970, exports of goods and services were 5.55% of GDP; this figure grew to 11.89% in 2016. This contrasts with China (Germany), who were able to increase their share of exports to GDP from 2.49% (15.17%) in 1970 to 19.64% (46.12%) in 2016, respectively (see Fig. A.3 in the Online Appendix).

Moreover, while this paper analyzes in-depth the impact of the *launch* of a dispute resolution case at the WTO, it might be worthwhile to analyze the *different stages* from the request for consultations to the final release of the report of the Appellate Body and the authorization of retaliatory measures (as in Liebman and Tomlin, 2015) in a classical event study context. We leave these ideas for future research.

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Illustration of the Interactions Among Chinese REE Companies, the Chinese Government, Foreign REE Companies, and the WTO

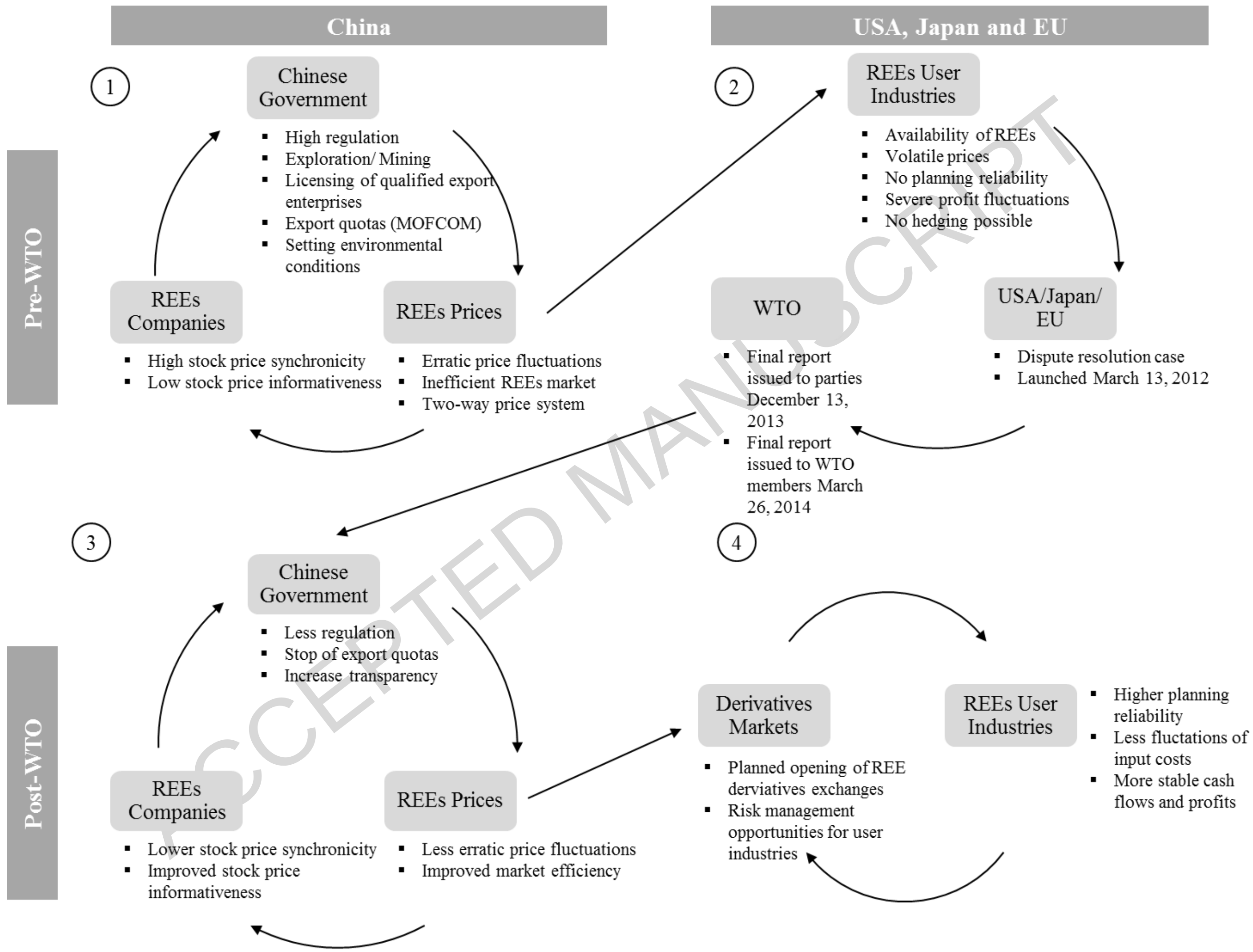

Fig. 1. This figure provides an illustration of the interactions among Chinese REE companies, governmental regulation, foreign REE companies and the WTO.

Illustration of the DiD and DDD Approaches

Panel A: Static DiD Approach

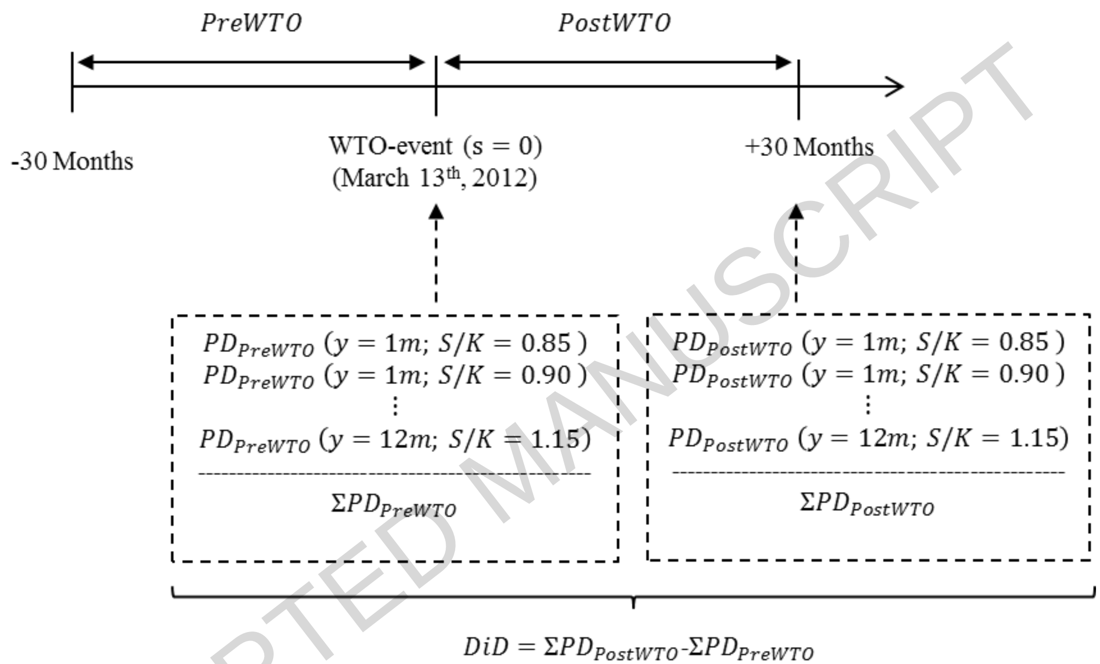

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Illustration of the DiD and DDD Approaches – *continued*

Panel B: Consecutive DDD Approach

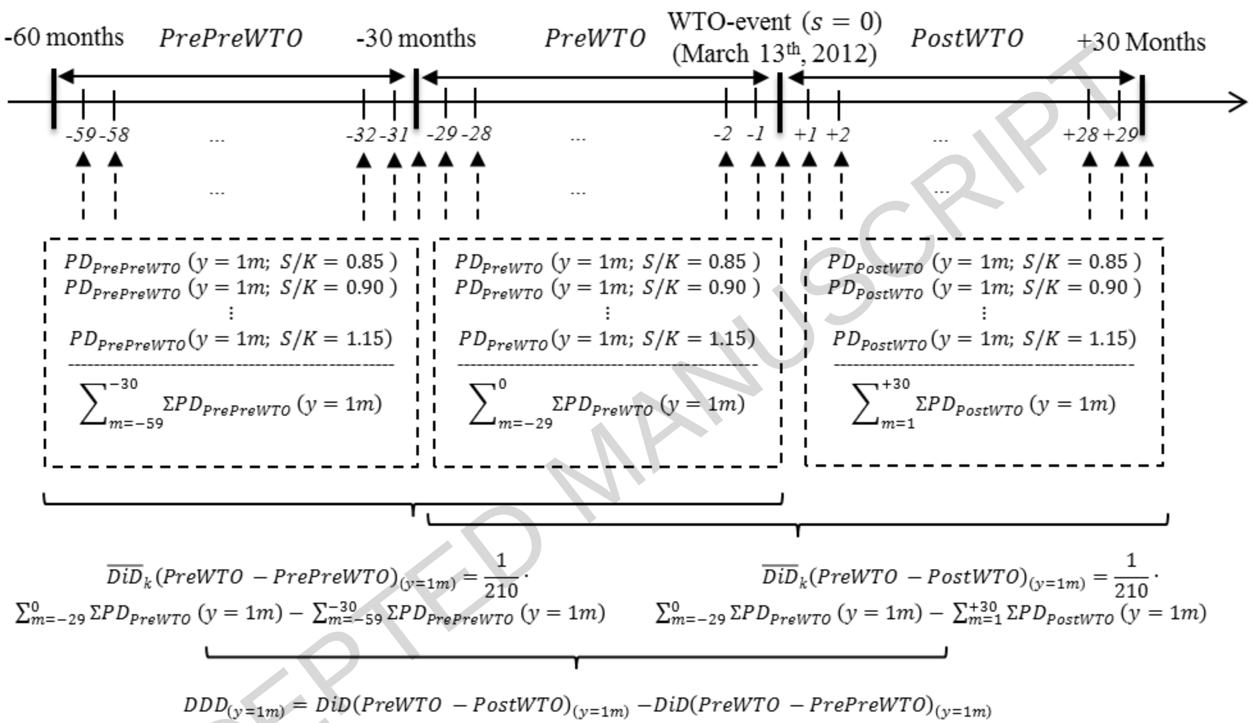
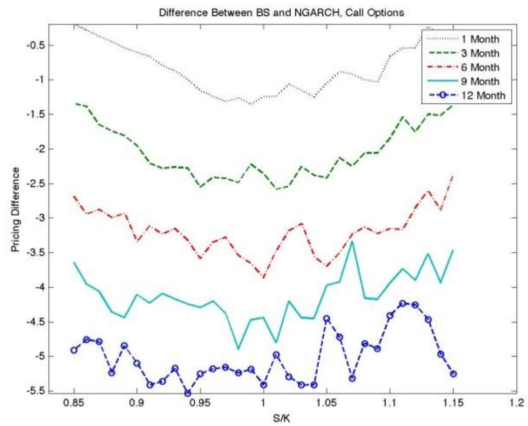


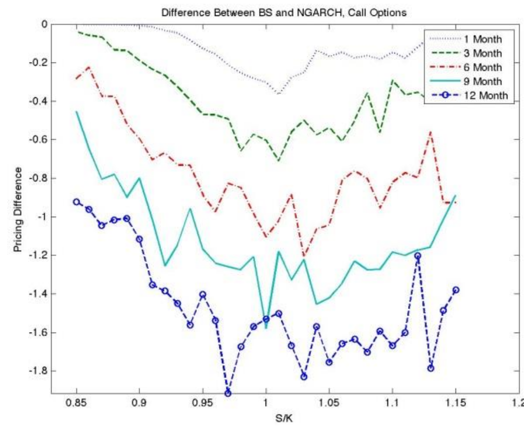
Fig. 2. This figure provides a visual approach toward evaluating the effect of the WTO trial on option pricing model uncertainty. Panel A shows the time line and windows for the static DiD approach; panel B outlines the idea of the DDD approach. The example shown is the methodology for one maturity only ($y = 1m$).

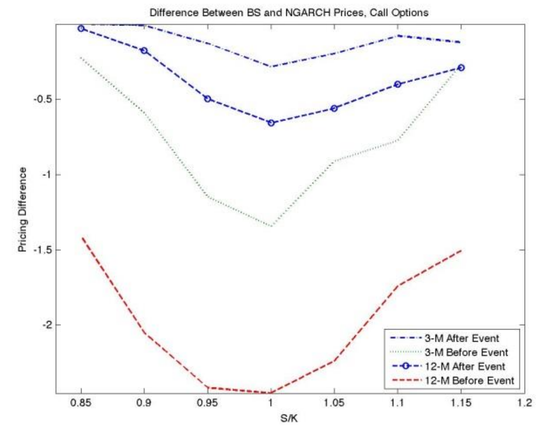
Option Price Differences between Option Pricing Models before and after the WTO Event (Lanthanum; Call Options)

Panel A: Call Option Price Differences between Black-Scholes (1973) and Duan (1995)

$PreWTO$ (1)	$PostWTO$ (2)	$PostWTO - PreWTO$ (3)

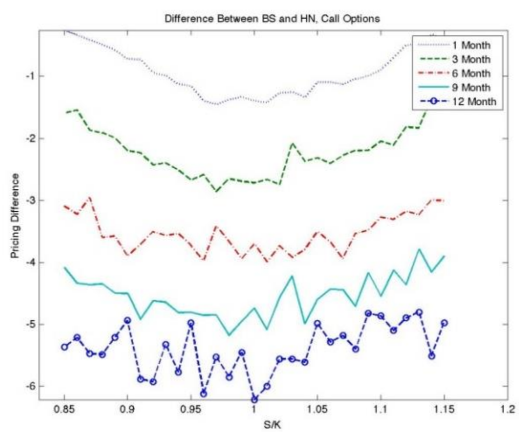


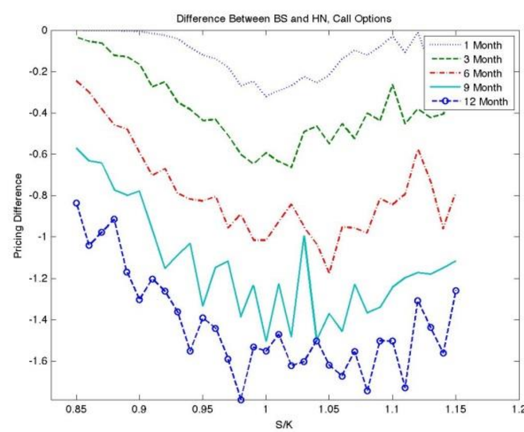


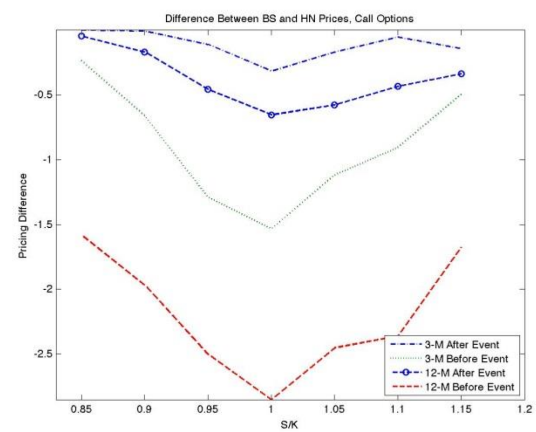


Panel B: Call Option Price Differences between Black-Scholes (1973) and Heston and Nandi (2000)

$PreWTO$ (1)	$PostWTO$ (2)	$PostWTO - PreWTO$ (3)







(continued)

Option Price Differences between Option Pricing Models before and after the WTO Event (Lanthanum; Call Options) – ***continued***

Panel C: Call Option Price Differences between Duan (1995) and Heston and Nandi (2000)

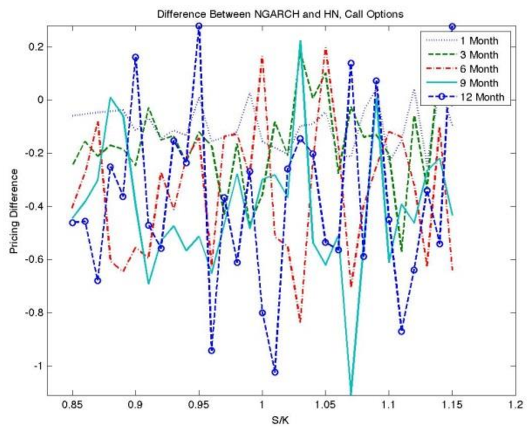
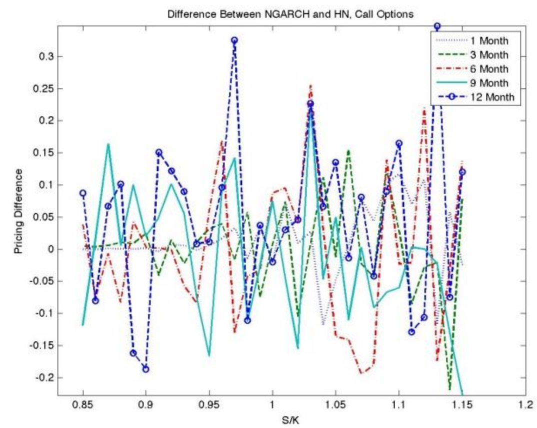
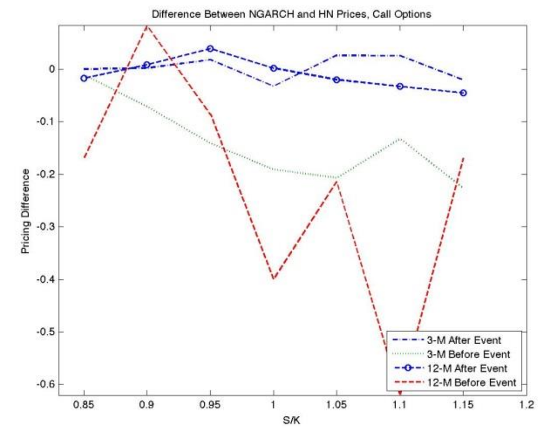
$PreWTO$ (1)	$PostWTO$ (2)	$PostWTO - PreWTO$ (3)
		

Fig. 3. This figure shows the pricing differences for different combinations of time to maturity and moneyness for lanthanum call options. Panel A (Panel B) [Pancel C] shows the differences between the option pricing models of Black and Scholes (1973) and Duan (1995) (Black and Scholes (1973) and Heston and Nandi (2000)) [Duan (1995) and Heston and Nandi (2000)], respectively.

Table 1
Descriptive Statistics of REE Prices – Weekly Base Observation Period

Element		Mean	Median	Std. Dev.	Skewness	Kurtosis	Min	Max
Panel A: (PreWTO)								
Lanthanum	FOB	52.57	49.95	45.54	0.65	2.24	4.05	140.05
	China	10.46	4.43	8.11	0.65	1.61	3.22	25.44
Cerium	FOB	51.47	41.50	47.02	0.78	2.50	3.35	150.55
	China	11.29	4.80	9.79	0.64	1.68	2.42	29.02
Neodymium	FOB	136.97	85.75	113.44	0.58	1.91	14.65	369.75
	China	76.35	39.01	65.60	1.05	2.94	12.45	232.91
Yttrium	FOB	120.89	125.00	45.63	-0.14	1.77	42.50	197.50
	China	23.55	7.61	21.96	0.88	2.36	5.78	77.16
Panel B: (PostWTO)								
Lanthanum	FOB	11.62	7.55	7.59	1.04	2.50	5.20	30.50
	China	6.62	4.85	3.56	0.79	2.32	2.61	14.44
Cerium	FOB	11.51	7.55	7.64	0.99	2.37	4.20	28.50
	China	6.60	4.85	3.63	0.73	2.23	2.36	14.68
Neodymium	FOB	81.58	74.50	21.23	1.76	5.55	59.50	146.00
	China	57.35	53.80	11.69	1.30	3.95	41.96	88.65
Yttrium	FOB	47.91	24.00	46.80	1.61	4.32	13.00	185.00
	China	14.80	12.09	7.06	0.94	2.87	6.62	33.96

This table presents the mean, median, standard deviation, skewness, kurtosis, minimum, and maximum of the time series of weekly REE FOB (foreign) and China (domestic) prices. Panel A ($PreWTO$) shows the prices for individual REEs for the period beginning thirty months prior to the WTO dispute resolution case on March 13, 2012. Due to data availability, the FOB prices for yttrium begin in August 2010. Panel B ($PostWTO$) shows the prices for the period up to thirty months after the WTO event. Price data comes from the Asian Metal database; all calculations based on USD/kg.

Table 2
Variance Ratio Tests – Weekly Base Observation Period

Panel A			Number q of base observations aggregated to form variance ratio			
Element			2	4	8	16
Lanthanum	FOB	$PreWTO$	1.401	1.997	2.816	4.112
			(3.024)***	(4.437)***	(6.036)***	(8.597)***
		$PostWTO$	1.274	1.641	1.998	2.897
			(2.011)**	(2.648)***	(2.926)***	(3.984)***
		Rel. Dif.	*-9.02%*	*-17.83%*	*-29.04%*	*-29.54%*
	China	$PreWTO$	1.485	2.283	3.520	5.208
			(2.467)**	(3.926)***	(5.877)***	(8.259)***
		$PostWTO$	1.227	1.659	2.477	3.778
			(2.611)	(4.048)	(5.738)*	(7.252)
		Rel. Dif.	*-17.36%*	*-27.34%*	*-29.61%*	*-27.45%*
Cerium	FOB	$PreWTO$	1.296	1.815	2.473	3.522
			(3.537)***	(4.232)***	(5.229)***	(7.307)***
		$PostWTO$	1.348	1.761	2.027	2.669
			(2.550)**	(3.171)***	(3.121)***	(3.730)***
		Rel. Dif.	*4.06%*	*-2.93%*	*-18.06%*	*-24.22%*
	China	$PreWTO$	1.450	2.301	3.744	5.602
			(3.304)***	(5.250)***	(7.506)***	(9.936)***
		$PostWTO$	1.112	1.453	1.955	2.375
			(1.190)	(2.523)**	(3.430)***	(3.365)***
		Rel. Dif.	*-23.34%*	*-36.83%*	*-47.78%*	*-57.60%*
Neodymium	FOB	$PreWTO$	1.336	2.006	2.936	4.323
			(6.526)***	(8.848)***	(10.875)***	(13.444)***
		$PostWTO$	1.405	1.811	2.211	2.257
			(1.788)*	(2.316)**	(2.835)***	(2.502)**
		Rel. Dif.	*5.19%*	*-9.722%*	*-24.674%*	*-47.791%*
	China	$PreWTO$	1.584	2.542	3.829	5.741
			(6.828)***	(10.070)***	(12.831)***	(15.996)***
		$PostWTO$	1.466	2.000	2.257	2.266
			(2.950)***	(3.984)***	(3.663)***	(2.822)***
		Rel. Dif.	*-7.44%*	*-21.33%*	*-41.06%*	*-60.53%*
Yttrium	FOB	$PreWTO$	1.470	2.198	3.166	4.554
			(4.394)***	(6.573)***	(8.637)***	(10.628)***
		$PostWTO$	1.345	1.786	2.176	3.211
			(2.976)***	(3.858)***	(4.075)***	(5.419)***
		Rel. Dif.	*-8.51%*	*-18.73%*	*-31.26%*	*-29.50%*
	China	$PreWTO$	1.359	2.123	3.166	4.645
			(3.044)***	(4.843)***	(6.665)***	(8.926)***
		$PostWTO$	1.151	1.511	1.760	1.548
			(1.211)	(2.412)**	(2.603)***	(1.358)
		Rel. Dif.	*-15.34%*	*-28.82%*	*-44.40%*	*-66.67%*

(continued)

Table 2 – *continued*

Panel B Element			Number q of base observations aggregated to form variance ratio 2	4	8	16
Lanthanum	FOB	$PreWTO$	1.434	2.215	3.455	5.480
			(10.250)***	(15.352)***	(19.623)***	(24.062)***
		$PostWTO$	1.182	1.434	1.653	1.947
			(2.085)**	(2.668)***	(2.536)***	(2.471)***
		Rel. Dif.	*-17.58%*	*-35.24%*	*-52.16%*	*-64.47%*
	China	$PreWTO$	1.330	1.984	3.279	5.643
			(7.400)***	(11.810)***	(17.297)***	(22.762)***
		$PostWTO$	1.140	1.400	1.818	2.016
			(1.614)	(2.459)**	(3.176)***	(2.652)***
		Rel. Dif.	*-14.22%*	*-29.42%*	*-44.56%*	*-64.27%*
Cerium	FOB	$PreWTO$	1.416	2.300	3.201	4.881
			(9.835)***	(13.896)***	(17.594)***	(20.845)***
		$PostWTO$	1.267	1.617	1.852	1.963
			(3.062)***	(3.789)***	(3.311)***	(2.515)***
		Rel. Dif.	*-10.56%*	*-22.98%*	*-42.13%*	*-59.79%*
	China	$PreWTO$	1.394	2.163	3.626	6.028
			(8.848)***	(13.957)***	(19.930)***	(25.646)***
		$PostWTO$	1.097	1.329	1.693	1.878
			(1.115)	(2.018)**	(2.692)***	(2.292)***
		Rel. Dif.	*-21.30%*	*-38.58%*	*-53.30%*	*-68.85%*
Neodymium	FOB	$PreWTO$	1.555	2.491	3.737	5.645
			(13.115)***	(18.838)***	(21.880)***	(24.948)***
		$PostWTO$	1.284	1.594	1.790	1.764
			(3.261)***	(3.650)***	(3.070)***	(1.994)**
		Rel. Dif.	*-17.43%*	*-35.98%*	*-52.10%*	*-68.75%*
	China	$PreWTO$	1.617	2.656	4.053	5.890
			(13.857)***	(19.872)***	(23.167)***	(24.940)***
		$PostWTO$	1.422	1.975	2.15	1.444
			(4.845)***	(5.987)***	(4.484)***	(1.158)
		Rel. Dif.	*-12.09%*	*-25.64%*	*-46.83%*	*-75.49%*
Yttrium	FOB	$PreWTO$	1.501	2.301	3.329	4.486
			(7.327)***	(10.172)***	(11.517)***	(11.586)***
		$PostWTO$	1.337	1.741	1.990	2.543
			(3.868)***	(4.553)***	(3.846)***	(4.026)***
		Rel. Dif.	*-10.94%*	*-24.31%*	*-40.21%*	*-43.32%*
	China	$PreWTO$	1.404	2.142	3.376	5.249
			(9.080)***	(13.704)***	(18.031)***	(21.671)***
		$PostWTO$	1.217	1.601	1.866	1.468
			(2.489)**	(3.691)***	(3.362)***	(1.220)
		Rel. Dif.	*-13.37%*	*-25.25%*	*-44.74%*	*-72.04%*

This table presents the results for the Lo and MacKinlay (1988) (panel A) and Wright (2000) (panel B) variance ratio tests of the random walk hypothesis. The variance ratios $\bar{M}_r(q)+1$ are presented in the main rows, and the heteroscedasticity-robust test statistic $z^*(q)$ is presented in brackets. Under the null hypothesis that REE prices follow a random walk, the value of the variance ratio is 1. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively. $PreWTO$ shows the results for the period beginning thirty months prior to the WTO event on March 13, 2012. $PostWTO$ shows results for the period up to thirty months after the WTO event. Due to data availability, FOB prices for Y begin in August 2010. Price data comes from the Asian Metal database; all calculations based on USD/kg. Rel. Dif. is the relative difference between the variance ratio before and after the launch of the WTO trial.

Table 3
Multiple Structural Change Test

Element		$\sup F(\ell+1\|\ell)$	Break Dates				
			$m=1$	$m=2$	$m=3$	$m=4$	$m=5$
Lanthanum	FOB	4	2010:30	2012:10	2007:41	2006:09	-
		(44.84)*					
	China	3	2011:10	2008:07	2012:33	-	-
		(50.38)*					
Cerium	FOB	3	2010:29	2012:09	2007:21	-	-
		(126.27)*					
	China	4	2011:06	2012:29	2009:36	2008:13	-
		(63.19)*					
Neodymium	FOB	3	2010:33	2012:18	2006:30	-	-
		(42.22)*					
	China	4	2011:05	2012:28	2006:29	2009:35	-
		(15.55)*					
Yttrium	FOB	5	2012:36	2011:12	2011:44	2013:16	2014:07
		(34.14)*					
	China	4	2011:08	2012:31	2007:27	2008:50	-
		(27.94)*					

This table presents the results of the Bai and Perron (1998, 2003a, 2003b) multiple structural change test allowing for serial correlation in the error terms and heterogeneous error distributions across breaks. We use the heteroscedasticity- and autocorrelation-consistent covariance matrix based on Andrews (1991) and Andrews and Monahan (1992) with a quadratic-spectral kernel and automatic bandwidth selection. Following Bai and Perron's (2003a, 2003b) recommendations, we select the number of breaks and the break dates using the sequential $\sup F(\ell+1|\ell)$ test. The number of breaks is shown in the main rows and the test statistic is shown in brackets, where * indicates statistical significance at the 5% level (see Bai and Perron, 2003a, 2003b). All results are based on weekly REE prices for the January 2, 2004-September 30, 2014 period. Due to data availability, the observation periods differ for some REEs: For FOB prices of La, Ce, and Nd, the time series begin in January 2004. The FOB prices for Y begin in August 2010. For China prices of Y, La, Ce, and Nd, the time series begin in January 2005. All time series end in September 2014. Price data comes from the Asian Metal database; all calculations based on USD/kg. Specification: Trimming $\varepsilon = 0.15$; maximum number of breaks $M = 5$.

Table 4
Descriptive Statistics for Stock Price Synchronicity and Firm Characteristics

Variable Name	#Obs	Mean	Standard Deviation	5th Pctl.	25th Pctl.	Median	75th Pctl.	95th Pctl.
$\boldsymbol{SYNCH}$	4,307	-0.369	0.638	-1.505	-0.736	-0.319	0.067	0.594
$\boldsymbol{Top\ Gov}$	4,307	0.093	0.178	0.000	0.000	0.000	0.078	0.514
$\boldsymbol{Volume}$	4,307	3.560	2.553	0.658	1.707	2.895	4.774	8.717
$\boldsymbol{Size}$	4,307	22.167	1.307	20.405	21.278	21.971	22.862	24.656
$\boldsymbol{Leverage}$	4,307	0.489	0.195	0.161	0.351	0.493	0.628	0.786
$\boldsymbol{Std}(\boldsymbol{RoA}))$	4,307	0.036	0.063	0.005	0.013	0.023	0.041	0.100
$\boldsymbol{M/B}$	4,307	2.988	23.764	0.835	1.529	2.397	3.792	7.937
$\boldsymbol{Ind\_Num}$	4,307	6.154	0.835	4.691	6.510	6.532	6.532	6.532
$\boldsymbol{Ind\_Size}$	4,307	29.566	1.048	27.988	29.545	29.914	30.173	30.260

$SYNCH$ refers to the stock price synchronicity measure estimated using Equation (4) and transformed using Equation (4) for firms with A shares only. All other variables are defined as in Appendix A.2.

Table 5
Change in Stock Price Synchronicity of REE Companies after the WTO Event – Univariate Evidence

		Panel A: Equally-Weighted FOB			Panel B: Equally-Weighted China			Panel C: Usage-Weighted FOB			Panel D: Usage-Weighted China		
$[\tau_1; \tau_2]$	*# REECs*	[-2]	[-1]	[Pre]	[-2]	[-1]	[Pre]	[-2]	[-1]	[Pre]	[-2]	[-1]	[Pre]
[-1; +1]	10	-0.573***	0.062	-0.54***	0.184	0.053	0.314*	-0.784***	-0.201	-0.341	0.172	-0.214*	0.172
		(-2.509)	(0.405)	(-2.355)	(1.119)	(0.2)	(1.426)	(-4.072)	(-1.191)	(-1.312)	(1.032)	(-1.63)	(0.844)
	25	-0.444***	0.048	-0.362***	0.003	0.041	0.147	-0.311***	-0.096	-0.339***	0.002	-0.122	0.228***
		(-3.523)	(0.42)	(-2.788)	(0.037)	(0.342)	(1.203)	(-2.206)	(-0.978)	(-2.674)	(0.015)	(-1.242)	(1.982)
	50	-0.303***	-0.057	-0.295***	0.015	-0.081	0.007	-0.306***	-0.114*	-0.252***	-0.029	-0.067	-0.018
		(-3.204)	(-0.75)	(-3.48)	(0.208)	(-0.928)	(0.079)	(-3.446)	(-1.457)	(-3.175)	(-0.363)	(-0.888)	(-0.201)
[-1; +2]	10	-0.676***	-0.249*	-0.517***	-0.387***	0.087	0.367*	-0.771***	-0.466***	-0.344***	-0.38***	-0.247	0.059
		(-4.617)	(-1.432)	(-3.365)	(-2.3)	(0.291)	(1.646)	(-6.362)	(-3.226)	(-1.98)	(-2.255)	(-1.057)	(0.398)
	25	-0.602***	-0.332***	-0.452***	-0.444***	-0.151	-0.016	-0.572***	-0.543***	-0.479***	-0.504***	-0.393***	0
		(-5.775)	(-2.259)	(-4.825)	(-3.49)	(-0.975)	(-0.12)	(-4.606)	(-4.927)	(-3.97)	(-4.14)	(-3.364)	(0.001)
	50	-0.502***	-0.358***	-0.457***	-0.441***	-0.397***	-0.294***	-0.497***	-0.457***	-0.396***	-0.455***	-0.491***	-0.209***
		(-5.456)	(-3.97)	(-5.642)	(-4.969)	(-3.927)	(-3.007)	(-5.278)	(-5.424)	(-4.767)	(-5.678)	(-4.953)	(-2.262)
[Pre; Post]	10	-0.676***	-0.309**	-0.574***	-0.324***	-0.335***	-0.152*	-0.779***	-0.385***	-0.393***	-0.362***	-0.362***	-0.167*
		(-6.018)	(-1.815)	(-5.816)	(-3.026)	(-2.926)	(-1.402)	(-9.743)	(-2.575)	(-2.047)	(-3.214)	(-3.714)	(-1.443)
	25	-0.618***	-0.319***	-0.491***	-0.388***	-0.268***	-0.268***	-0.573***	-0.437***	-0.485***	-0.435***	-0.354***	-0.18***
		(-8.1)	(-2.757)	(-5.847)	(-4.957)	(-3.906)	(-3.351)	(-7.159)	(-4.344)	(-4.792)	(-5.787)	(-5.412)	(-2.261)
	50	-0.558***	-0.23***	-0.445***	-0.385***	-0.388***	-0.365***	-0.498***	-0.377***	-0.415***	-0.399***	-0.376***	-0.29***
		(-8.962)	(-2.091)	(-6.548)	(-7.012)	(-5.384)	(-6.396)	(-6.834)	(-4.952)	(-6.24)	(-7.528)	(-6.862)	(-4.746)

This table reports the changes in stock price synchronicity of REE companies (*REECs*) after the WTO event. *REECs* are identified by applying Equation (6) using four different REE indices for the four most important elements (Ce, La, Nd, and Y) to proxy for price developments in the REE market. The equally-weighted REE index based on FOB (foreign) prices is displayed in panel A, and China (domestic) prices are in panel B; the usage-weighted REE index based on FOB prices is displayed in panel C, and China prices are in panel D. The selection periods for the *REECs* are one year, two years, or the entire *preWTO* period. The periods $[\tau_1; \tau_2]$ for which we compare stock price synchronicity are $[\tau_{-1}; \tau_{+1}]$, $[\tau_{-1}; \tau_{+2}]$, and $[\tau_{Pre}; \tau_{Post}]$. The changes in stock price synchronicity are calculated for portfolios of 10, 25, and 50 companies, with the highest R^2 in Equation (6). *t*-statistics are in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 6
Correlation Matrix for Stock Price Synchronicity and Firm Characteristics

		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
(1)	$SYNCH$	1								
(2)	$Top\ Gov$	0.1446*	1							
(3)	$Volume$	0.1156*	-0.1032*	1						
(4)	$Size$	0.1377*	0.1130*	-0.4211*	1					
(5)	$Leverage$	-0.0397*	0.0481*	-0.0037	0.3183*	1				
(6)	$Std(RoA)$)	-0.0308*	0.0571*	0.0047	-0.1548*	0.0239	1			
(7)	M/B	-0.0271	-0.0071	0.0365*	-0.0247	-0.0289	0.0062	1		
(8)	Ind_Num	-0.0087	-0.0664*	-0.0259	0.0079	0.0433*	0.0159	0.0262	1	
(10)	Ind_Size	-0.0907*	-0.1096*	-0.1259*	0.0628*	0.0147	-0.0099	0.0168	0.9438*	1

This table gives Pearson correlation coefficients for all variables presented in Table 4. All variables are considered in subsequent analyses (see appendix A.2 for variable descriptions and calculation methods). * indicates correlations are statistically significant at least at the 5% level.

Table 7
Change in Stock Price Synchronicity of REE Companies after the WTO Event – Multivariate Evidence

# *REECs*	Weighting		(1)	(2)	(3)	(4)	(5)	(6)
50	Equally	$PostWTO(\tau_2) \times REEC$	-0.170*** (-3.33)					
50		$PreWTO(\tau_1) \times REEC$	0.202*** (3.86)					
25		$PostWTO(\tau_2) \times REEC$		-0.146** (-2.11)				
25		$PreWTO(\tau_1) \times REEC$		0.259*** (4.51)				
10		$PostWTO(\tau_2) \times REEC$			-0.213* (-1.88)			
10		$PreWTO(\tau_1) \times REEC$			0.239*** (3.05)			
50	Usage	$PostWTO(\tau_2) \times REEC$				-0.155*** (-3.06)		
50		$PreWTO(\tau_1) \times REEC$				0.192*** (3.85)		
25		$PostWTO(\tau_2) \times REEC$					-0.251*** (-3.39)	
25		$PreWTO(\tau_1) \times REEC$					0.140** (2.17)	
10		$PostWTO(\tau_2) \times REEC$						-0.258** (-2.36)
10		$PreWTO(\tau_1) \times REEC$						0.246*** (3.88)

(*continued*)

Table 7 – *continued*

			Control Variables			
$Top\ Gov$	0.413***	0.418***	0.423***	0.415***	0.413***	0.422***
	(7.97)	(8.03)	(8.14)	(7.98)	(7.94)	(8.12)
$Volume$	0.038***	0.038***	0.038***	0.038***	0.038***	0.038***
	(10.39)	(10.41)	(10.48)	(10.41)	(10.43)	(10.42)
$Size$	0.135***	0.134***	0.135***	0.134***	0.135***	0.134***
	(16.28)	(16.25)	(16.30)	(16.26)	(16.27)	(16.26)
$Leverage$	-0.507***	-0.507***	-0.508***	-0.507***	-0.507***	-0.507***
	(-10.25)	(-10.25)	(-10.26)	(-10.26)	(-10.25)	(-10.23)
$Std(RoA))$	-0.183*	-0.186*	-0.187*	-0.189*	-0.178	-0.182*
	(-1.66)	(-1.77)	(-1.77)	(-1.72)	(-1.64)	(-1.73)
M/B	-0.001*	-0.001*	-0.001*	-0.001*	-0.001*	-0.001*
	(-1.77)	(-1.76)	(-1.77)	(-1.77)	(-1.76)	(-1.76)
Ind_Num	0.533***	0.541***	0.548***	0.534***	0.545***	0.546***
	(15.17)	(15.53)	(15.73)	(15.23)	(15.59)	(15.67)
Ind_Size	-0.444***	-0.450***	-0.456***	-0.445***	-0.455***	-0.455***
	(-15.28)	(-15.63)	(-15.84)	(-15.34)	(-15.72)	(-15.76)
Constant	6.457***	6.601***	6.714***	6.472***	6.709***	6.688***
	(9.54)	(9.82)	(10.01)	(9.60)	(9.95)	(9.95)
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	4,307	4,307	4,307	4,307	4,307	4,307
Adjusted R^2	0.187	0.185	0.183	0.186	0.185	0.184
F-statistic	78.512	78.869	78.062	78.503	78.212	79.579

We run standard OLS regressions (using robust standard errors) to identify the factors that determine stock price synchronicity as calculated in Equation (4). The coefficients and respective *t*-statistics are in parentheses below. The independent variables are $PostWTO(\tau_2) \times REEC$, $PreWTO(\tau_1) \times REEC$, $Top\ Gov$, $Volume$, $Size$, $Leverage$, $Std(RoA))$, M/B, Ind_Num, and Ind_Size (see appendix A.2 for variable descriptions and calculation methods). $PostWTO(\tau_2)$ is equal to 1 for the 2.5-year period after the WTO event, and 0 otherwise; $PreWTO(\tau_1)$ is equal to 1 for the 2.5-year period before the WTO event, and 0 otherwise. The selection period for the *REECs* is two years prior to the WTO event. The regressions are applied for the equally-weighted and the usage-weighted FOB REE Indices to form portfolios of 10, 25, and 50 companies with the highest R^2 in Equation (6). Specifications (1)-(3) are the multivariate pendants of the univariate results from Table 5 in panel (A), Equally-Weighted FOB REEs Index for a selection period two years before the WTO-event and the comparison [Pre; Post]. Similarly, specifications (4)-(6) are the multivariate pendants of the univariate results from Table 5 in panel (C), Equally-Weighted FOB REE Index for a selection period of two years before the WTO event and the comparison [Pre; Post]. For better orientation, the corresponding univariate results are framed with dotted lines in Table 5. Investigating the variance inflation factors (VIFs) reveals no evidence of multicollinearity, because all VIFs are well below the critical value of 5 (see Kutner et al., 2005). ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 8

Change in Stock Price Synchronicity of REE Industry Companies after the WTO Event – Multivariate Evidence (Alternative Estimation Strategies)

Panel A: Two-Stage Least Squares with Newey-West Standard Errors								
# *REECs*	Weighting		(1)	(2)	(3)	(4)	(5)	(6)
50	Equally	$PostWTO(\tau_2) \times REEC$	-0.183*** (-3.25)					
50		$PreWTO(\tau_1) \times REEC$	0.222*** (3.79)					
25		$PostWTO(\tau_2) \times REEC$		-0.157** (-2.07)				
25		$PreWTO(\tau_1) \times REEC$		0.284*** (4.42)				
10		$PostWTO(\tau_2) \times REEC$			-0.222** (-1.99)			
10		$PreWTO(\tau_1) \times REEC$			0.275*** (3.32)			
50	Usage	$PostWTO(\tau_2) \times REEC$				-0.166*** (-2.98)		
50		$PreWTO(\tau_1) \times REEC$				0.210*** (3.88)		
25		$PostWTO(\tau_2) \times REEC$					-0.264*** (-3.26)	
25		$PreWTO(\tau_1) \times REEC$					0.165** (2.29)	
10		$PostWTO(\tau_2) \times REEC$						-0.268*** (-2.59)
10		$PreWTO(\tau_1) \times REEC$						0.265*** (3.38)
		Control Variables	Yes	Yes	Yes	Yes	Yes	Yes
		Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
		Observations	4,307	4,307	4,307	4,307	4,307	4,307

(*continued*)

Table 8– ***continued***

# *REECs*	Weighting	Panel B: Random Effects Estimation	(1)	(2)	(3)	(4)	(5)	(6)
50		$PostWTO(\tau_2) \times REEC$	-0.166***					
			(-2.65)					
50		$PreWTO(\tau_1) \times REEC$	0.131*					
			(1.90)					
25	Equally	$PostWTO(\tau_2) \times REEC$		-0.146*				
				(-1.69)				
25		$PreWTO(\tau_1) \times REEC$		0.191**				
				(2.02)				
10		$PostWTO(\tau_2) \times REEC$			-0.209			
					(-1.56)			
10		$PreWTO(\tau_1) \times REEC$			0.203			
					(1.41)			
50		$PostWTO(\tau_2) \times REEC$				-0.148**		
						(-2.37)		
50		$PreWTO(\tau_1) \times REEC$				0.118*		
						(1.73)		
25	Usage	$PostWTO(\tau_2) \times REEC$					-0.267***	
							(-3.06)	
25		$PreWTO(\tau_1) \times REEC$					0.067	
							(0.72)	
10		$PostWTO(\tau_2) \times REEC$						-0.258*
								(-1.92)
10		$PreWTO(\tau_1) \times REEC$						0.186
								(1.34)
		Control Variables	Yes	Yes	Yes	Yes	Yes	Yes
		Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
		Observations	4,307	4,307	4,307	4,307	4,307	4,307

This table replicates Table 7, but shows results for alternative estimation strategies to the OLS regressions used in Table 7 in order to identify factors determining stock price synchronicity as calculated in Equation (4). Independent variables are equal to those in Table 7, but for clarity we do not report the coefficients and *t*-statistics for the control variables. Panel A shows the results for two-stage least squares regressions using Newey-West standard errors with lag(1); in panel B, random effects models are used. Investigating the variance inflation factors (VIFs) reveals no evidence of multicollinearity, because all VIFs are well below the critical value of 5 (see Kutner et al., 2005). ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 9
Static Difference-in-Differences (DiD) Between Option Pricing Models Before and After the WTO Event

		Call Options				Put Options			
		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
		PD	$\|BS-NGARCH\|$	$\|BS-HN\|$	$\|NGARCH-HN\|$	PD	$\|BS-NGARCH\|$	$\|BS-HN\|$	$\|NGARCH-HN\|$
Cerium	$PreWTO$	188.32	26.91	92.53	68.88	150.67	14.73	68.36	67.58
	$PostWTO$	57.88	27.54	27.83	2.51	16.83	7.62	7.49	1.72
	DiD	-130.44	0.63	-64.70	-66.36	-133.84	-7.11	-60.87	-65.86
	DiD in%	***-69.26%***	***2.33%***	***-69.92%***	***-96.35%***	***-88.83%***	***-48.29%***	***-89.04%***	***-97.46%***
Lanthanum	$PreWTO$	230.76	106.24	114.86	9.67	183.98	83.11	91.92	8.96
	$PostWTO$	55.11	26.39	26.17	2.54	16.95	7.68	7.69	1.58
	DiD	-175.65	-79.84	-88.68	-7.13	-167.03	-75.43	-84.23	-7.38
	DiD in%	***-76.12%***	***-75.16%***	***-77.21%***	***-73.69%***	***-90.79%***	***-90.76%***	***-91.63%***	***-82.39%***
Neodymium	$PreWTO$	138.68	69.33	34.09	35.26	93.96	46.29	10.91	36.76
	$PostWTO$	51.12	25.10	22.98	3.04	15.42	6.43	6.41	2.57
	DiD	-87.56	-44.23	-11.11	-32.22	-78.55	-39.86	-4.49	-34.19
	DiD in%	***-63.14%***	***-63.79%***	***-32.59%***	***-91.38%***	***-83.59%***	***-86.11%***	***-41.19%***	***-93.01%***
Yttrium	$PreWTO$	377.26	188.32	137.99	50.96	382.06	190.91	117.97	73.18
	$PostWTO$	77.10	36.52	36.95	3.62	30.75	13.75	13.97	3.03
	DiD	-300.16	-151.79	-101.03	-47.34	-351.31	-177.16	-104.00	-70.15
	DiD in%	***-79.56%***	***-80.61%***	***-73.22%***	***-92.90%***	***-91.95%***	***-92.80%***	***-88.16%***	***-95.87%***

This table shows the differences among three option pricing models (Black-Scholes, 1973; Duan, 1995; and Heston and Nandi, 2000) for put and call options on REEs (cerium, lanthanum, neodymium, and yttrium) (see Fig. 2 for an illustration). Column (2) shows the sum of thirty-five absolute differences between the Black-Scholes (1973) and Duan (1995) call option prices. We calculate the prices for 35 [= 7 × 5] distinct combinations of *time to maturity* and *strike price* (seven *strike prices* [85%, 90%, …, 115%] and five *times to maturity* [one, three, six, nine, and twelve months]). Columns (3) and (4) are similar, but show instead the differences in call option prices between Black-Scholes (1973) and Heston and Nandi (2000) (column (3)), and Duan (1995) and Heston and Nandi (2000) (column (4)). Column (1) is equal to the sum of columns (2), (3), and (4) (see Table 10 for a sample calculation for lanthanum, which is framed with a dotted line and shaded in grey). Columns (5), (6), (7), and (8) follow the same system for put options. The differences between the option pricing models are calculated at the WTO event based on the REE prices thirty months before ($PreWTO$) and are compared to thirty months after ($PostWTO$) the WTO event, but considering REE prices beginning at the WTO event and for the subsequent thirty months. To show the difference-in-difference ("DiD") between the option model pricing differences in response to the WTO event, we subtract from the respective squared $PreWTO$ option price differences (PD_{PreWTO}) the absolute $PostWTO$ option price differences ($PD_{PostWTO}$). The line "*DiD in%*" shows the relative changes in percentage terms.

Table 10
Sample Calculation for the Option Price Differences between Option Pricing Models (Lanthanum; $PreWTO$; Call Options)

$PreWTO$ Lanthanum				
(1) y	(2) S/K	(3) $\lvert BS - NGARCH \rvert$	(4) $\lvert BS - HN \rvert$	(5) $\lvert NGARCH - HN \rvert$
1m	0.85	0.22	0.28	0.06
	0.9	0.62	0.68	0.06
	0.95	1.15	1.19	0.04
	1	1.23	1.52	0.29
	1.05	1.07	1.23	0.15
	1.1	0.56	0.74	0.17
	1.15	0.38	0.44	0.06
	SUM	5.23	6.08	0.85
3m	0.85	1.36	1.57	0.21
	0.9	1.94	2.17	0.23
	0.95	2.44	2.49	0.04
	1	2.37	2.43	0.06
	1.05	2.13	2.32	0.19
	1.1	2.04	2.18	0.14
	1.15	1.53	1.46	0.07
	SUM	13.82	14.63	0.94
6m	0.85	2.83	2.81	0.01
	0.9	3.60	3.66	0.06
	0.95	3.52	3.37	0.16
	1	3.28	3.89	0.61
	1.05	3.41	3.46	0.06
	1.1	3.17	3.31	0.15
	1.15	2.08	2.90	0.82
	SUM	21.88	23.40	1.87
9m	0.85	3.92	4.41	0.49
	0.9	4.43	4.93	0.50
	0.95	4.58	5.36	0.78
	1	5.02	4.73	0.29
	1.05	4.50	4.63	0.12
	1.1	4.00	4.08	0.08
	1.15	4.03	4.78	0.75
	SUM	30.48	32.92	3.01
12m	0.85	4.70	5.49	0.79
	0.9	4.99	5.59	0.60
	0.95	5.13	5.40	0.27
	1	5.80	5.88	0.08
	1.05	5.05	5.42	0.37
	1.1	5.15	5.41	0.26
	1.15	4.01	4.63	0.62
	SUM	34.83	37.82	3.00
PD	**230.76**	106.24	114.86	9.67

This table shows detailed calculations for each absolute pricing difference between all three option pricing models for all distinct lanthanum call option combinations for *times to maturity* and *strike prices*. The calculations correspond to the framed dotted cells, which are shaded in grey in Table 9 and for the last line PD shows the aggregated values also shown in Table 9 in column (1). y is equal to *time to maturity* and S/K is the ratio of the spot price to the *strike price*.

Table 11
Consecutive Triple-Difference (DDD) Approach Between Option Pricing Models Before and After the WTO Event

Cerium	$\lvert BS-NGARCH \rvert$			$\lvert BS-HN \rvert$			$\lvert NGARCH-HN \rvert$			No obs
S/K	$\overline{DiD}\binom{PreWTO\ -}{PrePreWT}$	$\overline{DiD}\binom{PreWTO\ -}{PostWTO}$	t-stat (DDD)	$\overline{DiD}\binom{PreWTO\ -}{PrePreWT}$	$\overline{DiD}\binom{PreWTO\ -}{PostWTO}$	t-stat (DDD)	$\overline{DiD}\binom{PreWTO\ -}{PrePreWT}$	$\overline{DiD}\binom{PreWTO\ -}{PostWTO}$	t-stat (DDD)	
0.85	27.255	26.924	-1.80	3.125	1.862	-9.65	28.745	26.901	-14.56	150
0.9	9.446	9.305	-0.69	3.417	2.012	-9.55	11.734	9.453	-16.06	150
0.95	9.116	9.283	0.73	3.482	2.164	-7.49	12.460	10.045	-17.55	150
1	12.115	12.369	1.11	3.420	2.197	-6.61	15.540	13.344	-17.89	150
1.05	11.836	11.754	-0.47	3.300	1.943	-10.31	14.723	12.969	-16.33	150
1.1	12.731	12.499	-1.71	2.940	1.742	-11.13	15.095	13.852	-14.66	150
1.15	23.521	23.216	-2.81	2.554	1.526	-11.41	25.372	24.536	-13.48	150
y										
1m	6.007	5.798	-5.11	0.689	0.361	-8.39	5.917	5.700	-11.98	210
3m	28.969	28.664	-3.42	2.082	1.215	-11.74	27.585	26.847	-18.55	210
6m	26.744	26.598	-0.99	3.427	2.065	-12.39	27.508	25.784	-28.17	210
9m	6.722	6.749	0.14	4.456	2.736	-12.63	12.584	9.877	-37.55	210
12m	7.287	7.439	0.65	5.231	3.227	-12.45	14.741	11.148	-44.48	210

Lanthanum	$\lvert BS-NGARCH \rvert$			$\lvert BS-HN \rvert$			$\lvert NGARCH-HN \rvert$			No obs
S/K	$\overline{DiD}\binom{PreWTO\ -}{PrePreWT}$	$\overline{DiD}\binom{PreWTO\ -}{PostWTO}$	t-stat (DDD)	$\overline{DiD}\binom{PreWTO\ -}{PrePreWT}$	$\overline{DiD}\binom{PreWTO\ -}{PostWTO}$	t-stat (DDD)	$\overline{DiD}\binom{PreWTO\ -}{PrePreWT}$	$\overline{DiD}\binom{PreWTO\ -}{PostWTO}$	t-stat (DDD)	
0.85	1.694	2.326	7.57	1.049	1.450	5.60	0.963	1.327	17.30	150
0.9	1.897	2.600	6.37	1.261	1.693	4.52	1.001	1.383	19.54	150
0.95	1.996	2.694	5.07	1.351	1.783	3.45	1.028	1.403	20.62	150
1	1.986	2.656	4.59	1.316	1.774	3.44	1.068	1.354	15.61	150
1.05	1.887	2.551	5.89	1.112	1.648	5.11	1.134	1.309	10.80	150
1.1	1.749	2.303	6.44	0.883	1.385	6.23	1.187	1.263	5.46	150
1.15	1.513	1.955	6.69	0.706	1.091	6.35	1.044	1.109	4.54	150
y										
1m	0.312	0.471	6.77	0.268	0.376	5.96	0.128	0.252	11.47	210
3m	0.951	1.382	8.52	0.720	1.009	6.55	0.523	0.717	15.12	210
6m	1.918	2.599	8.29	1.188	1.675	6.52	1.119	1.373	14.83	210
9m	2.657	3.521	7.82	1.526	2.160	6.36	1.581	1.887	16.80	210
12m	3.248	4.231	7.17	1.782	2.510	5.74	1.952	2.305	16.91	210

(continued)

Table 11 – *continued*

Neodymium	$\lvert BS-NGARCH\rvert$			$\lvert BS-HN\rvert$			$\lvert NGARCH-HN\rvert$			No bs
S/K	$\overline{D\iota D}(PreWTO-PrePreWTO)$	$\overline{D\iota D}(PreWTO-PostWTO)$	t-stat (DDD)	$\overline{D\iota D}(PreWTO-PrePreWTO)$	$\overline{D\iota D}(PreWTO-PostWTO)$	t-stat (DDD)	$\overline{D\iota D}(PreWTO-PrePreWTO)$	$\overline{D\iota D}(PreWTO-PostWTO)$	t-stat (DDD)	
0.85	4.550	6.666	10.77	-0.034	2.049	10.63	8.065	8.127	8.10	150
0.9	4.444	6.899	11.08	0.028	2.448	10.96	9.017	9.062	5.81	150
0.95	5.610	8.290	10.44	0.200	2.848	10.36	11.296	11.363	6.28	150
1	5.902	8.595	10.37	0.326	2.971	10.26	11.747	11.835	6.20	150
1.05	6.653	9.061	11.68	0.078	2.411	11.42	12.044	12.113	4.44	150
1.1	14.594	16.686	11.66	0.003	2.010	11.21	18.974	19.056	5.20	150
1.15	11.130	12.854	10.97	-0.088	1.648	11.03	14.661	14.750	6.11	150
y										
1m	3.532	4.156	10.88	-0.107	0.511	10.81	4.045	4.067	5.97	210
3m	22.544	24.112	14.63	-0.036	1.494	14.28	22.995	23.045	7.34	210
6m	6.436	8.923	15.56	0.074	2.500	15.21	10.269	10.337	7.37	210
9m	2.676	5.829	15.87	0.176	3.287	15.66	11.053	11.160	8.03	210
12m	2.586	6.304	15.61	0.258	3.912	15.44	12.926	13.039	7.36	210

Yttrium	$\lvert BS-NGARCH\rvert$			$\lvert BS-HN\rvert$			$\lvert NGARCH-HN\rvert$			No bs
S/K	$\overline{D\iota D}(PreWTO-PrePreWTO)$	$\overline{D\iota D}(PreWTO-PostWTO)$	t-stat (DDD)	$\overline{D\iota D}(PreWTO-PrePreWTO)$	$\overline{D\iota D}(PreWTO-PostWTO)$	t-stat (DDD)	$\overline{D\iota D}(PreWTO-PrePreWTO)$	$\overline{D\iota D}(PreWTO-PostWTO)$	t-stat (DDD)	
0.85	-0.746	0.998	8.97	2.942	4.686	8.98	3.769	3.825	7.57	150
0.9	-0.696	1.307	9.23	3.472	5.466	9.21	4.227	4.270	5.39	150
0.95	-0.576	1.632	8.58	3.810	6.026	8.64	4.420	4.476	5.00	150
1	-0.445	1.780	8.43	3.925	6.114	8.28	4.316	4.371	5.32	150
1.05	-0.340	1.634	9.40	3.520	5.514	9.67	3.807	3.869	4.91	150
1.1	-0.226	1.460	10.28	3.044	4.754	10.26	3.202	3.243	2.84	150
1.15	-0.195	1.238	9.88	2.537	3.939	9.70	2.621	2.669	3.26	150
y										
1m	-0.364	0.156	9.50	0.903	1.423	9.37	1.416	1.436	5.29	210
3m	-0.763	0.541	12.08	2.184	3.492	12.12	2.952	2.991	6.23	210
6m	-0.678	1.367	12.64	3.558	5.598	12.58	4.132	4.188	6.25	210
9m	-0.384	2.200	12.53	4.555	7.135	12.57	4.869	4.935	5.57	210
12m	-0.114	2.913	12.52	5.406	8.422	12.50	5.462	5.538	5.47	210

This table shows the mean values of the DiDs from the "control" ($\overline{D\iota D}_k(PreWTO - PrePreWTO)$) and "treatment" periods ($\overline{D\iota D}_k(PreWTO - PostWTO)$) for the REEs cerium, lanthanum, neodymium, and yttrium based on call options. "t-stat (DDD)" is the t −statistic testing DDD_k against zero. y is equal to *time to maturity*, and S/K is the ratio of the spot price to the *strike price*.

Do Announcements of WTO Dispute Resolution Cases Matter? Evidence from the Rare Earth Elements Market

Highlights

- Rare earth elements (REEs) are important for green- and high-technology products
- Announcements of WTO dispute resolution cases have the power to change market dynamics
- The price-generating process of REEs changes after the announcement
- Stock price informativeness of companies in the REEs industry increases after the announcement
- Model uncertainty for option pricing models decreases after the announcement