Center For International Economics

Working Paper Series

Working Paper No. 2017-05

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

Juliane Proelss, Denis Schweizer, and Volker Seiler

April 2017

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

Juliane Proelss*, Denis Schweizer†, and Volker Seiler‡

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 the 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 the model uncertainty for option pricing models decreases, which we measure by the lower pricing differences among them.

JEL Classification: C22, C58, F13, G14, G18, G28, Q02, Q38

Keywords: Market Efficiency, Rare Earth Elements, Stock Price Informativeness, Structural Break Tests, Variance Ratio Tests, World Trade Organization (WTO)

---

* Concordia University, Assistant Professor of Finance, John Molson School of Business Building, 1450 Guy, Montreal, Quebec, Canada H3H 0A1, Phone: +1 514-848-2424, ext. 2242, Fax: +1-514-848-4500, Email: Juliane.Proelss@concordia.ca.

† Corresponding Author: Concordia University, Associate Professor of Finance, John Molson School of Business Building, 1450 Guy, Montreal, Quebec, Canada H3H 0A1, Phone: +1 514-848-2424, ext. 2926, Fax: +1-514-848-4500, Email: Denis.Schweizer@concordia.ca.

‡ University of Paderborn, Akademischer Rat, International Economics and Macroeconomics, Warburger Straße 100, 33098 Paderborn, Germany and Visiting Research Professor, New York University, Leonard N. Stern School of Business, Phone: +49 5251 - 60 2970, Fax: +49 5251 - 60 4372, Email: Volker.Seiler@wiwi.uni-paderborn.de.

**Acknowledgments: We thank Christopher F. Baum, Jennifer Conrad, Douglas Cumming, Wenxuan Hou, Nengsheng Fang, Christian Koziol, Hong Li, Benjamin Liebman, Oliver Linton, Marcel Prokopczuk, Morten Ravn, Jay Ritter, Timo Teräsvirta, Craig Wilson, and Sjur Westgaard, as well as participants at the NSVCM Workshop on non- and semiparametric volatility and correlation models 2014 (Paderborn, Germany), the Commodity Markets Workshop 2015 (Oslo, Norway), the Forum for Economists International Conference 2015 (Amsterdam, Netherlands), and the 5th Conference on the Chinese Capital Markets 2015 (Winnipeg, Canada) for helpful comments and suggestions. We also thank Qi Sun for excellent research assistance. Denis Schweizer gratefully acknowledges the financial support provided through the Manulife Professorship. A previous version was named "Do WTO Rulings Really Matter? Evidence from Rare Earth Markets."

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

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 is overly restrictive, which 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 the 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 the model uncertainty for option pricing models decreases, which we measure by the lower pricing differences among them.

JEL Classification: C22, C58, F13, G14, G18, G28, Q02, Q38

Keywords: Market Efficiency, Rare Earth Elements, Stock Price Informativeness, Structural Break Tests, Variance Ratio Tests, World Trade Organization (WTO)

1. Introduction

The rare earth elements (REEs) market has gained increasing attention in recent years because these elements 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).[4] However, well above 80% of global REE mine production and more than 50% of worldwide REE reserves are located in China (see, for example, Müller, Schweizer, and Seiler, 2016). 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[5]) (see Figure 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 (see Figure 2).

– Please insert Figures 1 and 2 about here –

[4] 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).

[5] Müller, Schweizer, and Seiler (2015) provide an overview of the export quota announcements for the period July 7, 2008 to July 8, 2014.

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, WTO event). This has led to a series of reports concerning how to better secure access to REEs by the U.S. Congress (see Figure 1, step 2, as well as the European Commission, 2012; Humphries, 2012; Morrison and Tang, 2012; and Bailey Grasso, 2013).

The WTO generally focuses on negotiating new agreements for reducing trade burdens among their member states. But 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, Rehbein, and Starks, 1990; Liebman and Tomlin, 2007, 2008; and Lenway, Morck, and Yeung, 1996, for research on shareholder behavior).[6] 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) are the first to investigate the effects of the *announcement* of filing 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).

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

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 their REE policy to somewhat reduce their influence on REEs (see Figure 1, step 3). We observe direct evidence for this behavior in the Chinese government's statement at the end of 2014 that they will 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).[7] 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.[8]

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.

[7] See TMR (2014), WTO (2014a, 2014b, 2014c, 2014d, 2014e, 2014f), and the online appendix OA1 of Müller, Schweizer, and Seiler (2016) for more information on the dispute resolution case.

[8] For an overview of the main events surrounding the WTO dispute resolution case launched by the U.S., the EU, and Japan against China concerning China's export policies with regards to REE, see Figure 1 in the online appendix of Müller, Schweizer, and Seiler (2016).

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, Kim, and Qiu, 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 will 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 Figure 1, step 4).

Fourth, as a whole, our study contributes to the extant and growing literature on the effects of trade disputes and the WTO's rulings thereof in several ways. While the 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, Rehbein, and Starks, 1990; Lenway, Morck, and Yeung, 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 we also observe an increase in stock price informativeness for REE companies.

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.[9]

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 the random walk hypothesis, 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).

[9] 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).

Formally, we calculate the variance ratios as follows:

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

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.

2.2. Multiple Structural Change Tests

Besides using variance ratio tests to gauge whether the commencement of a WTO trial had 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 $\hat{T}_i$ $(i=1,\dots,\ell+1)$ where $\hat{T}_0=0$ and $\hat{T}_{\ell+1}=T$. An additional break is included as long as the overall minimal 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:

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

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 follow the methodology of Gul, Kim, and Qiu (2010) and Tan et al. (2015) closely. 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 (see Müller, Schweizer, and Seiler, 2016, for a detailed overview of the developments during that trial). The measure of choice is stock price synchronicity, as used by Gul, Kim, and Qiu (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.[10] $R_{i,\tau}^2$ is obtained from the regression in Equation (7):[11]

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

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

[10] 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.

[11] 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.

Shenzhen stock exchange) from CSMAR.[12] $IR_t^{(j)}$ represents the industry index return on day t for the respective industry j, based on the six industry classifications from *Industry Code A*.

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,[13] 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, Goldstein, and Jiang, 2007, for a more detailed discussion). In line with Gul, Kim, and Qiu (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

[12] 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.
[13] 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.

that developed markets such as the U.S. have lower R^2 and lower (usually negative) $SYNCH_{i,\tau}$ than emerging markets.

Finally, to identify implicitly with a statistical method, and analogously to our previous approach, we run the following regression for companies in the REE industry. 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. 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 Figure 2; Morrison and Tang, 2012; and Bailey Grasso, 2013, for more details about the different price schemes).

In the next step, we select the 10, 25, and 50 companies with the highest R^2 in Equation (8) 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 operated directly in the REE industry or held stakes in companies in the REE industry (see Table A1 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, Schweizer, and Seiler (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:

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

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 (6) (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 A2 in the appendix for variable descriptions and calculations). ξ_τ are year fixed effects to control for potential macroeconomic trends in China (see appendix A4 for detailed variable descriptions and calculation methods). 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.[14] 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

[14] Note that larger R^2s imply higher $SYNCH$, and thus higher market information and lower company-specific information than with lower R^2s and lower $SYNCH$.

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 option 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 Figure 3 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).[15]

We sum 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:

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

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

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

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)).

[15] We consider GARCH option pricing because we find that the REE returns exhibit GARCH effects (see Figure A1 in the appendix).

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 A5 for more 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,[16] 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 Figure 3, 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

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

($PreWTO$ and $PostWTO$ period), which would indicate a clear (causal) influence of the WTO event.

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:

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

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

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

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

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

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

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 ($\overline{DiD}_k(PreWTO - PrePreWTO)$) are lower than those from the "treatment" period ($\overline{DiD}_k(PreWTO - PostWTO)$), which would mean that DDD_k is statistically significantly greater than zero.

– Please insert Figure 3 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 a usage-weighted REE index (based on domestic prices—see appendix A3 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. Following Müller, Schweizer, and Seiler (2016), 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.

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 A3 in the appendix shows the descriptive statistics of the log returns.[17] 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.

[17] 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.

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 (see also Figure 2). 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 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). To do so, 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 1, Table 2).[18] 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.[19] We then compare the results from the pre-WTO period

[18] 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).

[19] 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

(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 $\bar{M}_r(q) + 1 = 1.401$ 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.

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 2 of Table 2, and underline our previous findings: After the announcement of the commencement of the WTO dispute resolution case by

upon request. The test statistic is robust to many forms of heteroscedasticity, which is critical given the apparent difference in variance before and after the peak for the REEs and related indices (see Figure 2).

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.[20] 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 –

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, Chopra, and Bessler (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).[21] More interestingly, we find breakpoints that are relatively close to the announcement of the WTO event (the dates for

[20] 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.

[21] Note that the break dates identified by Bai and Perron's (2003a, 2003b) $\sup F(\ell+1|\ell)$ test in Table 4 are not in chronological order, but are as identified by the testing procedure.

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.

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 strategies and for 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, Kim, and Qiu (2010), but much larger than the -1.742 reported by Piotroski and Roulstone (2004) for U.S. firms.[22] The rather small difference

[22] Comparing the means for the other firm-specific characteristics in our sample with the means reported in Gul, Kim, and Qiu (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, Kim, and Qiu (2010), which can be explained by the fact that trading volume in

from Gul, Kim, and Qiu (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 statistically supported by the evidence in Table 5, which shows that the stock price synchronicity of REE companies decreased after the WTO event.

However, 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 the 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

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, Kim, and Qiu (2010) divide by total net assets instead of book value of equity to calculate M/B.

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 shown in Figure 2, FOB prices are always higher than China prices. Therefore, 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).[23] We are primarily interested in the coefficients for the variables $PostWTO(\tau_2)\times REEC$ and $PreWTO(\tau_1)\times REEC$ (the DiD coefficients).[24] 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 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

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

[24] 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.

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, Kim, and Qiu, 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, Kim, and Qiu, 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, Kim, and Qiu, 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.[25]

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

[25] 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.

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.[26]

– 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 A4 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 (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.

[26] 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. (2015) for a similar argumentation in a related context.

First, 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. Second, 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, Schweizer, and Seiler, 2015, for example, already identified 14 REE mining and refining companies). This leads to the third reason, 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.

In summary, we present 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 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.[27]

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 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

[27] 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, Seppi, and Spatt, 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.

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 Figure 4, 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 Figure 4).

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 4 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). 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).[28]

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 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 statistically after the WTO event (see again Table 11).

[28] 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.

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 –

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; Desai and Hines Jr., 2008; Liebman and Tomlin, 2008; and Liebman and Tomlin, 2015). 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 and the subsequent price runups on the Shanghai and Shenzhen stock exchanges (see Hung,

2009), as well as the opening of Asian stock markets and the unprecedented inflow of funds into these markets after the Asian financial crisis in 1997 (see Hoque, Kim, and Pyun, 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 Figure A2 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 ceased intervening in the REE market 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), we show that even the announcement of a trial triggers economically significant changes in the REE market. We interpret our results as strong support for the notion that governments accused of violating GATT react to the announcement with "measurable" policy changes even before an official WTO ruling.

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 Figure 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, Sarin, and Shastri, 1998; and Danielsen, Van Ness, and Warr, 2007).

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; and Baur, 2014, for related research on the gold and oil market). Finally, due to the often extreme price movements of REEs, future research may also want to concentrate on volatility modeling and predictions (see Batten, Ciner, and Lucey, 2010; and Haugom et al., 2014, for the oil and metal markets).

References

Alexander, S.S. (1961): Price movements in speculative markets: Trends or random walks, in: Industrial Management Review, Vol. 2 (2), 7-26

Andrews, D.W.K. (1991): Heteroskedasticity and autocorrelation consistent covariance matrix estimation, in: Econometrica, Vol. 59 (3), 817-858

Andrews, D.W.K., and Monahan, J.C. (1992): An improved heteroscedasticity and autocorrelation consistent covariance matrix estimator, in: Econometrica, Vol. 60 (4), 953-966

Anstett, T.F. (1986): Availability of rare-earth, yttrium, and related thorium oxides – market economy countries. A minerals availability appraisal, in: Bureau of Mines Information Circular No. 9111

APS Panel on Public Affairs/The Materials Research Society (2011): Energy critical elements: Securing materials for emerging technologies, American Physical Society, Washington

Awokuse, T.O., Chopra, A., and Bessler, D.A. (2009): Structural change and international stock market interdependence: Evidence from Asian emerging markets, in: Economic Modelling, Vol. 26 (3), 549-559

Ayadi, O.F., and Pyun, C.S. (1994): An application of variance ratio test to the Korean securities market, in: Journal of Banking and Finance, Vol. 18 (4), 643-658

Bai, J., and Perron, P. (1998): Estimating and testing linear models with multiple structural changes, in: Econometrica, Vol. 66 (1), 47-78

Bai, J., and Perron, P. (2003a): Computation and analysis of multiple structural change models, in: Journal of Applied Econometrics, Vol. 18 (1), 1-22

Bai, J., and Perron, P. (2003b): Critical values for multiple structural change tests, in: Econometrics Journal, Vol. 6 (1), 72-78

Bailey Grasso, V. (2013): Rare earth elements in national defense: Background, oversight issues, and options for congress, Congressional Research Service Report R41744, http://www.fas.org/sgp/crs/natsec/R41744.pdf

Batten, J.A., Ciner, C., and Lucey, B.M. (2010): The macroeconomic determinants of volatility in precious metals markets, in: Resources Policy, Vol. 35 (2), 65-71

Baur, D.G. (2014): Gold mining companies and the price of gold, in: Review of Financial Economics, Vol. 23 (4), 174-181

Bhattacharya, S., and Ritter, J.R. (1983): Innovation and communication: Signalling with partial disclosure, in: Review of Economic Studies, Vol. 50 (2), 331-346

Binnemans, K., Jones, P.T., Blanpain, B., Van Greven, T., Yang, Y., Walton, A., and Buchert, M. (2013): Recycling of rare earths: A critical review, in: Journal of Cleaner Production, Vol. 51, 1-22

Black, F., and Scholes, M. (1973): The pricing of options and corporate liabilities, in: Journal of Political Economy, Vol. 81 (3), 637-645

Bloomberg News (2014): Chinese exchange starts rare earths trading trial today, http://www.bloomberg.com/news/2014-03-27/chinese-exchange-starts-rare-earths-trading-trial-today.html

Bond, E.W., and Trachtman, J. (2016): China-rare earths: Export restrictions and the limits of textual interpretation, in: World Trade Review, Vol. 15 (2), 189-209

Boyer, M.M., and Filion, D. (2007): Common and fundamental factors in stock returns of Canadian oil and gas companies, in: Energy Economics, Vol. 29 (3), 428-453

Casassus, J., and Collin-Dufresne, P. (2005): Stochastic convenience yield implied from commodity futures and interest rates, in: Journal of Finance, Vol. 60 (5), 2283-2331

Chan, K., and Hameed, A. (2006): Stock price synchronicity and analyst coverage in emerging markets, in: Journal of Financial Economics, Vol. 80 (1), 115-147

Charles, A., and Darné, O. (2009): Variance-ratio tests of random walk: An overview, in: Journal of Economic Surveys, Vol. 23 (3), 503-527

Chen, Q., Goldstein, I., and Jiang, W. (2007): Price informativeness and investment sensitivity to stock price, in: Review of Financial Studies, Vol. 20 (3), 619-650

Choi, H., and Varian, H. (2009): Predicting the present with Google Trends, Google Inc. Working Paper

Choi, H., and Varian, H. (2012): Predicting the present with Google Trends, in: Economic Record, Vol. 88 (s1), 2-9

Chow, K.V. and Denning, K.C. (1993): A simple multiple variance ratio test, in: Journal of Econometrics, Vol. 58 (3), 385-401

Cochrane, J. H. (1988): How big is the random walk in GNP? in: Journal of Political Economy, Vol. 96 (5), 893-920

Conrad, J. (1989): The price effect of option introduction, in: Journal of Finance, Vol. 44 (2), 487-498

Danielsen, B.R., Van Ness, B.F., and Warr, R.S. (2007): Reassessing the impact of option introductions on market quality: A less restrictive test for event-date effects, in: Journal of Financial and Quantitative Analysis, Vol. 42 (4), 1041-1062

Davey, W.J. (2003): The WTO dispute settlement mechanism, Available at SSRN: http://ssrn.com/abstract=419943

Desai, M.A., and Hines Jr., J.R. (2008): Market reactions to export subsidies, in: Journal of International Economics, Vol. 74 (2), 459-474

Detemple, J., and Jorion, P. (1990): Option listing and stock returns: An empirical analysis, in: Journal of Banking & Finance, Vol. 14 (4), 781-801

Duan, J.-C. (1995): The GARCH option pricing model, in: Mathematical Finance, Vol. 5 (1), 13-32

Epstein, L.G., and Schneider, M. (2008): Ambiguity, information quality, and asset pricing, in: Journal of Finance, Vol. 63 (1), 197-228

European Commission (2012): EU challenges China's rare earth export restrictions, European Commission – Press Release UP/12/239, http://europa.eu/rapid/press-release_IP-12-239_en.htm

Faff, R., and Chan, H. (1998): A multifactor model of gold industry stock returns: evidence from the Australian equity market, in: Applied Financial Economics, Vol. 8 (1), 21-28

Fan, J., and Wong, T.J. (2005): Do external auditors perform a corporate governance role in emerging markets? Evidence from East Asia, in: Journal of Accounting Research, Vol. 43 (1), 35-72

GAO (2010): Rare earth materials in the defense supply chain, United States Government Accountability Office Report, GAO-10-617R, http://www.gao.gov/new.items/d10617r.pdf

Gibson, R., and Schwartz, E. S. (1990): Stochastic convenience yield and the pricing of oil contingent claims, in: Journal of Finance, Vol. 45 (3), 959-976

Goonan, T.G. (2011): Rare earth elements – end use and recyclability, U.S. Geological Survey Scientific Investigations Report 2011 – 5094, http://pubs.usgs.gov/sir/2011/5094/

Gul, F.A., Kim, J.B., and Qiu, A.A. (2010): Ownership concentration, foreign shareholding, audit quality, and stock price synchronicity: Evidence from China, in: Journal of Financial Economics, Vol. 95 (3), 425-442

Haugom, E., Langeland, H., Molnár, P., and Westgaard, S. (2014): Forecasting volatility of the U.S. oil market, in: Journal of Banking and Finance, Vol. 47, 1-14

Heston, S.L., and Nandi, S. (2000): A closed-form GARCH option valuation model, in: Review of Financial Studies, Vol. 13 (3), 585-625

Hoque, H.A.A.B, Kim, J.H., and Pyun, C.S. (2007): A comparison of variance ratio tests of random walk: A case of Asian emerging stock markets, in: International Review of Economics & Finance, Vol. 16 (4), 488-502

Humphries, M. (2012): Rare earth elements: The global supply chain, Congressional Research Service Report R41347, http://www.fas.org/sgp/crs/natsec/R41347.pdf

Hung, J.-C. (2009): Deregulation and liberalization of the Chinese stock market and the improvement of market efficiency, in: Quarterly Review of Economics and Finance, Vol. 49 (3), 843-857

Ibáñez, A., and Zapatero, F. (2004): Monte Carlo valuation of American options through computation of the optimal exercise frontier, in: Journal of Financial and Quantitative Analysis, Vol. 39 (2), 253-275

Jackson, W.D., and Christiansen, G. (1993): International strategic minerals inventory summary report – rare earth oxides. U.S. Geological Survey Circular 930-N, http://pubs.usgs.gov/circ/1993/0930n/report.pdf

Jacob, M., Johan, S., Schweizer, D., and Zhan, F. (2015): Corporate finance and governance implications from the removal of government support programs, in: Journal of Banking and Finance, Forthcoming.

Johnson, S., La Porta, R., Lopez-De-Silanes, F., and Shleifer, A. (2000): Tunneling, in: American Economic Review, Vol. 90 (2), 22-27

Kendall, M.G. (1953): The analysis of economic time-series – Part I: Prices, in: Journal of the Royal Statistical Society. Series A (General), Vol. 116 (1), 11-25

Kim, J.H. (2006): Wild bootstrapping variance ratio tests, in: Economics Letters, Vol. 92 (1), 38-43

Kumar, R., Sarin, A., and Shastri, K. (1998): The impact of options trading on the market quality of the underlying security: An empirical analysis, in: Journal of Finance, Vol. 53 (2), 717-732

Kutner, M.H., Nachtsheim, C.J., Neter, J., and Li, W. (2005): Applied Linear Statistical Models, 5th edition. New York: McGraw-Hill

Lenway, S., Morck, R., and Yeung, B. (1996): Rent seeking, protectionism and innovation in the American steel industry, in: The Economic Journal, Vol. 106 (435), 410-421

Lenway, S., Rehbein, K., and Starks, L. (1990): The impact of protectionism on firm wealth: The experience of the steel industry, in: Southern Economic Journal, Vol. 56 (4), 1079-1093

Liebman, B.H. (2006): Safeguards, China, and the price of steel, in: Review of World Economics, Vol. 142 (2), 354-373

Liebman, B.H., and Tomlin, K.M. (2007): Steel safeguards and the welfare of U.S. steel firms and downstream consumers of steel: a shareholder wealth perspective, in: Canadian Journal of Economics, Vol. 40 (3), 812-842

Liebman, B.H., and Tomlin, K.M. (2008): Safeguards and retaliatory threats, in: Journal of Law & Economics, Vol. 51 (2), 351-376

Liebman, B.H., and Tomlin, K.M. (2015): World Trade Organization sanctions, implementation, and retaliation, in: Empirical Economics, Vol. 48 (2), 715-745

Liu, C.Y., and He, J. (1991): A variance-ratio test of random walks in foreign exchange rates, in: Journal of Finance, Vol. 46 (2), 773-785

Lo, A.W., and MacKinlay, A.C. (1988): Stock market prices do not follow random walks: Evidence from a simple specification test, in: Review of Financial Studies, Vol. 1 (1), 41-66

Long, K.R., Van Gosen, B.S., Foley, N.K., and Cordier, D. (2010): The principal rare earth elements deposits of the United States – a summary of domestic deposits and a global perspective, in: U.S. Geological Survey Scientific Investigations Report 2010 – 5220, http://pubs.usgs.gov/sir/2010/5220/

Massachusetts Institute of Technology's Energy Initiative/APS Panel on Public Affairs/The Materials Research Society (2010): Critical elements for new energy technologies, An MIT Energy Initiative Workshop Report

McCarty, D., and Casey, S. (2015): Bankruptcy as rare-earth prices drop, Bloomberg, http://www.bloomberg.com/news/articles/2015-06-25/molycorp-files-for-bankruptcy-proposes-debt-restructuring-plan

Miller, J.W., and Zheng, A. (2015): Molycorp files for bankruptcy protection, Wall Street Journal, http://www.wsj.com/articles/SB10907564710791284872504581069270334872848

Morrison, W.M., and Tang, R. (2012): China's rare earth industry and export regime: Economic and trade implications for the United States, Congressional Research Service Report R42510, http://www.fas.org/sgp/crs/row/R42510.pdf

Müller, M., Schweizer, D., and Seiler, V. (2016): Wealth effects of rare earth prices and China's rare earth elements policy, in: Journal of Business Ethics, forthcoming

Newey, W.K., and West, K.D. (1987): A simple, positive semi-definite, heteroscedasticity and autocorrelation consistent covariance matrix, in: Econometrica, Vol. 55 (3), 703-708

Piotroski, J.D., and Roulstone, D. T. (2004): The influence of analysts, institutional investors, and insiders on the incorporation of market, industry, and firm-specific information into stock prices, in: The Acccounting Review, Vol. 79 (4), 1119-1151

Politis, D.N., and Romano, J.P. (1995): Bias-corrected nonparametric spectral estimation, in: Journal of Time Series Analysis, Vol. 16 (1), 67-104

Politis, D.N., and White, H. (2004): Automatic block-length selection for the dependent bootstrap, in: Econometric Reviews, Vol. 23 (1), 53-70

Poterba, J.M., and Summers, L.H. (1988): Mean reversion in stock prices: Evidence and implications, in: Journal of Financial Economics, Vol. 22 (1), 27-59

Reuters (2015): Rare earths miner Molycorp files for Chapter 11 bankruptcy, http://www.reuters.com/article/2015/06/25/us-molycorp-bankruptcy-idUSKBN0P50QS20150625

Routledge, B.R., Seppi, D.J., and Spatt, C.S. (2000): Equilibrium forward curves for commodities, in: Journal of Finance, Vol. 55 (3), 1297-1338

Shen, H. (2014): China to add rare earths to the Shanghai Futures Exchange, http://investorintel.com/rare-earth-intel/china-plans-add-rare-earths-shanghai-futures-exchange/

Shleifer, A., and Vishny, R.W. (1994): Politicians and firms, in: Quarterly Journal of Economics, Vol. 109 (4), 995-1025

Skinner, D.J. (1989): Options markets and stock return volatility, in: Journal of Financial Economics, Vol. 23 (1), 61-78

Tan, Y., Tian, X., Zhang, X., and Zhao, H. (2015): The real effects of privatization: Evidence from China's split share structure reform, Available at SSRN: http://ssrn.com/abstract=2481838

TMR (Technology Metals Research) (2014): The WTO panel report on Chinese rare earth exports, http://www.techmetalsresearch.com

Tse, P.-K. (2011): China's rare-earth industry, U.S. Geological Survey Open-File Report 2011 – 1042, http://pubs.usgs.gov/of/2011/1042/

Tufano, P. (1998): The determinants of stock price exposure: Financial engineering and the gold mining industry, in: Journal of Finance, Vol. 53 (3), 1015-1052

USGS (2002): Rare earth elements – critical resources for high technology. U.S. Geological Survey Fact Sheet 087 – 02, http://pubs.usgs.gov/fs/2002/fs087-02/

Van Gosen, B.S., Verplanck, P.L., Long, K.R., Gambogi, J., and Seal, R.R. II (2014): The rare-earth elements – Vital to modern technologies and lifestyles, U.S. Geological Survey Fact Sheet 2014 – 3078, http://pubs.usgs.gov/fs/2014/3078/

Whang, Y.-J., and Kim, J. (2003): A multiple variance ratio test using subsampling, in: Economics Letters, Vol. 79 (2), 225-230

Wright, J.H. (2000): Alternative variance-ratio tests using ranks and signs, in: Journal of Business & Economic Statistics, Vol. 18 (1), 1-9

WSJ (2015): World news: China ends quotas on rare-earth exports, in: Wall Street Journal (Eastern Edition), 06 January 2015, A.6

WTO (World Trade Organization) (2014a): China files notice of appeal in rare earth disputes, http://www.wto.org/english/news_e/news14_e/ds432_433apl_25apr14_e.htm

WTO (World Trade Organization) (2014b): China – measures related to the exportation of rare earths, tungsten and molybdenum, dispute settlement: Dispute DS431, http://www.wto.org/english/tratop_e/dispu_e/cases_e/ds431_e.htm#bkmk431abr

WTO (World Trade Organization) (2014c): China – measures related to the exportation of rare earths, tungsten, and molybdenum. Reports of the Appellate Body, http://wto.org/english/tratop_e/dispu_e/431_432_433abr_e.pdf

WTO (World Trade Organization) (2014d): China – measures related to the exportation of rare earths, tungsten, and molybdenum. Reports of the Panel, http://www.wto.org/english/tratop_e/dispu_e/431_432_433r_e.pdf

WTO (World Trade Organization) (2014e): DSB adopts appellate body reports on China's rare earths measures, http://www.wto.org/english/news_e/news14_e/dsb_29aug14_e.htm

WTO (World Trade Organization) (2014f): US appeals rare earth panel report, http://wto.org/english/news_e/news14_e/ds431apl_08apr14_e.htm

Figure 1: Illustration of the Interactions Among Chinese REE Companies, the Chinese Government, Foreign REE Companies, and the WTO

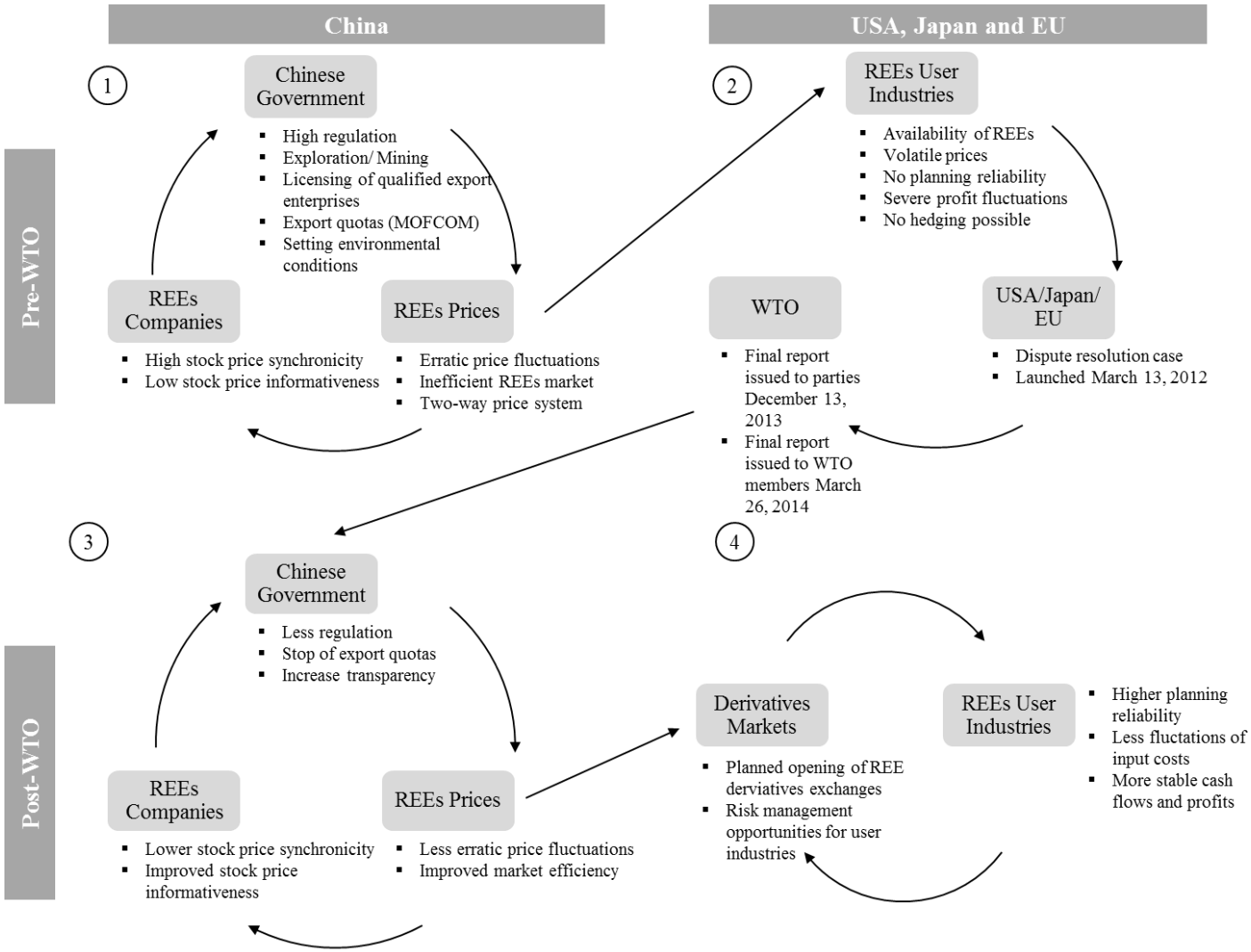
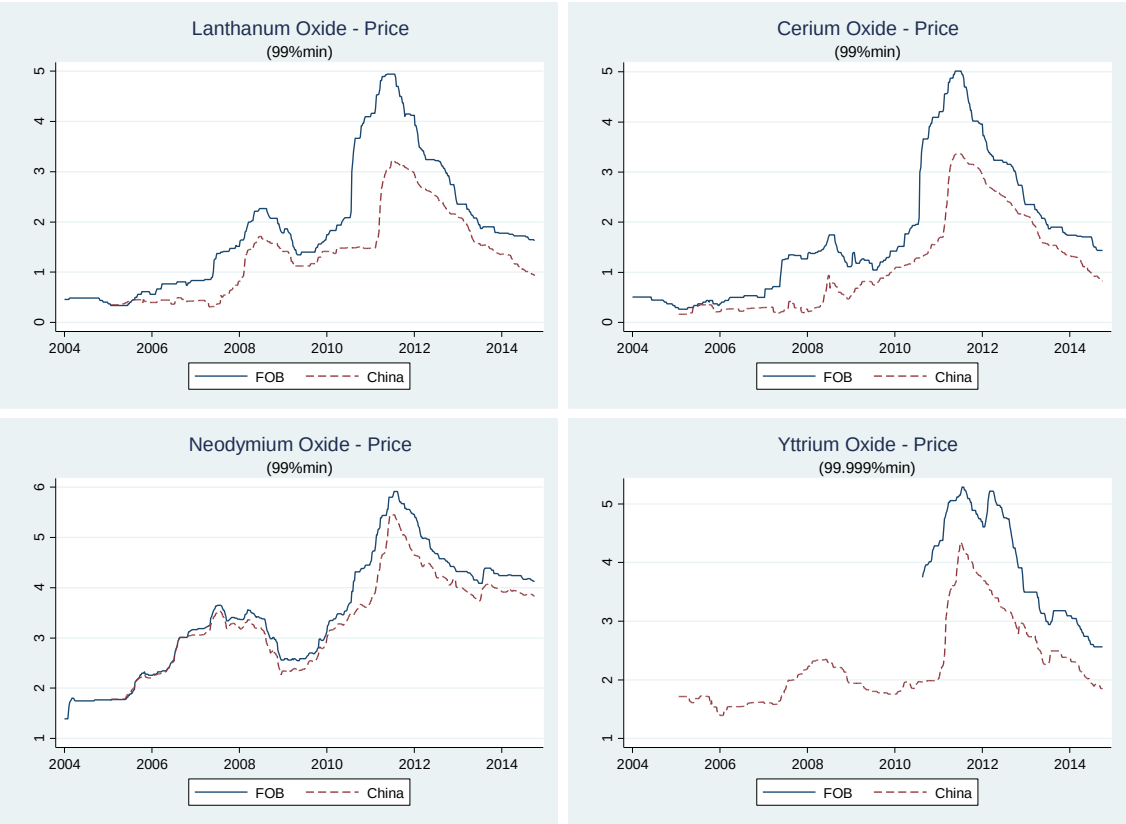


Figure 2: Price Developments of the Four Rare Earth Elements with the Highest Usage

This figure shows REE price and index developments over time. For REEs the foreign (FOB) and domestic (China) price developments are shown. All data come from the Asian Metals database. For FOB prices of La, Ce, and Nd, the time series begins in January 2004. For FOB prices of Y, the time series begins in August 2010. For China prices of Y, La, Ce and Nd, the time series begins in January 2005. All time series end in September 2014. Panel A shows the prices for individual REEs; panel B shows the prices for total usage for an equally weighted and usage-weighted REE index comprised of the four elements (about 90% of all REE usage, with cerium at 32.93%, lanthanum 30.16%, neodymium 17.84%, and yttrium 9.06%). Usage statistics: Own calculations based on Goonan (2011). % min below the respective REEs in the graphs denotes purity. Figure is similar to Müller, Schweizer, and Seiler (2015), but only for the four elements that we consider in the further analyses and enter the index and for an extended period.

Panel A: Rare Earth Elements



Panel B: Rare Earth Elements Indices

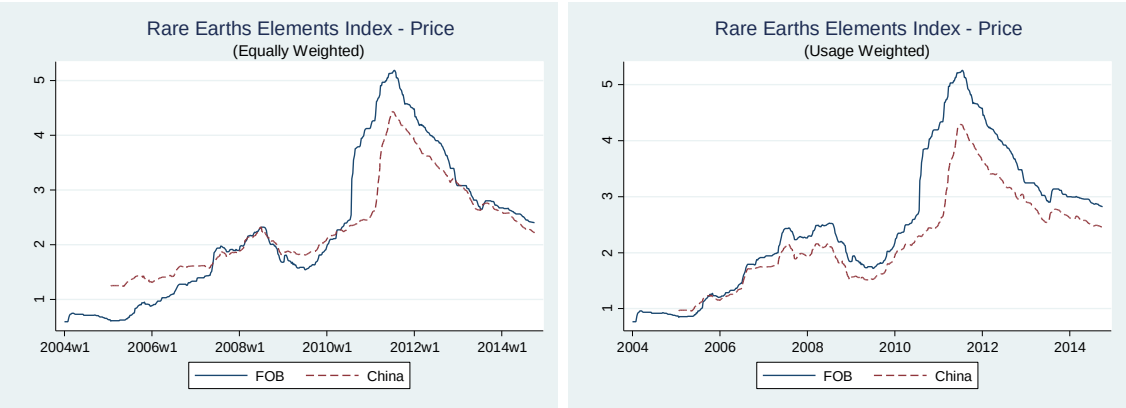
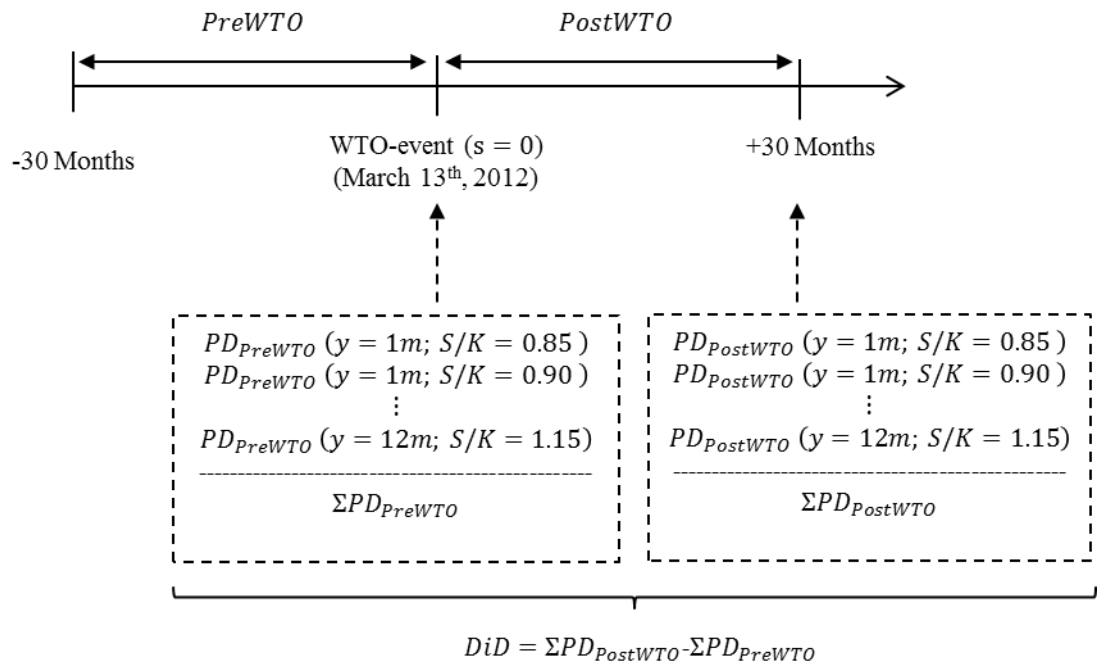


Figure 3: Illustration of the DiD and DDD Approaches

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 ($y = 1m$).

Panel A: Static DiD Approach



(continued)

Figure 3: Illustration of the DiD and DDD Approaches—*continued*

Panel B: Consecutive DDD Approach

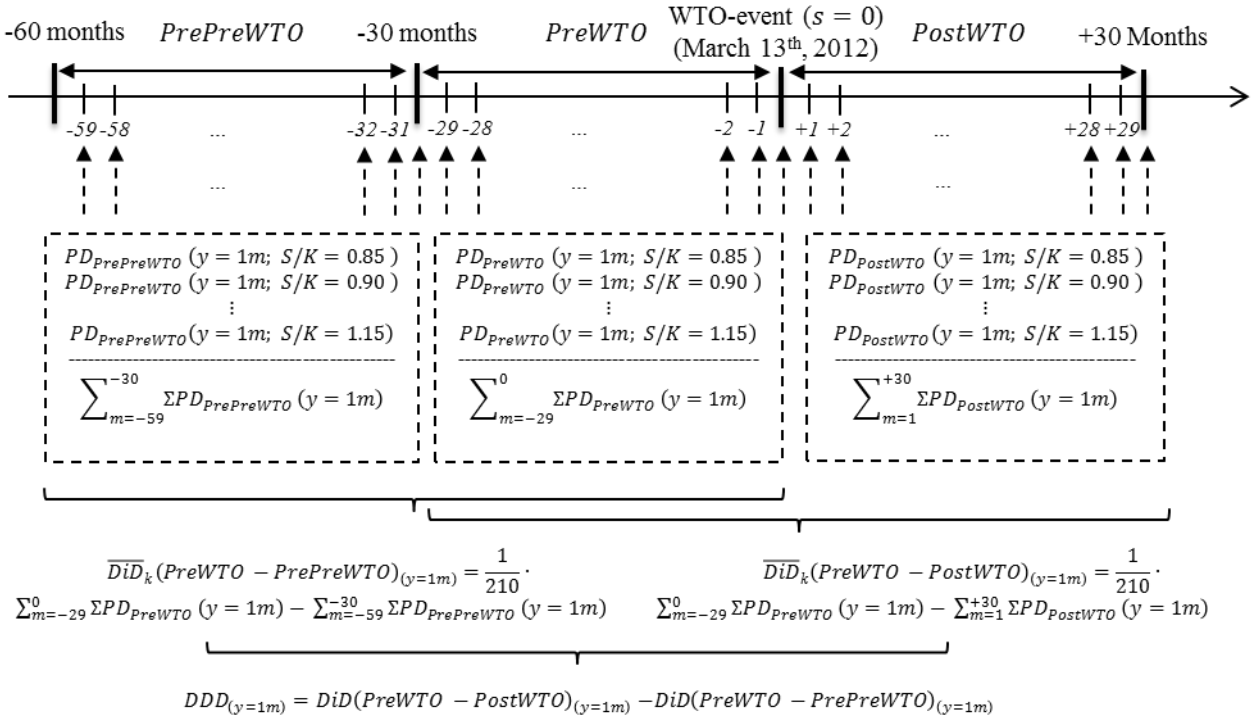
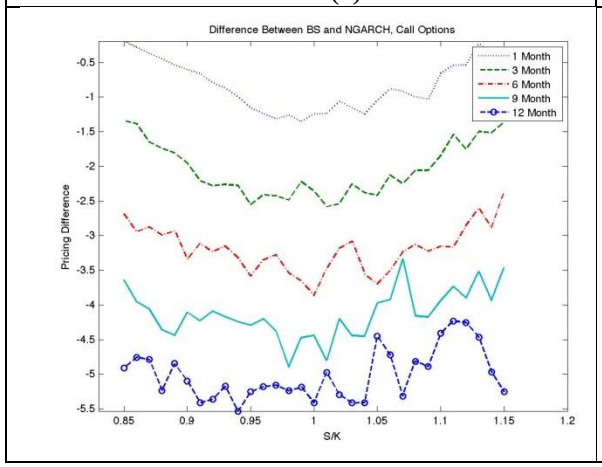
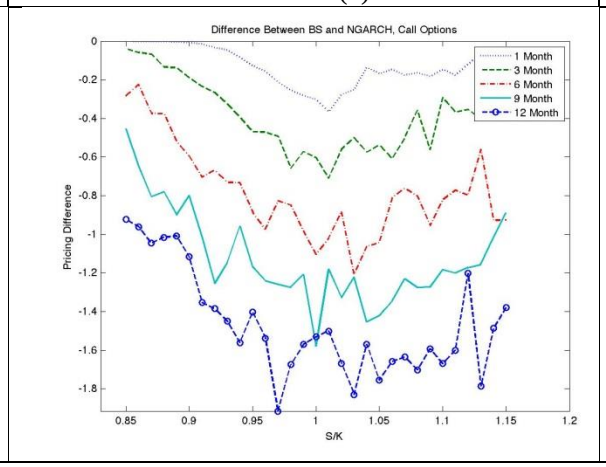
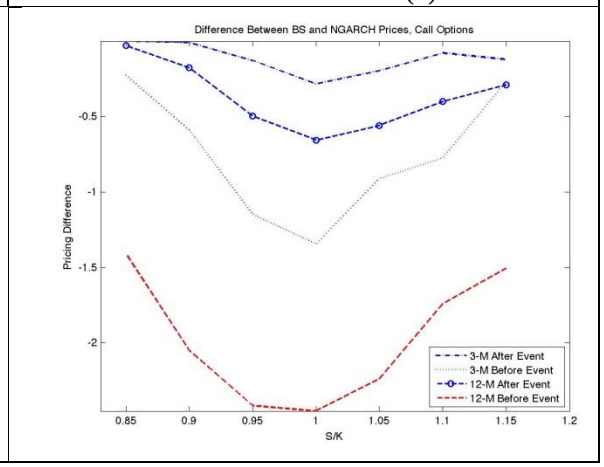
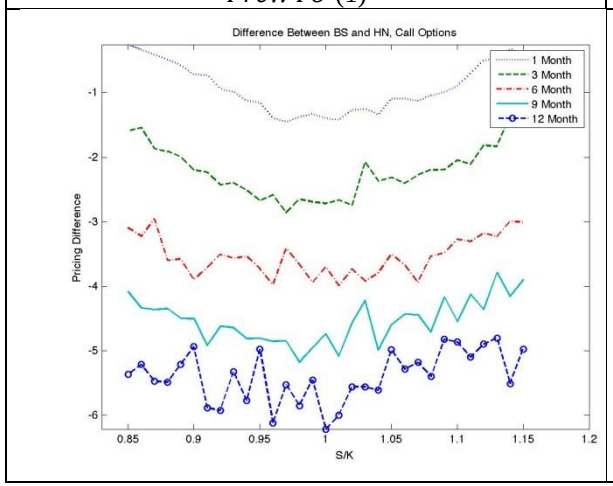
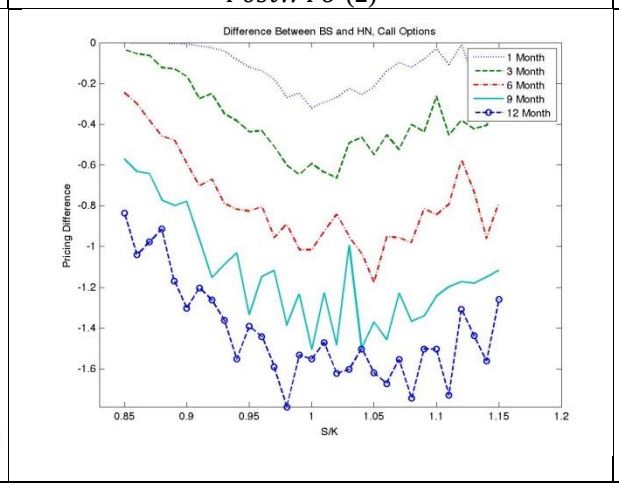
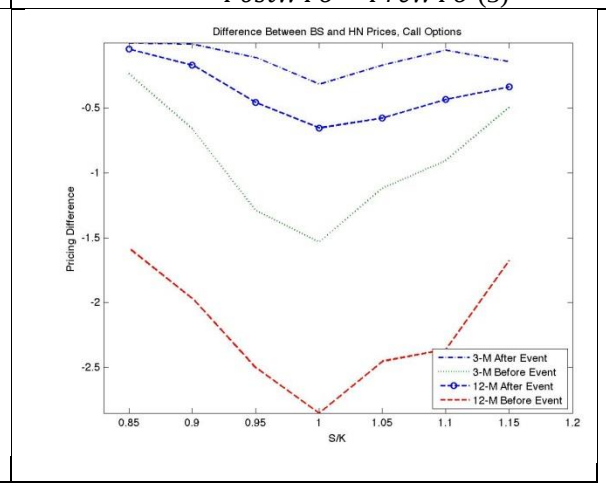


Figure 4: Option Price Differences between Option Pricing Models before and after the WTO Event (Lanthanum; Call Options)

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)

Figure 4: 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)

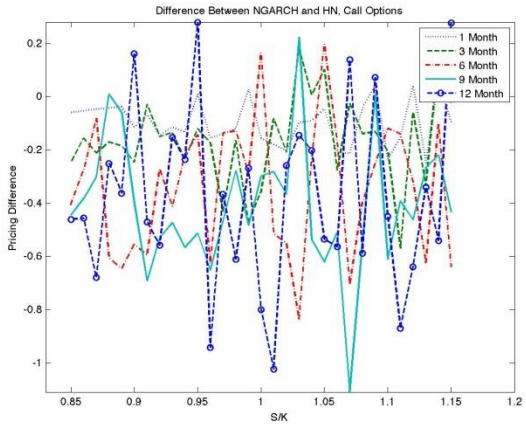
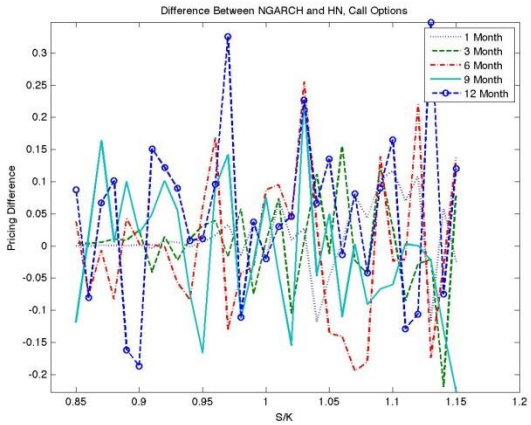
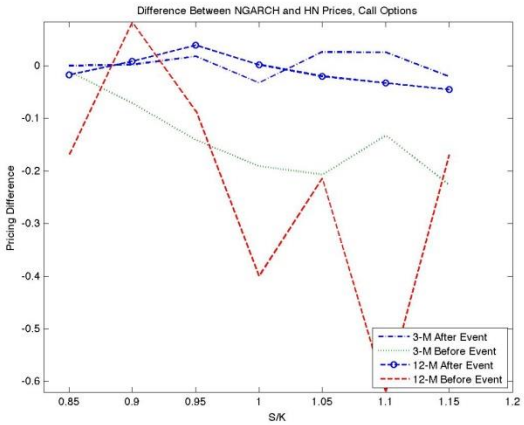
$PreWTO$ (1)	$PostWTO$ (2)	$PostWTO - PreWTO$ (3)
		

Table 1: Descriptive Statistics of REE Prices—Weekly Base Observation Period

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.

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

Table 2: Variance Ratio Tests—Weekly Base Observation Period

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.

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: Variance Ratio Tests—Weekly Base Observation Period—*continued*

Panel B			Number q of base observations aggregated to form variance ratio			
Element			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%*

Table 3: Multiple Structural Change Test

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 $M = 5$.

			Break Dates				
Element		sup$F(\ell+1\|\ell)$	$m=1$	$m=2$	$m=3$	$m=4$	$m=5$
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)*					

Table 4: Descriptive Statistics for Stock Price Synchronicity and Firm Characteristics

$SYNCH$ refers to the stock price synchronicity measure estimated using Equation (5) and transformed using Equation (4) for firms with A shares only. All other variables are defined as in appendix A4.

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
$\boldsymbol{Std(RoA)}$	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

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

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 (11). *t*-statistics are in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

		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)

Table 6: Correlation Matrix for Stock Price Synchronicity and Firm Characteristics

This table gives Pearson correlation coefficients for all variables presented in Table 4. All variables are considered in subsequent analyses (see appendix A4 for variable descriptions and calculation methods). * indicates correlations are statistically significant at least at the 5% level.

		(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

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

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 A4 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.

# *REECs*	Weighting		(1)	(2)	(3)	(4)	(5)	(6)
50		$PostWTO(\tau_2)\times REEC$	-0.170***					
			(-3.33)					
50		$PreWTO(\tau_1)\times REEC$	0.202***					
			(3.86)					
25		$PostWTO(\tau_2)\times REEC$		-0.146**				
	Equally			(-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		$PostWTO(\tau_2)\times REEC$				-0.155***		
						(-3.06)		
50		$PreWTO(\tau_1)\times REEC$				0.192***		
						(3.85)		
25		$PostWTO(\tau_2)\times REEC$					-0.251***	
	Usage						(-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: Change in Stock Price Synchronicity of REE Industry Companies after the WTO Event—Multivariate Evidence—*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

Table 8: Change in Stock Price Synchronicity of REE Industry Companies after the WTO Event—Multivariate Evidence (Alternative Estimation Strategies)

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.

Panel A: Two-Stage Least Squares with Newey-West Standard Errors

# *REECs*	Weighting		(1)	(2)	(3)	(4)	(5)	(6)
50		$PostWTO(\tau_2)\times REEC$	-0.183***					
			(-3.25)					
50		$PreWTO(\tau_1)\times REEC$	0.222***					
			(3.79)					
25	Equally	$PostWTO(\tau_2)\times REEC$		-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: Change in Stock Price Synchronicity of REE Industry Companies after the WTO Event—Multivariate Evidence (Alternative Estimation Strategies)—*continued*

Panel B: Random Effects Estimation								
# *REECs*	Weighting		(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

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

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 Figure 3 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.

		Call Options				**Put Options**			
		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
		PD	$\lvert BS - NGARCH \rvert$	$\lvert BS - HN \rvert$	$\lvert NGARCH - HN \rvert$	PD	$\lvert BS - NGARCH \rvert$	$\lvert BS - HN \rvert$	$\lvert NGARCH - HN \rvert$
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%***

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

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*.

$PreWTO$ Lanthanum

(1)	(2)	(3)	(4)	(5)
y	S/K	$\lvert BS - NGARCH \rvert$	$\lvert BS - HN \rvert$	$\lvert NGARCH - HN \rvert$
	0.85	0.22	0.28	0.06
	0.9	0.62	0.68	0.06
	0.95	1.15	1.19	0.04
1m	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
	0.85	1.36	1.57	0.21
	0.9	1.94	2.17	0.23
	0.95	2.44	2.49	0.04
3m	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
	0.85	2.83	2.81	0.01
	0.9	3.60	3.66	0.06
	0.95	3.52	3.37	0.16
6m	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
	0.85	3.92	4.41	0.49
	0.9	4.43	4.93	0.50
	0.95	4.58	5.36	0.78
9m	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
	0.85	4.70	5.49	0.79
	0.9	4.99	5.59	0.60
	0.95	5.13	5.40	0.27
12m	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

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

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*.

Cerium	$\lvert BS - NGARCH\rvert$			$\lvert BS - HN\rvert$			$\lvert NGARCH - HN\rvert$			Nobs
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	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	$\lvert BS - NGARCH\rvert$			$\lvert BS - HN\rvert$			$\lvert NGARCH - HN\rvert$			Nobs
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	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: Consecutive Triple-Difference (DDD) Approach Between Option Pricing Models Before and After the WTO Event—*continued*

Neodymium	$\lvert BS - NGARCH\rvert$			$\lvert BS - HN\rvert$			$\lvert NGARCH - HN\rvert$			Nobs
S/K	$\overline{DiD}\left(PreWTO - PrePreWTO\right)$	$\overline{DiD}\left(PreWTO - PostWTO\right)$	t-stat (DDD)	$\overline{DiD}\left(PreWTO - PrePreWTO\right)$	$\overline{DiD}\left(PreWTO - PostWTO\right)$	t-stat (DDD)	$\overline{DiD}\left(PreWTO - PrePreWTO\right)$	$\overline{DiD}\left(PreWTO - PostWTO\right)$	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	$\lvert BS - NGARCH\rvert$			$\lvert BS - HN\rvert$			$\lvert NGARCH - HN\rvert$			Nobs
S/K	$\overline{DiD}\left(PreWTO - PrePreWTO\right)$	$\overline{DiD}\left(PreWTO - PostWTO\right)$	t-stat (DDD)	$\overline{DiD}\left(PreWTO - PrePreWTO\right)$	$\overline{DiD}\left(PreWTO - PostWTO\right)$	t-stat (DDD)	$\overline{DiD}\left(PreWTO - PrePreWTO\right)$	$\overline{DiD}\left(PreWTO - PostWTO\right)$	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

ONLINE APPENDIX

Figure A1: Conditional Variance from GARCH(1,1) for REEs

This figure shows the conditional variance from a GARCH(1,1) process for the return series of the individual REEs. All data come from the Asian Metals database for the January 2005-September 2014 period. The estimated variance equations are as follows: Cerium: $\sigma_t^2 = 0.0502\epsilon_{t-1}^2 + 0.9560\sigma_{t-1}^2$. Lanthanum: $\sigma_t^2 = 0.1752\epsilon_{t-1}^2 + (-0.0890)\sigma_{t-1}^2$. Neodymium: $\sigma_t^2 = 0.4574\epsilon_{t-1}^2 + 0.5166\sigma_{t-1}^2$. Yttrium: $\sigma_t^2 = 0.2093\epsilon_{t-1}^2 + 0.7238\sigma_{t-1}^2$.

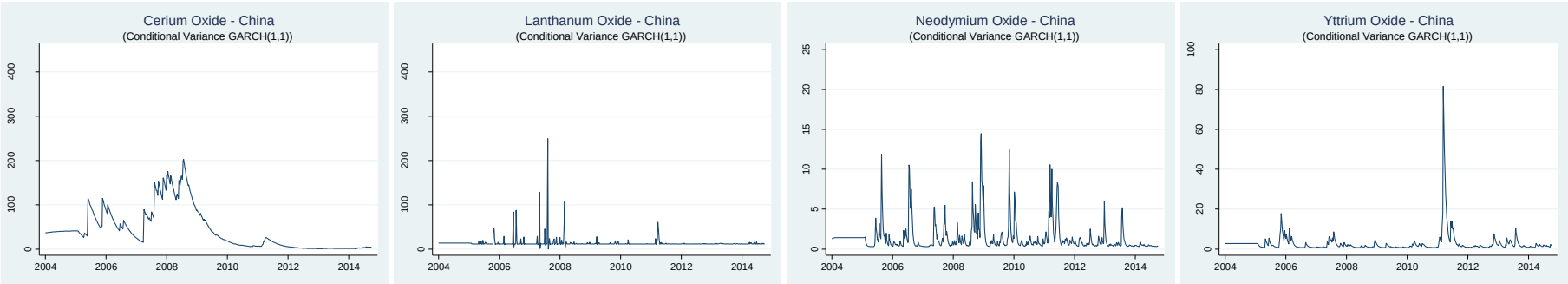


Figure A2: Google Trends Search
This figure shows the Google Trends index for the search term "Rare Earth" over the January 2004 to September 2014 period.

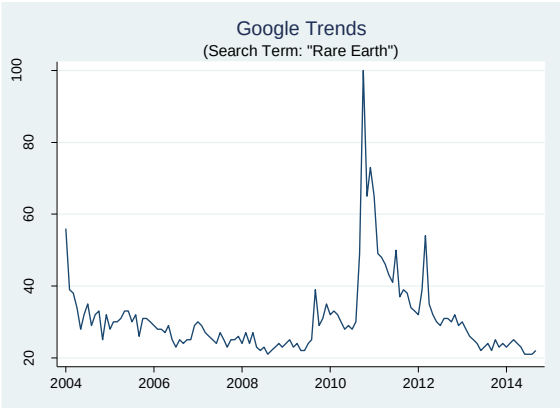


Table A1: Business Profiles of Identified REE Companies

This table gives an overview of the business profiles of the five firms with the highest correlations among the 12 (= 3 (τ=-1, -2, and pre-selection periods) × 2 (REE index based on FOB and China prices) × 2 (equally and usage-weighted REE indices) possible selection combinations (based on lowest rank-sum). Business profiles and additional information come from Reuters or Bloomberg.

Company	Profile	Additional Information
Huayi Electric Co., Ltd.	Huayi Electric Company Limited is principally engaged in the manufacture and distribution of wind power generation equipment and high voltage electrical appliances. The company's main products are high-and low-voltage power distribution products, including outdoor switches, medium- and low-voltage switch sets, switch components, high-voltage switches and substations, as well as wind power generation products, such as wind power generator sets. In addition, through its subsidiaries, the company is involved in providing first stage development and technical services for wind power stations, as well as research and development of software of wind power control systems. The company's products are distributed in both domestic and overseas markets.	
Hubei Biocause Pharmaceutical Co., Ltd.	Hubei Biocause Pharmaceutical Co., Ltd. develops and produces pharmaceuticals focusing on the treatment of colitis, high blood pressure, and infectious diseases. The company also manufactures and markets carbinol and polypropylene chemical products.	The company owns a 2.23% stake in Inner Mongolia Baotou Steel Union Co. Ltd., and a 0.68% stake in China Minmetals Rare Earth Co. Ltd., which are both active in the REE market.
Shanghai Electric Co., Ltd.	Shanghai Electric Group Company Limited is principally engaged in the design, manufacture, and distribution of electric power and industrial equipment. The company's main business is new energy, including the manufacture and sale of wind turbines and components and nuclear power equipment; efficient and clean energy, including the manufacture and sale of thermal power equipment and power transmission and distribution equipment; industrial equipment, including the manufacture and sale of elevators and motors; modern services, including the contracting of construction projects of thermal power and transmission and distribution projects, as well as other businesses. The company mainly operates in domestic and overseas markets.	
Xining Special Steel Co., Ltd.	Xining Special Steel Co., Ltd. is principally engaged in the smelting and rolling processing of special steel products, as well as coal industries and ore mining. The company's primary products include alloy structural steels, alloy tool steels, carbon structural steels, bearing steels, spring steels, carbon tool steels, stainless steels, raw coals, lime stones, refined ferrous powder, and cokes. The company operates in domestic and overseas markets.	

(continued)

Table A1: Business Profiles of Identified REE Companies—*continued*

Company	Profile	Additional Information
Xinjiang Ba Yi Iron & Steel Co., Ltd.	Xinjiang Ba Yi Iron & Steel Co., Ltd. is principally engaged in the smelting, rolling, and processing of iron and steel, as well as the production and distribution of pressing steel products. The company provides deformed steel bars, high-speed wire rods, hot-rolled plates, round steel, cold-rolled plates, galvanized plates, colour-coated plates, and medium plates, among others. The company distributes its products primarily in the domestic market, with the Xinjiang autonomous region as its major market.	The company is part of the Baosteel group, which is active in the production of steel and non-ferrous metals, including the processing of REEs.

Table A2: Variable Definitions and Descriptions

This appendix provides a detailed overview of the calculation methods and databases used here. Variable name means the name used in all tables and figures; database shows which database was used to obtain the information; variable ID is the name of the data item in the respective database, and description and calculation method describe how the variable was derived or calculated. CSMAR:TRADING refers to the CSMAR China Stock Market Trading Database, and CSMAR:FINANCIAL refers to the CSMAR China Stock Market Financial Statements Database.

Variable Name	Database	Variable ID	Description and Calculation Method
$SYNCH$	CSMAR:TRADING	See Equation (6)	Stock price synchronicity
MR	CSMAR:TRADING	CDRETWDTL	Volume-weighted daily aggregated market index return with reinvested cash dividends (total value-weighted), including all A shares (listed on either the Shanghai or Shenzhen stock exchange)
IR	CSMAR:TRADING	Based on DRETWD	Industry index return for the respective industry based on the six-digit industry classifications from *Industry Code A*. See necessary conditions below Equation (5), which must hold to enter the index
$REER$	Asian Metal Database	./.	REE indices that cover the four most important elements (Ce, La, Nd, and Y), calculated as usage- and equally weighted indices based on FOB (foreign) and China prices.
$PreWTO(\tau_1)$	./.	./.	Dummy variable that equals 1 for the pre-WTO period, and 0 otherwise
$PostWTO(\tau_2)$	./.	./.	Dummy variable that equals 1 for the post-WTO period, and 0 otherwise
$Top\ Gov$	CSMAR:TRADING	Based on NSHRSTT, NSHRTTL	Percentage of shares held by the state at the end of the fiscal year

(*continued*)

Table A2: Variable Definitions and Descriptions—*continued*

$Volume$	CSMAR:TRADING	Based on MNSHRTRD, NSHRTTL	Trading volume computed as the total number of shares traded in a year, divided by the total number of shares outstanding at the end of the fiscal year
$Size$	CSMAR:FINANCIAL	A001000000	Firm size computed as the log of total assets at the end of the fiscal year
$Leverage$	CSMAR:FINANCIAL	Based on A002000000, A001000000	Leverage computed as the total liabilities divided by total assets at the end of the fiscal year
$Std(RoA))$	CSMAR:FINANCIAL	Based on D000101000, A001000000;	Standard deviation of a firm's earnings stream measured by the standard deviation of a firm's return on assets (ROA) over the preceding four-year period, including the current year
M/B	CSMAR:TRADING, & CSMAR:FINANCIAL	Based on MSMVTTL, A003000000	Market-to-book ratio, computed as the total market value of equity, divided by total shareholder equity at the end of the fiscal year
Ind_Num	CSMAR:FINANCIAL	./.	Natural log of the number of firms in the industry to which a firm belongs. See necessary conditions below Equation (5), which must hold to enter in the equation
Ind_Size	CSMAR:FINANCIAL	Based on A001000000	Industry size, measured as the log of the year-end total assets of all sample firms in the industry to which a firm belongs. See necessary conditions below Equation (5), which must hold to enter in the equation

Table A3: Descriptive Statistics of REE Returns

This table shows the mean (in %), standard deviation (in %), skewness, kurtosis, minimum (in %), and maximum (in %) of the weekly REE log returns for the January 2, 2004-September 30, 2014 period, based on FOB (foreign) and China (domestic) prices (end-of-week) (panel A) and related indices (panel B). Price data comes from the Asian Metal database; all calculations based on USD/kg.

Element		Mean	Std. Dev.	Skewness	Kurtosis	Min	Max
Panel A: Light Rare Earth Elements – Oxides							
Lanthanum	FOB	0.2005	5.5216	6.1584	87.4396	-19.70	79.54
	China	0.1421	3.8554	4.2875	33.1932	-12.24	32.73
Cerium	FOB	0.2185	6.1732	6.8522	97.0774	-23.51	91.25
	China	0.2241	3.5074	2.7797	17.7017	-9.93	26.06
Neodymium	FOB	0.4788	4.0202	2.0750	14.1507	-15.12	23.86
	China	0.3878	3.6276	0.9166	8.9949	-14.14	18.40
Yttrium	FOB	na	na	na	na	na	na
	China	0.1865	6.0704	2.3510	35.4982	-34.05	43.33
Panel B: Rare Earth Indices – Oxides							
Equally Weighted REE Index	FOB	0.3233	4.4089	6.6652	94.0005	-14.71	65.79
	China	0.1919	2.8956	3.0077	22.5206	-8.66	24.40
Usage-Weighted REE Index	FOB	0.3674	4.0667	5.1808	62.0995	12.78	53.98
	China	0.2931	2.9463	1.8093	12.7585	-9.46	20.13

Table A4: Change in Stock Price Synchronicity of REE Industry Companies after the WTO Event—Multivariate Evidence (China REE Index)

This table replicates Tables 7 and 8, except it uses the equally weighted China REE Index for the selection of REECs instead of the FOB REE Index. Examining 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.

		OLS			**Two-Stage Least Squares with Newey-West Standard Errors**			**Random Effects Estimation**		
# REECs		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
50	$PostWTO(\tau_2)\times REEC$	-0.170***			-0.183***			-0.131**		
		(-3.13)			(-2.92)			(-2.10)		
50	$PreWTO(\tau_1)\times REEC$	0.015			0.033			-0.013		
		(0.26)			(0.58)			(-0.19)		
25	$PostWTO(\tau_2)\times REEC$		-0.133*			-0.149*			-0.090	
			(-1.66)			(-1.70)			(-1.02)	
25	$PreWTO(\tau_1)\times REEC$		0.022			0.042			0.033	
			(0.29)			(0.55)			(0.32)	
10	$PostWTO(\tau_2)\times REEC$			-0.080			-0.091			-0.017
				(-0.59)			(-0.62)			(-0.13)
10	$PreWTO(\tau_1)\times REEC$			0.030			0.048			0.048
				(0.31)			(0.57)			(0.32)
	Control Variables	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
	Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
	Observations	4,307	4,307	4,307	4,307	4,307	4,307	4,307	4,307	4,307

Appendix A5: Black-Scholes (1973) Simulation with Moving Block Bootstrap

We use a moving block bootstrap methodology to simulate the sample distribution. We thus follow Politis and White (2004) in terms of selecting the optimal block length for the dependent bootstrap. The optimal block size is chosen by $b_{opt} = \left(\frac{2G^2}{D}\right)^{1/3} N^{1/3}$. Both G and D involve parameters $\sum_{k=-\infty}^{\infty}|k|R(k)$, $\sum_{k=-\infty}^{\infty}R(k)$, and $g(\omega) = \sum_{k=-\infty}^{\infty}R(k)\cos(\omega k)$. As Politis and Romano (1995) suggest, $\sum_{k=-\infty}^{\infty}|k|R(k)$ is estimated by $\sum_{k=-M}^{M}\lambda(k/M)\,|k|\hat{R}(k)$, where $\hat{R}(k) = N^{-1}\sum_{i=1}^{N-|k|}(X_i - \bar{X}_N)\left(X_{i+|k|} - \bar{X}_N\right)$. $\lambda(t)$ is a function with symmetrically trapezoidal shape around zero:

$$\lambda(t) = \begin{cases} 1 & if\ |t| \in [0, 1/2] \\ 2(1-|t|) & if\ |t| \in [1/2, 1]. \\ 0 & otherwise. \end{cases}$$

Similarly, $g(\omega) = \sum_{k=-\infty}^{\infty}R(k)\cos(\omega k)$ could also be estimated by $g(\hat{\omega}) = \sum_{k=-M}^{M}\lambda(k/M)\hat{R}(k)\cos(\omega k)$. With the predefined functions, we would arrive at $\hat{G} = \sum_{k=-M}^{M}\lambda(k/M)\,|k|\hat{R}(k)$, $\hat{D} = 4\hat{g}^2(0) + \frac{2}{\pi}\int_{-\pi}^{\pi}(1+\cos\omega)\,\hat{g}^2(\omega)d\omega$. Then, the estimator for the block size choice would be: $\hat{b}_{opt} = \left(\frac{2\hat{G}^2}{\hat{D}}\right)^{1/3} N^{1/3}$.

It is important to choose the optimal bandwidth M. As suggested by Politis and Romano (1995), practically, we would look for the smallest integer $\hat{m}$, with the condition that $\hat{R}(k)$ is not significantly different from zero while $k > \hat{m}$; then $M = 2\hat{m}$.

Recent discussion papers

2017-05	Juliane Proelss Denis Schweizer Volker Seiler	Do Announcements of WTO Dispute Resolution Cases Matter? Evidence from the Rare Earth Elements Market
2017-04	Yuanhua Feng Thomas Gries	Data-driven local polynomial for the trend and its derivatives in economic time series
2017-03	Sonja Brangewitz Behnud Mir Djawadi Angelika Endres Britta Hoyer	Network Formation and Disruption - An Experiment Are efficient networks too complex?
2017-02	Volker Seiler Evgenia Taach	Efectiveness of Official Development Assistance – Further Evidence from DAC Countries
2017-01	Nico Stöckmann	“The Impact of Refugees on Per Capita Income” A Gravity Model Approach
2016-06	Jochen Manegold	Stackelberg Competition among Intermediaries in a Differentiated Duopoly with Product Innovation
2016-05	Republished as CIE Working Paper 2017-05	
2016-04	Marlon Fritz Thomas Gries Yuanhua Feng	Growth Trends and Systematic Patterns of Booms and Busts – Testing 200 Years of Business Cycle Dynamics-
2016-03	Thomas Gries Claus-Jochen Haake	An Economic Theory of ‘Destabilization War’
2016-02	Daniel Kaimann Nadja Maraun Joe Cox	Identifying preferences and heterogeneity of consumers in multiplayer video games
2016-01	Republished as CIE Working Paper 2016-05	
2015-05	Sonja Brangewitz Claus-Jochen Haake Philipp Möhlmeier	Strategic Formation of Customer Relationship Networks
2015-04	Sonja Brangewitz Jochen Manegold	Competition and Product Innovation of Intermediaries on a Differentiated Duopoly
2015-03	Achim Voß Mark Schopf	Lobbying over Exhaustible-Resource Extraction
2015-02	Maximilian Müller Denis Schweizer Volker Seiler	Wealth Effects of Rare Earth Prices and China`s Rare Earth Elements Policy
2015-01	Yuanhua Feng Chen Zhou	An iterative plug-in algorithm for realized kernels
2014-13	Republished as CIE Working Paper 2015-02	
2014-12	Sonja Brangewitz Behnud Djawadi René Fahr Claus-Jochen Haake	Quality Choices and Reputation Systems in Online Markets – An Experimental Study
2014-11	Daniel Kaimann Joe Cox	The Interaction of Signals: A Fuzzy set Analysis of the Video Game Industry
2014-10	Thomas Gries Ha Van Dung	Institutional environment, human capital, and firm growth: Evidence from Vietnam