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RESEARCH ARTICLE



Return dynamics during periods of high speculation in a thinly traded commodity market

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Abstract

This article studies the effects of speculation in a thinly traded commodity futures market, paying particular attention to periods characterized by highspeculative activity of long-short speculators. Using the speculation ratio as a daily measure for long-short speculation, we employ generalized autoregressive conditional heteroscedasticity regressions to study its impact on return dynamics. Our results for the Chicago Mercantile Exchange feeder cattle futures market suggest that futures returns are predominantly explained by fundamentals, but their volatility is significantly driven by the speculation ratio. This relationship holds for periods of high- and low-speculative activity alike.

K E Y W O R D S

commodity markets, return dynamics, speculation, thinly traded markets

1 | INTRODUCTION

Both in the scientific literature and the general public, the reason for the surge of commodity prices throughout the 2000s has been the subject of considerable debate. A number of public commentators claimed that the price hikes were caused by various financial institutions increasing their investments in commodity futures. Most prominently, the so-called Masters (2008) hypothesis asserts that in particular long-only positions by commodity index funds drove up commodity prices throughout this period of time. However, the debate on the validity of the Masters hypothesis seems mostly settled, with numerous studies, for example, by Sanders, Irwin, and Merrin (2010), Irwin and Sanders (2012), Büyükşahin and Harris (2011) or Hamilton and Wu (2015), and Brunetti, Büyükşahin, and Harris (2016), rejecting the hypothesis's central claim.

Yet, comparatively little attention has been paid to the role of traditional long-short speculators. One strand of the literature examining this trader type has analyzed the direction of causality between changes in the positions of these traders and changes in commodity prices. Büyükşahin and Harris (2011) and Alquist and Gervais (2013) both study the market for crude oil futures and use trader position data from the Commodity Futures Trading Commission (CFTC). Whereas the former use highly disaggregated data at daily frequency, the latter rely upon more aggregated, weekly data that are publicly available. Neither find evidence for the notion that changes in trader positions lead price changes. Only Robles, Torero, and vonBraun (2009), who study four agricultural markets, find some support for the connection between changes in trader positions and changes in prices. However, the authors' findings are limited to a small number of rolling-window subsamples which are drawn from monthly data and thus contain very few observations.

Another stream of the literature has investigated the case of long-short speculation using different generalized autoregressive conditional heteroscedasticity (GARCH) models, which, in addition to investigating effects on returns,

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also allow for examining how speculation affects the volatility of returns. Manera, Nicolini, and Vignati (2013a) find that in nine fuel and agricultural commodity markets, the impact of speculation, as measured by Working's (1960) *T*, on returns is either not statistically significant or significantly negative. Further studies by Manera, Nicolini, and Vignati (2013b, 2016) show that in different fuel and agricultural markets, various speculation measures based on weekly CFTC trader position data also have a significantly negative impact on the volatility of futures returns. This implies that speculation exerts, contrary to public opinion, a stabilizing effect on commodity returns. These findings are echoed by the results of Brunetti et al. (2016), who study the impact of changes in trader positions on realized volatility. The authors use daily disaggregated position data from the CFTC and show that the activities of swap dealers and, in particular, hedge funds reduce the volatility of the futures markets for crude oil, natural gas, and corn.

Conversely, when measuring speculation using the speculation ratio of Garcia, Leuthold, and Zapata (1986), which divides trading volume by open interest, Manera et al. (2013b) find a positive impact of speculation on return volatility. This finding is in line with earlier results of Streeter and Tomek (1992) and Du, Yu, and Hayes (2011), who also use the speculation ratio to measure the impact that speculators have on volatility in the markets for soybeans and crude oil, respectively. The difference in the results of these studies based on the ratio of trading volume to open interest and those based on trader position data is that the latter are unable to capture trading activities that occur throughout the day.

Lastly, Etienne, Irwin, and Garcia (2015) study how speculation explains the occurrence of periods characterized by explosive prices. They find that increases in the positions of noncommercial traders significantly increase the probability for the occurrence of positive price bubbles and significantly decrease the probability for negative price bubble events.¹

Our investigation extends this earlier work in three ways: First, unlike most studies, this article analyzes the effects of long-short speculation using daily data. As neither scalpers nor day traders hold positions for more than a few hours, daily observations promise to capture the effects of long-short speculation more accurately than lower frequency data. When measuring speculation, we rely on the speculation ratio of Garcia et al. (1986). This ratio measures speculation by linking trading volume to the number of open positions in the market, an idea that dates at least back to Peck (1981). The reasoning for using the speculation ratio as a measure for speculative activity is that long-short speculators, as opposed to hedgers, will prefer to hold their positions for relatively short periods of time. Long-short speculators, such as day traders, take intraday positions based on expectations of how prices will move over the next minutes or hours, whereas scalpers open and close positions almost instantaneously. Naturally, the profit margins of such trades are very small, which is why these traders must trade relatively large volumes to be profitable. In the process, they provide liquidity and immediacy to the market, which is why such traders are said to make markets (e.g., Büyükşahin & Harris, 2011; Du et al., 2011).

Second, this article identifies time periods characterized by high levels of long–short speculation to particularly study market conditions under which one would expect to observe the most powerful effects of long–short speculation. The reason for this is that changes in the positions of hedgers are typically more strongly motivated by changes in fundamentals than those of speculators. Consequently, increases in the relative market dominance of speculators will add to the risk of prices moving away from their fundamental level and instead fueling bubbles. To single out such time periods, we adopt a standard Markov-switching approach (Hamilton, 1989).

Third, we focus on a rather illiquid market. Again, the idea is that speculators' impact should be greater if the number of hedgers is small and if the underlying commodity is nonstorable. In this case, commercial traders cannot adjust inventories to cushion against undesirable price movements. Whereas the majority of research focuses on relatively large commodity futures markets with several hundred thousand open positions at the end of each trading day, this article considers the effects of speculation in the market for Chicago Mercantile Exchange (CME) feeder cattle. This market is arguably the least liquid agricultural commodity market included in the CFTC's Supplemental Commitments of Traders (SCOT) reports. Moreover, livestock can, unlike other commodities, generally not be stored.²

If, despite these conditions, classical long–short speculation does not significantly drive returns or volatility in this market, speculators are unlikely to influence return dynamics in more favorable market environments. In contrast, if speculation amplifies or reinforces return dynamics, regulators should be alert for allowing such unfavorable market conditions to develop. Lastly, should speculation exhibit negative, that is, stabilizing, effects on returns, and volatility,

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¹For a detailed literature survey regarding the role of speculation during the financialization of commodity markets, refer to Boyd, Harris, and Li (2018).

²Over the last 10 years, the average daily trading volume of CME feeder cattle futures amounted to around 8,000 contracts. During the same period, other livestock futures, for example, CME lean hogs and live cattle futures, featured average trading volumes of around 40,000 and 50,000 contracts/day.

regulators should refrain from reining in speculative trading. Our results show that while returns are predominantly driven by fundamentals, the speculation ratio does exert economically and statistically significant effects on volatility. This effect is independent of the underlying regime of high or low long-short speculation in the market, that is, the volatility-increasing effect of the speculation ratio is not greater or smaller in one of the regimes than during the other.

The remainder of this article is structured as follows: First, we explain the econometric methodology then the data of our analysis. Thereafter, the subsequent sections present the key results of this article and discuss several robustness exercises, after which we conclude.

2 | ECONOMETRIC METHODOLOGY

Our methodology comprises a two-step approach. First, we use weekly CFTC data and a Markov-switching model (Hamilton, 1989) to determine periods of increased activity by long–short speculators. Second, we use daily data to assess the effects of long–short speculation on returns and volatility by employing a GARCH model (Bollerslev, 1986), while incorporating the results of the first step in the form of a time dummy. This allows for studying the impact of increased long–short speculation on futures return dynamics.³

We determine regimes of high long-short speculation based on a constant-only Markov-switching regression model. A measure S_t of speculation in week t can therefore be in either one of two states, namely high or low. The two states are unobservable and follow a Markov process. Hence, S_t can be described by the model:

$$S_t = \mu_{s_t} + \varepsilon_t,\tag{1}$$

where μ_{s_t} resembles the state-dependent intercept term and ε_t denotes random disturbances. If week *t* is characterized by high-speculative activity, then $\mu_{s_t} = \mu_h$. Conversely, if week *t* is a period of low speculation, then $\mu_{s_t} = \mu_l$.

In opposition to autoregressive models, the dynamic regression model of Equation (1) allows for quick changes in the level of S_t . We estimate Equation (1) using a quasi-Newton algorithm within a standard maximum-likelihood routine. On the basis of the estimation results, we compute for each week the one-step ahead prediction for the probability of being in the state of increased speculative activity. After that, we define a daily regime dummy D_t equal to one if in that week the predicted probability of being in the high-speculation state exceeds the sample mean probability of being in this state. Otherwise the dummy is set to zero.

Given the regime dummy, the effects of long-short speculation on returns and volatility are examined using daily data and the following GARCH(1, 1) model:

$$r_{t} = \alpha_{0} + \alpha_{1}r_{t-1} + \sum_{i=1}^{4} \beta_{i}X_{i,t} + \beta_{5}R_{t-1} + \beta_{6}D_{t-1} + \beta_{7}I_{t-1} + \varepsilon_{t},$$
(2a)

$$\sigma_t^2 = \gamma_0 + \gamma_1 \varepsilon_{t-1}^2 + \gamma_2 \sigma_{t-1}^2 + \sum_{i=1}^4 \delta_i X_{i,t} + \delta_5 R_{t-1} + \delta_6 D_{t-1} + \delta_7 I_{t-1},$$
(2b)

where r_t denotes returns and σ_t^2 their conditional variance. Returns are explained by previous period returns r_{t-1} and a number of explanatory regressors. These consist of a set of macroeconomic controls contained in $X_{i,t}$, the lagged speculation ratio R_{t-1} to proxy for speculative activity, the lagged regime dummy D_{t-1} and an interaction term $I_{t-1} = R_{t-1} \cdot D_{t-1}$. The use of lagged terms avoids an endogeneity problem due to simultaneity. Lastly, ε_t denotes random disturbances.

Our analysis includes as macroeconomic controls the oil price, a stock return index, the riskless interest rate, and the exchange rate. Rising oil prices will typically increase the production and transportation costs of a large variety of physical goods, including agricultural commodities and thus curb their supply. Similarly, increases in a broad variety of stock returns generally reflect economic booms, which in turn are likely to boost the demand for commodities.

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³Although the CFTC data are only available at weekly frequency, it seems fair to assume that they provide the most accurate measure of speculative activity in the market. On the other hand, the speculation ratio of Garcia et al. (1986) can easily be computed from daily data on trading volume and open interest. By combining these two types of data, we intend to capture speculative activity as precisely as possible at daily frequency.

Therefore, we expect these two determinants to have a positive relationship with the futures return of feeder cattle. Increases in the riskless rate will increase producers' opportunity costs of holding inventories and thus most likely reduce the return of a commodity. This effect is, however, presumably lower for feeder cattle than for other commodities, as feeder cattle cannot be stored in the same way as other commodities. The exchange rate is also likely to have an inverse relationship with the concerned futures price, as an appreciation of the dollar exchange rate will cause a depreciation of the dollar denominated contract price. Empirical evidence for the inverse relationship between exchange rates and interest rates on the one hand and the prices of food and other commodities on the other is, for example, given by Akram (2009) and S.-L. Chen, Jackson, Kim, and Resiandini (2014).

The variance equation of the model consists of the autoregressive conditional heteroscedasticity (ARCH) term ε_{t-1}^2 , the GARCH term σ_{t-1}^2 and the same explanatory regressors incorporated in the mean equation. The parameter γ_1 resembles the ARCH effect, that is, how strongly the conditional variance reacts to new information arriving in the market, whereas the GARCH effect γ_2 measures how persistent shocks to the conditional variance are over time. For the returns process to be stationary in variance, it must hold that $\gamma_1 + \gamma_2 < 1$.

As this article seeks to study the effects of long-short speculators on returns and volatility, we are particularly interested in the six parameters associated with the speculation ratio R_{t-1} , the speculation regime dummy D_{t-1} and the interaction term I_{t-1} . The interpretation of β_5 and δ_5 is straightforward. If positive, speculation exerts destabilizing effects on prices by increasing returns and their volatility such that price changes become more severe. Conversely, if the coefficients are negative, speculative activity reduces returns and volatility.

In opposition to the Masters hypothesis, this latter point has been the main finding for the role of commodity index funds. This is in line with classical theory concerning the impact of speculation: The key argument for why speculators will exert stabilizing effects is that they are seen as arbitrageurs who buy when prices are below their fundamental value and sell if prices are above it. Otherwise, if speculators bought when prices were high or sold when prices were low, they would, on average, lose money and eventually exit the market (Friedman, 1953). Even if noise traders (Black, 1986; Kyle, 1985), that is, traders who respond to signals other than fundamentals, are present in the market, their impacts would always be offset by aggressive arbitrage trading from rational speculators (Fama, 1965). However, as demonstrated by De Long, Shleifer, Summers, and Waldman (1990a, 1990b), the actions of noise traders need not always be perfectly balanced by opposing trades from sophisticated arbitrageurs, either because the latter are risk averse or because they intend to profit from the actions of positive feedback traders. In these cases, speculation bears the potential for fueling bubbles by moving prices away from their fundamental values.

Parameters β_6 and δ_6 as well as β_7 and δ_7 capture nonlinear effects of speculation on the stability of prices. In the spirit of De Long et al. (1990a, 1990b), statistically significant estimates for these parameters would mean that sophisticated investors' risk aversion and trading strategy depend on the level of noise traders in the market. Parameters β_6 and δ_6 shift the constants in the two equations of the model. A positive value of β_6 suggests that returns are generally higher in periods of high-speculative activity, whereas a positive value of δ_6 would suggest that the volatility is elevated in these periods. Parameters β_7 and δ_7 alter the slope of the mean and variance equation with respect to the speculation ratio. Positive values of β_7 and δ_7 indicate that the activities of long–short speculators amplify the effects of an increase in the speculation ratio, whereas negative values of β_7 and δ_7 suggest that price or volatility-increasing effects of long–short speculation are reduced during periods of increased speculative activity.

3 | DATA

The CFTC issues a series of reports on the open positions taken by different types of traders in commodity futures markets.⁴ Among these are the SCOT reports, which are publicly available and have provided weekly data since January 2006 on market open interest, the number of traders and the aggregate positions of different trader types for 13 agricultural commodity futures markets.⁵ The SCOT reports distinguish between commercial traders, noncommercial traders, commodity index traders, and traders that do not report the nature of their futures transactions. Traders classified as commercial are typically seen as hedgers, whereas noncommercial traders are typically seen as speculators.

The behavior of nonreporting traders is, however, less clear. On the one hand, nonreported positions typically comprise those of traders that are too small to exceed the CFTC's reporting levels. These traders are presumably mainly

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⁴See Irwin and Sanders (2012) for a description and comparison of the different CFTC reports.

⁵Note that CBOT Soybean Meal futures were originally not part of the SCOT reports but were added to them in April 2013.

hedgers. On the other hand, Irwin and Sanders (2012) note that traders have incentives not to be classified as speculative traders, which is why a substantial fraction of nonreported positions could also stem from speculators. Table 1 summarizes the position data for the different markets included in the SCOT reports. The market for feeder cattle stands out for three reasons. First, it is the smallest both in terms of open interest and the number of traders, suggesting that it is the least liquid market. Second, the market for feeder cattle features very high open interest shares of noncommercial and nonreporting traders. Note that the share of nonreported positions is almost three times higher than that in the other markets.⁶ Therefore, the number of speculators is presumably rather high in this market.

Third, feeder cattle are nonstorable, which reduces the ability of commercial traders to cushion adverse price developments with inventory adjustments. Numerous papers have investigated the role of asset storability for the functioning of futures markets. For example, Skadberg and Futrell (1966) and Peck (1976) discuss various theoretical considerations for why futures contracts with nonstorable underlyings will typically behave differently from those contracts that are based on storable commodities. Studying a variety of different storable and nonstorable commodities, Yang, Bessler, and Leatham (2001) find that futures contracts based on nonstorable commodities, and in particular cattle contracts, perform worse in predicting future cash prices than futures contracts derived from storable commodities. Yang, Bessler, and Fung (2004) study the long-run relationship between the futures price and open interest, which can be seen as a proxy for hedging demand (see, e.g., N.-F. Chen, Cuny, & Haugen, 1995). They find evidence for the existence of such a relationship in futures markets with storable underlyings, but not for futures markets that are based on nonstorable assets.

Lastly, another branch of the literature focuses on how asset storability affects futures markets' ability to provide effective hedging opportunities. Baillie and Myers (1991), Yang and Awokuse (2003), and Zhang and Choudhry (2015) find that markets for nonstorable commodities feature far lower levels of hedging effectiveness than those for storable commodities. Consequently, if a considerable number of hedgers choose for any of these reasons not to participate in these futures markets, these markets will be particularly prone to experience forms of speculative pressures. Given this set of market characteristics, the effects of speculation should be particularly strong in the market for feeder cattle futures. Thus, we select the market for CME feeder cattle futures for our analysis.

As explained above, to determine high-speculation regimes, one must first compute an observable measure S_t for the level of speculative activity in week t. Obviously, this calculation hinges on the question of which trader type actually engages in speculative activity as opposed to who engages in hedging. Whereas the case is fairly straightforward for the positions of commercial and noncommercial traders, the literature has proposed several ways for how to allocate nonreported positions to hedging and speculative positions. Rutledge (1977) and Sanders et al. (2010) allocate nonreported positions to hedging and speculative positions in the same proportions that they observe for reported positions. Alquist and Gervais (2013) assume that all nonreporting traders are speculators, whereas Manera et al. (2013a) assume a 30/70 split among hedging and speculative positions.

These high values for the share of speculators among nonreporting traders are, however, presumably too high for the feeder cattle market. As noted by the CME Group (2017), the US cattle farming and feedlot industry is characterized by a large number of small and often family-owned operations. Consequently, the overwhelming majority of producers in this market will fall below the CFTC's reporting levels, resulting in a large number of hedgers among nonreporting traders. Therefore, we assume that only 20% of nonreported positions stem from speculative activity, whereas we assume the remaining 80% to stem from hedging.⁷ Hence, S_t is defined as

$$S_t = \frac{NCL_t + NCS_t + \alpha \cdot (NRL_t + NRS_t)}{2 \cdot MOI_t},$$
(3)

where NCL_t and NCS_t denote long and short positions of noncommercial speculators, respectively. Analogously, the positions of nonreporting traders are denoted NRL_t and NRS_t . Lastly, MOI_t resembles market open interest in week t, whereas $\alpha = 1/5$ for the baseline regressions.

Given S_t , we follow the procedure outlined in the methodology section to obtain the daily regime dummy D_t . This yields a total of 19 different high-speculation regimes, which on average last 11 weeks. The shortest of these regimes last

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⁶The analysis by Sanders et al. (2010) shows that the market for feeder cattle futures has also featured such high levels of speculative positions in earlier periods.

⁷Section 5 tests the sensitivity of our results with respect to changes in this assumption. Specifically, we explore how our results change when we assume 0%, 60%, or 80% of nonreported positions to stem from speculators.

TABLE 1 Open interest, number of traders, and open interest shares of different trader types in the CFTC SCOT reports

Future contract	Contract size	ΜΟΙ	No. traders	C (%)	NC (%)	I (%)	NR (%)
CBOT corn	5,000 bu	1,750,031	774	66	81	27	26
ICE sugar	112,000 lb	957,932	231	85	66	32	16
CBOT soybeans	5,000 bu	767,441	550	68	84	25	23
CBOT wheat	5,000 bu	521,595	404	50	90	42	18
CBOT soybean meal	100 tons	405,534	278	83	71	25	21
CBOT soybean oil	60,000 lb	369,395	259	81	74	27	17
CME live cattle	40,000 lb	358,571	395	60	83	32	25
ICE cotton	50,000 lb	271,573	277	79	77	31	12
CME hogs	40,000 lb	257,189	273	53	86	36	26
ICE cocoa	10 mtr. t	203,182	193	100	73	17	10
ICE coffee	37,500 lb	198,109	383	79	86	26	10
KCBOT wheat	5,000 bu	173,443	193	75	68	27	31
CME feeder cattle	50,000 lb	43,885	177	37	82	22	59

Note: MOI refers to average market open interest, whereas *C*, *NC*, *I*, and *NR* refer to the average open interest shares (in percent) of commercial, noncommercial, index, and nonreporting traders, respectively. All series range from January 2006 to December 2017, except that of soybean meal futures, which starts on April 2, 2013. As these shares include both long and short positions, they add up to 200. Slight deviations result from rounding. Abbreviations: CBOT, Chicago Board of Trade; CFTC, Commodity Futures Trading Commission; CME, Chicago Mercantile Exchange; ICE, Intercontinental Exchange Futures U.S.; KCBOT, Kansas City Board of Trade; SCOT, Supplemental Commitments of Traders.

only a single week, whereas the longest of them last half a year. Figure 1 displays the time series of the speculation measure S_t and the resulting high-speculation regimes, which are indicated by shaded backgrounds.

We use daily data from Thomson Reuters Datastream on prices, trading volume, and open interest in the market for CME feeder cattle futures as well as the macroeconomic controls, which comprise the West Texas Intermediate crude oil price, the 3-month US *T*-bill middle rate, the US broad exchange rate index and the Morgan Stanley Capital International-USA (MSCI-USA). On the basis of the availability of data for all variables, the sample consists of a total of 2,523 observations ranging from November 16, 2007 to December 28, 2017. The price data consist of a continuous futures price series obtained from switching the contract on the first day of a new trading month. Given the data on open interest and trading volume *VOL*_t, we calculate the speculation ratio R_t as the quotient of trading volume to open interest:

$$R_t = \frac{VOL_t}{MOI_t}.$$
(4)



FIGURE 1 Open interest share of speculators and high-speculation regimes

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The speculation ratio has been proposed by Garcia et al. (1986) as a measure for speculative activity. The idea behind measuring speculative activity using market open interest and trading volume is that hedgers generally hold their positions over a longer period of time, while speculators prefer intraday trade and avoid holding positions overnight. Consequently, the trading activities of hedgers will increase end-of-day market open interest, while the activities of speculators will increase trade volume (Bessembinder & Seguin, 1993; Leuthold, 1983; Rutledge, 1977). The ratio of trading volume to market open interest can therefore be interpreted as a measure for the relative dominance of speculators to hedgers. Wiley and Daigler (1998) and Ederington and Lee (2002) provide empirical evidence for the hypothesis that hedgers in futures markets prefer to hold their positions relatively longer than speculators. Whereas Wiley and Daigler (1998) do so for financial futures markets, Ederington and Lee (2002) show this for commodity futures markets. Table 2 provides descriptive statistics for the variables included in the analysis along with the results of augmented Dickey-Fuller (ADF) (1979) tests.⁸ Moreover, we compute for each of the two speculation regimes the correlation coefficients for the relationship between the returns and their conditional variance (based on a standard GARCH(1,1) estimation) and the relationship between the returns and the speculation ratio. During periods of low speculation the correlation coefficients are equal to -0.03 and -0.04, respectively. During periods of high speculation they equal -0.04 and -0.05, respectively. Obviously, the correlations vary only slightly between the different regimes. In absolute value, however, they are greater during the high-speculation regime.

4 | RESULTS

Columns (1) and (2) of Table 3 report the results for two variants of the main regression. The first specification excludes the regime dummy D_{t-1} and the interaction term I_{t-1} . The second includes all of the regressors contained in Equations (2a) and (2b). Concerning the macroeconomic controls, the results show that crude oil and the MSCI returns have significantly positive effects on feeder cattle futures returns. Both estimated parameters are significant at the 1% level. Conversely, the coefficient estimates for the *T*-bill and the exchange rate are close to zero in the mean equation and neither is statistically significant.

Regarding the effects on volatility, we find that neither crude oil nor the MSCI returns have a statistically significant effect. Instead, the *T*-bill and the exchange rate both have a significantly negative impact on volatility. Concerning the ARCH and GARCH terms, the regressions yield significantly positive estimates for both coefficients γ_1 and γ_2 . Whereas the ARCH term is in both specifications small and close to zero, the GARCH term is rather high and close to unity, suggesting that shocks to volatility die out relatively slowly. The variance stationarity constraint is met in both specifications.

Turning to the regressors pertaining to speculation, we observe that none of them is statistically significant on conventional levels in the mean equation. Nonetheless, it is worth noting that the negative coefficient estimate for the speculation ratio becomes less negative with the introduction of the interaction term. The coefficient estimate for the speculation ratio in the variance equation is strongly positive and highly significant. This suggests that increases in the speculation ratio are the main driver of daily volatility in the market for feeder cattle futures. This finding persists when we also include the regime dummy and the interaction term in the regression, of which neither is statistically significant.

Our results imply that speculation does contribute to greater uncertainty in the market regarding short-term return developments in the form of volatility clusters. This is in line with earlier findings of Streeter and Tomek (1992) and Manera et al. (2013b), who also find that the speculation ratio is a driver of volatility. However, as implied by the insignificant estimates for the speculation parameters in the mean equation, these effects only pertain to the short term, beyond which speculators do not appear to alter return dynamics. This key result is closely related to the findings of earlier studies investigating the Masters hypothesis, for example, Sanders et al. (2010), Brunetti et al. (2016), or Etienne et al. (2015). As with commodity index traders, long–short speculators do not appear to fuel price bubbles and permanently drive prices away from their fundamental value, lending support to the concept developed by Friedman (1953) and Fama (1965) that speculation stabilizes markets.

⁸The test results reveal that the logarithms of prices and the macroeconomic control variables are generally not stationary. To induce stationarity, this article considers the logarithmic returns (in percent) of these variables. Futures returns are, for example, defined as $r_t = (\ln(F_t) - \ln(F_{t-1}))\cdot 100$, where F_t denotes the level futures price. The results of a series of ARCH-LM tests suggest the use of GARCH models as the logarithmic return series features highly significant ARCH effects.

TABLE 2 Descriptive statistics and ADF test statistics

	Descriptive statistics						ADF test	
	Mean	SD	Skew	Kurt.	Min	Max	Log	Log FD
Volume	7,610.67	3,915.20	1.22	5.01	29.00	28,848.00	-3.16	-74.91
Open interest	38,677.27	9,596.93	0.21	2.82	17,981.00	65,612.00	-2.50	-40.48
Price	143.29	36.69	0.73	2.94	86.55	242.93	-1.35	-45.73
3-Month T-bill	0.35	0.59	2.64	10.37	0.01	3.32	-2.65	-63.89
Crude oil	76.78	24.66	-0.03	2.00	26.19	145.31	-1.80	-51.94
Exchange rate	107.08	9.65	0.71	2.05	93.95	129.08	-0.79	-49.07
MSCI	1,449.88	299.43	-0.12	2.26	688.64	2,106.89	-0.33	-51.72
Speculation ratio	0.19	0.08	0.94	4.51	0.00	0.70	-10.52	-74.92

Note: The critical values of the ADF test are -2.570, -2.860, and -3.430 for the 10%, 5%, and 1% level of statistical significance, respectively. Lag lengths are determined based on the Bayesian Schwarz information criterion. A series of ARCH-LM tests reveal the presence of highly significant (p < 0.01) ARCH effects in the logarithmic return series, suggesting the use of GARCH models.

Abbreviations: ADF, augmented Dickey-Fuller; ARCH, autoregressive conditional heteroscedasticity; FD, First differences; GARCH, generalized autoregressive conditional heteroscedasticity; MSCI, Morgan Stanley Capital International.

It should nonetheless be noted that the speculation ratio is not a direct measure of speculation but relates trading volume to open interest as an indication of the presence of speculators in the market. Hence, increases in the speculation ratio need not necessarily reflect greater speculative activity but could also resemble increases in trading volume due to increased information flows in the market.⁹

5 | ROBUSTNESS ANALYSIS

In this section, we examine the robustness of the previous results in four dimensions. In particular, this section discusses several alternatives for constructing the regime dummy for periods of high or low-speculative activity by long–short speculators. First, we test the robustness of our results with respect to the minimum length of any high or low-level speculation regime. Second, as mentioned in the methodology section, determining the speculative activity by long–short speculators or any other trader type hinges on the question of how to deal with the positions of nonreporting traders. Therefore, this section also experiments with different assumptions regarding such positions. Third, we explore how our results change depending on the use of alternative measures of speculative activity by long–short speculators. Lastly, we test the robustness of our results with respect to interaction terms between the speculation ratio and changes in the fundamentals.

As explained above, the way in which the regime dummy was constructed implied a number of very short regimes, some lasting only for a single week. Now, we change the construction of the dummy by imposing a minimum length restriction for both types of regimes. In particular, we require that any regime, regardless of high or low-speculative activity, lasts for at least 5 weeks. Hence, regimes shorter than 1 month are ruled out. Figure 2 demonstrates the results of this restriction graphically. If regimes must at least last for 5 weeks, we obtain a total of 13 high-speculation regimes in contrast to the initial 19. Instead of 11 weeks, these regimes now last 16 weeks on average.

Column (3) of Table 3 reports the results in comparison to the baseline results in column (2). The results in the mean and the variance equation are highly similar, in particular concerning those parameter estimates that are statistically significant. The most notable deviation from the baseline model occurs for the coefficient of the speculation ratio in the variance equation. This estimate is, unsurprisingly, roughly 5% lower in the restricted case when only longer-term regimes are considered. This is because more of the speculation ratio's impact on volatility is now absorbed in the interaction term, which remains, however, statistically insignificant.

Concerning the treatment of nonreported positions, we consider three alternative ways of how to allocate these positions across hedging and speculative positions. Whereas we initially assumed that $\alpha = 0.2$ of nonreported positions

TABLE 3 Results of GARCH regressions

	(1)	(2)	(3)	(4)	(5)	(6)
Mean equation						
r_{t-1}	0.108**	0.109**	0.109**	0.109**	0.110**	0.109**
	(0.020)	(0.020)	(0.020)	(0.020)	(0.020)	(0.020)
Oil_t	0.058**	0.058**	0.059**	0.058**	0.059**	0.058**
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
Tbill _t	0.001	0.001	0.001	0.001	0.001	0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$Exrate_t$	0.010	0.011	0.009	0.012	0.014	0.012
	(0.058)	(0.059)	(0.059)	(0.059)	(0.061)	(0.061)
MSCIt	0.035**	0.035**	0.035**	0.035**	0.035**	0.035**
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
R_{t-1}	-0.332	-0.262	-0.280	-0.378	-0.074	-0.185
	(0.262)	(0.314)	(0.307)	(0.296)	(0.324)	(0.331)
D_{t-1}		0.024	0.005	-0.050	0.100	0.058
		(0.105)	(0.107)	(0.117)	(0.102)	(0.102)
I_{t-1}		-0.329	-0.328	0.074	-0.604	-0.331
		(0.567)	(0.573)	(0.623)	(0.548)	(0.545)
Constant	0.092*	0.092	0.101*	0.110*	0.045	0.063
	(0.050)	(0.063)	(0.060)	(0.058)	(0.066)	(0.068)
Variance equation		~ /	× ,	× ,		. ,
Oil	0.034	0.024	0.037	0.032	0.037	0.047
Ou_t	(0.110)	(0.104)	(0.105)	(0.032)	(0.110)	(0.109)
Th:11	(0.110)	(0.104)	(0.105)	(0.107)	(0.110)	(0.108)
$1Duu_t$	-0.009	-0.008	-0.008	-0.008	-0.010	-0.009
E-marte.	(0.003)	(0.003)	(0.003)	(0.004)	(0.004)	(0.004)
$Exrate_t$	-1.205*	-1.216***	-1.141*	-1.165*	-1.203*	-1.150°
MECH	(0.488)	(0.445)	(0.452)	(0.459)	(0.504)	(0.508)
MSCIt	-0.080	-0.080	-0.080*	-0.078	-0.082	-0.080
D	(0.049)	(0.050)	(0.047)	(0.050)	(0.053)	(0.052)
R_{t-1}	7.175**	7.933**	7.382**	7.20/**	9.462**	8.256**
D	(0.818)	(1.158)	(1.060)	(0.920)	(1.448)	(1.372)
D_{t-1}		0.394	0.112	0.099	0.818*	0.385
Ŧ		(0.466)	(0.461)	(0.533)	(0.459)	(0.424)
I_{t-1}		-0.513	1.103	0.430	-2.523	-1.674
_		(2.151)	(2.047)	(2.490)	(1.912)	(1.826)
Constant	-5.347**	-5.594**	-5.395**	-5.297**	-6.117**	-5.631**
	(0.257)	(0.340)	(0.330)	(0.291)	(0.491)	(0.448)
GARCH terms						
ε_{t-1}^2	0.028**	0.027**	0.027**	0.027**	0.027**	0.027**
• •	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
σ^2 .	0.947**	0.947**	0.945**	0.945**	0.950**	0.948**
\mathbf{v}_{t-1}	(0.008)	(0.008)	(0,009)	(0,000)	(0.007)	(0.000)
	(0.008)	(0.008)	(0.009)	(0.009)	(0.007)	(0.008)

Note: Standard errors are in parentheses.

Abbreviations: GARCH, generalized autoregressive conditional heteroscedasticity; MSCI, Morgan Stanley Capital International.

p < 0.05.*p < 0.01.

were speculative, the regression is now repeated for $\alpha = 0$, $\alpha = 0.6$ and $\alpha = 0.8$, that is, we assume that nonreported positions are either all hedging or to be mildly or strongly dominated by speculators. The latter values reflect the idea that a larger number of speculators might choose not to classify themselves as speculators (Irwin & Sanders, 2012).

Columns (4), (5) and (6) of Table 3 report the results concerning these alternative proportions of hedgers and speculators. The estimates of the mean equation are again highly alike for all assumptions underlying the dummy construction. The results for the variance equation are also similar to before.

Next, we consider alternatives in constructing the regime dummy with respect to the underlying measure of speculative activity. As explained in the methodology section, the baseline regression built upon a dummy that had been constructed from the total open interest share that speculators take in the market. The literature has, however,

^{*}p < 0.1.

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FIGURE 2 High-speculation regimes with minimum length restriction

employed a number of alternative measures for assessing speculative activity using the trader position data from the CFTC. The most common of these is Working's *T* index of speculative activity. Let S_t^{WT} denote Working's *T*, then using the same notation as above, the index is defined as

$$S_t^{WT} = \begin{cases} 1 + \frac{SS_t}{HS_t + HL_t} & \text{if } HS_t \ge HL_t, \\ 1 + \frac{SL_t}{HS_t + HL_t} & \text{if } HS_t < HL_t, \end{cases}$$
(5)

where SL_t , SS_t , HL_t , and HS_t represent the long and short positions of speculators and hedgers, respectively. These are given by

$$SL_t = NCL_t + \alpha \cdot NRL_t, \tag{6}$$

$$SS_t = NCS_t + \alpha \cdot NRS_t, \tag{7}$$

$$HL_t = CL_t + (1 - \alpha) \cdot NRL_t, \tag{8}$$

and

$$HS_t = CS_t + (1 - \alpha) \cdot NRS_t.$$
⁽⁹⁾

The logic underlying this index is that any trade by a hedger requires a trade in the opposite direction by a speculator. If hedgers would on average like to take a short position, that is, $HS_t \ge HL_t$, then there must be a sufficient number of speculators taking a long position to clear the market. If, however, a large number of speculators take a short position, SS_t , this might be seen as an indication of "excessive speculation". Depending on the size of this speculative pressure, Working's *T* moves further and further away from its lower bound of unity.

Other measures of speculative activity, which have, for example, been used in Manera et al. (2013b, 2016), are the long-only, the short-only, and the net-long share of speculators. We denote these three shares by S_t^l , S_t^s , S_t^{nl} and define them as

$$S_t^l = \frac{SL_t}{MOI_t},\tag{10}$$

$$S_t^s = \frac{SS_t}{MOI_t},\tag{11}$$

and

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$$S_t^{nl} = \frac{SL_t - SS_t}{MOI_t}.$$
(12)

For each of these alternative speculation measures, we construct the regime dummy series of high-speculation regimes analogously to the method described above and repeat the main regression accordingly. Table 4 displays the results of these regressions.

Comparing the results for these alternative speculation measures with those of the baseline regression, we again find that the results are highly consistent across the different measures in both their economic and statistical interpretation.

ΤA	BLE 4	Regression	results for	different	speculation	measures
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	Long	Short	Net long	Working's T
Mean equation				
r_{t-1}	0.108**	0.108**	0.108**	0.109**
	(0.020)	(0.020)	(0.020)	(0.020)
Oil_t	0.058**	0.058**	0.058**	0.059**
	(0.009)	(0.009)	(0.009)	(0.009)
$Tbill_t$	0.001	0.001	0.001	0.001
	(0.001)	(0.001)	(0.001)	(0.001)
$Exrate_t$	0.014	0.011	0.011	0.013
	(0.058)	(0.060)	(0.059)	(0.060)
MSCI _t	0.034**	0.035**	0.034**	0.035**
	(0.008)	(0.008)	(0.008)	(0.008)
R_{t-1}	-0.091	-0.380	-0.133	-0.407
	(0.372)	(0.292)	(0.401)	(0.282)
D_{t-1}	0.071	-0.046	0.064	-0.113
	(0.104)	(0.132)	(0.103)	(0.153)
I_{t-1}	-0.488	0.206	-0.305	0.491
	(0.531)	(0.664)	(0.526)	(0.789)
Constant	0.056	0.102*	0.050	0.108*
	(0.079)	(0.055)	(0.082)	(0.054)
Variance equation				
Oilt	0.050	0.021	0.053	0.015
	(0.107)	(0.113)	(0.111)	(0.103)
Tbill,	-0.010**	-0.009**	-0.009**	-0.008*
	(0.003)	(0.003)	(0.003)	(0.003)
$Exrate_t$	-1.308**	-1.233**	-1.229*	-1.166**
	(0.502)	(0.475)	(0.503)	(0.416)
MSCIt	-0.086*	-0.078	-0.085*	-0.075
	(0.050)	(0.052)	(0.051)	(0.051)
R_{t-1}	9.100**	7.471**	9.581**	7.190**
	(1.791)	(0.996)	(1.902)	(0.887)
D_{t-1}	0.633	0.340	0.829	0.541
	(0.527)	(0.687)	(0.546)	(0.620)
I_{t-1}	-1.494	-1.437	-2.841	-0.591
	(2.225)	(3.012)	(2.265)	(2.839)
Constant	-6.162**	-5.412**	-6.151**	-5.189**
	(0.580)	(0.292)	(0.599)	(0.284)
GARCH terms				
ε^2	0.024**	0.028**	0.025**	0.028**
\mathbf{v}_{t-1}	(0.005)	(0.005)	(0.005)	(0.005)
-2	0.0037	0.003	0.003	0.003
o_{t-1}	(0.007)	(0,000)	(0.000)	(0,000)
	(0.007)	(0.008)	(0.008)	(0.009)

Note: Standard errors are in parentheses.

Abbreviations: GARCH, generalized autoregressive conditional heteroscedasticity; MSCI, Morgan Stanley Capital International.

 $^{\ast}p<0.1.$

***p* < 0.05.

***p < 0.01.

A criticism regarding the speculation ratio as proposed by Garcia et al. (1986) is its behavior during periods of rapidly changing fundamentals. If it is primarily hedgers rather than speculators that respond to such changes, the speculation ratio, which relates trading volume to open interest, might during these time periods be a poor proxy for the level of speculation. In fact, in these periods the ratio might even be a proxy for hedging activity. To test the robustness of our results with respect to this concern, we repeat the baseline regression but include interaction terms between the speculation ratio and the absolute values of the first differences of the macroeconomic control variables in both the mean and the variance equation. Table 5 displays the results of this regression.

None of the interaction terms has a statistically significant impact in the mean equation. However, changes in the oil price appear to significantly increase the impact of the speculation ratio on volatility, whereas changes in the *T*-bill rate and the MSCI appear to reduce it. The coefficient for R_{t-1} remains, however, positive and highly significant.

To sum up, this article has explored how the speculation ratio affects the return dynamics in a market environment that is particularly prone to suffer from detrimental effects due to speculation, that is, a thinly traded commodity market with a nonstorable underlying commodity during periods of increased speculative activity. To check the robustness of our results, we have explored a number of different ways to construct the regime dummy for periods of high speculation. These comprised different regime lengths, different assumptions regarding nonreported positions, and a variety of alternative speculation measures. Our key finding from the initial regressions, namely that the speculation ratio is an important driver of volatility but does not affect the level of returns and is independent of the speculation regime, has proven robust across all of these different specifications.

	Coefficient and standard error	
Mean equation		
rt_1	0.112**	(0.021)
Oilt	0.056**	(0.009)
Thill.	0.001	(0.001)
Exrate	-0.005	(0.061)
MSCI _t	0.037**	(0.007)
$ \Delta Oil_{t-1} \cdot R_{t-1}$	-0.012	(0.044)
$ \Delta T bill_{t-1} \cdot R_{t-1}$	-0.000	(0.002)
$\Delta Exrate_{t-1} \cdot R_{t-1}$	-0.009	(0.325)
$\Delta MSCI_{t-1} \cdot R_{t-1}$	-0.017	(0.040)
R_{t-1}	-0.200	(0.342)
D_{t-1}	0.051	(0.105)
I_{t-1}	-0.430	(0.579)
Constant	0.085	(0.064)
Variance equation		
Oilt	0.017	(0.055)
Tbill _t	0.032**	(0.002)
Exrate	0.235	(0.452)
MSCI _t	-0.123**	(0.033)
$ \Delta Oil_{t-1} \cdot R_{t-1}$	0.534**	(0.158)
$\Delta T bill_{t-1} \cdot R_{t-1}$	-0.217**	(0.031)
$\Delta Exrate_{t-1} \cdot R_{t-1}$	1.146	(1.238)
$\Delta MSCI_{t-1} \cdot R_{t-1}$	-0.995**	(0.375)
R_{t-1}	5.708**	(1.311)
D_{t-1}	-0.713*	(0.360)
I_{t-1}	5.156**	(1.775)
Constant	-4.290**	(0.297)
GARCH terms		
ε_{t-1}^2	0.031**	(0.006)
σ_t^2	0.929**	(0.008)
		-

TABLE 5 Regression results when including interaction terms with controls

Note: Standard errors are in parentheses.

Abbreviations: GARCH, generalized autoregressive conditional heteroscedasticity; MSCI, Morgan Stanley Capital International.

*p < 0.1.

***p* < 0.05.

***p < 0.01.

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6 | CONCLUSION

This article has examined how long-short speculation affects the return dynamics in a market environment where one would expect these effects to be particularly strong, that is, when trading volume is low, the number of speculators is high and inventories are expensive to hold. The analysis focused on the market for CME feeder cattle futures, which is one of the most illiquid markets included in the CFTC SCOT reports. It features by far the highest number of nonreporting positions and is additionally characterized by a nonstorable underlying. Within this market setting, we use a Markov-switching approach to single out periods characterized by high-speculative activity of long-short speculators.

The results of the mean equation in the GARCH analysis reveal that speculation as measured by the speculation ratio does not significantly impact returns, whereas the variance equation suggests that speculation does influence volatility. The parameters of the dummy variable for periods characterized by high-speculative activity are generally not significant, neither is the interaction term between the regime dummy and the speculation ratio. This implies that speculation does not arbitrarily drive returns away from levels implied by fundamentals, not even during high-speculation regimes. In particular, the results do not support the idea that speculation inflates prices permanently in the form of price bubbles. Only with respect to the short term do the results support the notion that speculation impacts returns by increasing their volatility. But, as we measured the activity of long–short speculators using the speculation ratio, these results can also be interpreted as a sign for increased volatility due to increased information flows that manifest themselves in higher trading volume.

Whereas previous research has largely rebutted the Masters hypothesis and has found no evidence that commodity index traders are to blame for price hikes in commodity markets, this study reaches a similar conclusion for the case of classical long–short speculators. Most notably, this study does so for a market environment where one would expect the strongest effects of speculation. Consequently, it seems unlikely that long–short speculation would destabilize prices in more liquid markets with lower levels of speculation. Even though the evidence strongly points toward fundamentals, the question of what caused the price hikes observed in numerous commodity markets during 2007–2008 remains unanswered and open for future research. Yet, the policy implications of this study are rather clear. As futures speculation is not found to drive prices away from fundamental values, regulators should be wary of reining in speculation in futures markets and thereby depriving commercial traders of their ability to hedge against undesirable price movements.

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