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Regime Shifts in the Fish Meal/Soybean Meal Price Ratio

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Abstract

Aquaculture growth has led to worries about overfishing and reduction in wild-caught food fish supply because of increased demand for fish meal. As such, the price ratio between fish meal and soybean meal has received much attention as an indicator of changing market conditions. In recent years, the price ratio between these two commodities has become more volatile. Several authors have suggested that the traditional relationship between fish meal and soybean meal has broken down and that this is evidence of increased demand pressure on fish meal. In this article, we investigate the hypothesis that there are two regimes for the relative price between fish meal and soybean meal. The empirical results support this hypothesis, with the low-price regime representing the traditional stable relative price. The continued linkages between the fish meal and the soybean meal markets indicate that aquaculture is reducing its dependency on marine proteins in favour of vegetable proteins.

Keywords: *Regime shifts; commodity markets; fish meal demand.*

JEL classifications: *Q11, Q13, Q22.*

1. Introduction

It is well documented that many food markets experience cyclical behaviour and structural shifts, features that have been documented in a number of studies (e.g. Holt and Craig, 2006; Wang and Tomek, 2007). Less attention has been given to the possibility that these markets can experience irregular regime switching caused by exogenous shocks. This is behaviour that is recognised as important for many financial and macroeconomic settings (e.g. Hamilton, 1989), and it can be important in commodity markets as well. In this article, we investigate one market where price movements are likely to be characterised by regime switching: the fish meal and soybean meal markets. We use the regime shifting model of Hamilton (1989, 1994).

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The price relationship between fish meal and soybean meal has received a great deal of attention in the literature, not only because of historical strong linkages between these two markets. During the last couple of decades price formation in the fish meal market appears to have changed radically as aquaculture has become the largest buyer using marine proteins as a prime ingredient in fish feed (Vukina and Anderson, 1993; Asche and Tveteras, 2004, 2008; Kristofersson and Anderson, 2006; Tveteras, 2010; Tveteras and Tveteras, 2010; Tveteras *et al.*, 2012).[2] Increased demand pressure on fishmeal has also led to concerns about the sustainability of fisheries targeted for fishmeal production and the use of these wild fish resources as feed instead of food (Naylor *et al.*, 2000).

Traditionally, fish meal was one of several protein sources for terrestrial animal feed and, as such, part of the much larger market for vegetable meals. Until the 1990s, the long-run relative price between fishmeal and the dominant protein source for feeds, soybean meal, was constant, although with short-run variation due to specific market shocks such as *El Niños* (Vukina and Anderson, 1993; Asche and Tveteras, 2004, 2008). From the late 1980s there has been a gradual shift in the users of fish meal from chicken and pork feed producers to aquaculture feed producers. Aquaculture has since increased its share of fish meal from virtually nothing in 1980 to 59% in 2008 (Jackson and Shepherd, 2010). The change in users suggests that despite the constant relative price, fish meal has some unique characteristics that differentiate it from vegetable proteins, represented by soybean meal.

Potentially, the unique characteristics of marine proteins can segment the fish meal market from the soybean meal market. Kristofersson and Anderson (2006) and Tveteras (2010) show that there was a regime shift in 1998 in the relative price between fish meal and soybean meal following a strong *El Niño*. These studies suggest that the underlying reason for the regime shift is an increased demand for fish meal because of its unique properties, leading prices to become more sensitive to supply side shocks such as *El Niño*. Regime shift, in this context, means that the market moves from a relatively stable price ratio between fish and soybean meals to a situation where this ratio becomes more volatile. As such, this could be evidence that the fish meal and soybean meal markets are segmenting. Several authors indicate that this is caused by rapidly growing demand from the aquaculture sector (Naylor *et al.*, 2000).[3] Moreover, Naylor *et al.* (2000) propose the 'fish meal trap' hypothesis, which claims that aquaculture production is limited by the availability of wild fish to be used for feed.

However, if the fish meal price is increasing compared to other vegetable meals, this provides strong incentives for innovation among feed producers to reduce fish meal use in the feed to avoid increased costs (Kristofersson and Anderson, 2006; Asche, 2008). Chicken and pig feed producers have already gone through such a process and, as a result, fish meal is hardly used in chicken feeds today. In pig feeds, it is now primarily used as a strategic input because of its beneficial effect on growth of weaner pigs. Reduction in the inclusion of fish meal in feed is also taking place in aquaculture. The share of fish meal in the feed of aquaculture species that

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[2] Fishmeal is also a fairly homogenous product, in contrast to fish in retail (Roheim *et al.*, 2011).

[3] Aquaculture has been the world's fastest growing food production technology in recent decades (Smith *et al.*, 2010a).

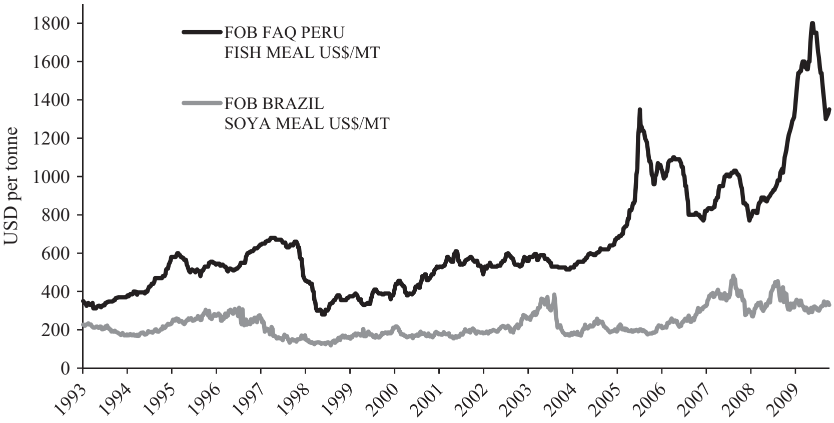

Figure 1. Fish meal price (FOB FAQ PERU) and Soybean meal price (FOB Brazil)

typically consume large amounts of marine proteins has declined substantially (Tacon and Metian, 2008). For example, in the early 1990s salmon feeds normally contained over 50% fish meal, while current feed can contain as little as 15%.

The prices of fish meal and soybean meal are shown in Figure 1 and their price ratio along with a linear trend in Figure 2. The linear trend regression of the price ratio shows evidence of an increasing trend, suggesting that the markets are segmenting, reflected by fishmeal becoming relatively more expensive than soybean meal.[4] However, a more thorough inspection of Figure 2 seems to indicate that in some periods, the relative price reverts back to the level typically observed in the 1990s, intersected by periods of significant deviations, associated with *El Niños*. Hence, it seems that the relative price has two different regimes; a constant relative price regime to which the market returns after different shocks and a high-price regime associated with periods of short supply.

In this article, we test the hypothesis that the relative price between fish meal and soybean meal has different regimes as well as whether the relative price is constant or changing in the different regimes. In our model specification, we allow the constant, trend and relative price components to change between two regimes. This allows for endogenous structural changes in the dynamics of the relative price, making it possible to test for the effect of market shocks without making *a priori* assumptions on the number and locations of structural changes. We apply the regime shifting model of Hamilton (1989, 1994), where the likelihood of being in a regime at a given time follows a Markov process. Since its introduction, the regime switching model has become popular in accounting for changes in the dynamics of economic variables. Specific applications can be found in business cycle modelling (Hamilton, 1989; Bansal *et al.*, 2004), and interest rate movements (Garcia and Perron, 1996; Gray, 1999; Ang and Bekaert, 2002). Tomek (1997) provides an application of the effect of macroeconomic regime shifts on commodity prices. Our application is novel in the sense that it is supply shocks in the agricultural commodity markets themselves that can cause the regime shifts.

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[4] Using the weekly data in Figure 2, this regression gives the result $P = 1.953 + 0.002^{*}t$, with $R^2 = 0.25$, where P is the relative price and both parameters has P-values < 0.001.

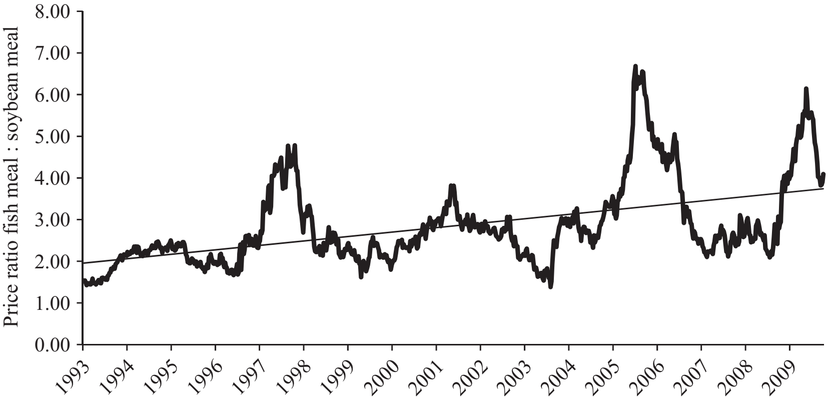

Figure 2. Fish meal: soybean meal price ratio and linear trend.

2. Background and Data

Most of the world's fish meal production is based on fisheries of small pelagic species.[5] Pelagic fish are used both as food and for reduction into fish meal and fish oil for use in derived products – mostly feed. Certain species are mainly fit for reduction purposes due to low profitability as food products (i.e. caused by a combination of relatively high marketing cost and low attractiveness to consumers). Wijkström (2009) puts species like sand eels, Gulf menhaden, and Norway pout in this category. Other species are mainly used for fish meal, but where a small share is also used for human consumption. Examples of this category are Peruvian anchovy, capelin, blue whiting and European sprat. Finally, there are the prime food grade pelagic fish including species like mackerel, herring and sardines.[6] These are species for which there are well developed food markets, but in times with excess supply part of the landings are used for fish meal.

A characteristic of pelagic fisheries is that while the quantity for human consumption is relatively stable, the 'surplus' destined for fish meal production can vary dramatically (Asche and Tveteras, 2004). Thus, in years when catches are low, such as in *El Niño* periods, the fish meal industry faces supply shortages. The pelagic fisheries have also generally been described as fully exploited or over-exploited by the FAO (Grainger and Garcia, 1996; FAO, 2010). A substantial expansion in the global fish meal production above an annual of 6–7 million tonnes is therefore unlikely unless prices for fish meal increase substantially. In recent years, fish meal production has been relatively stable largely because of increased use of cutoffs, as landings of wild fish have been diverted from reduction to human consumption. Production shares for fish meal in 2008 are shown in Figure 3. Peru and Chile jointly supplied 33% of the global fish meal production. Other important producers

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[5] Pelagic fish are migrating fish species that inhabit the surface waters, as opposed to demersal fish that inhabit deeper (bottom) waters.

[6] It is worthwhile to note that for some species the product mix is influenced by the management system and its interaction with the fish stocks (Homans and Wilen, 2005; Smith, 2008).

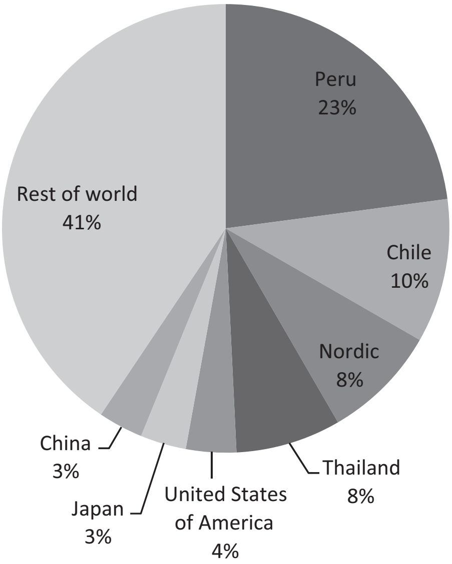

Figure 3. Global production share of fish meal in 2008.
Source: FAO Fishstat, Jackson and Shepherd, (2010).

are Thailand with 8% and the Nordic countries Denmark, Iceland and Norway, collectively with 8% of the global fish meal production.[7]

Feed producers' willingness to pay relatively higher prices for marine proteins compared with vegetable proteins such as soybean meal is determined by the additional value of fish meal when combined with other input feed ingredients. The added value arises as a result of several factors including better growth performance, reduced mortality, better taste and increased consumer acceptance. In Figure 4, one can see that in 2008, aquaculture was the main user of fish meal with a share of 59%, indicating that the aquaculture sector on average has a higher willingness to pay for fish meal. There are several possible explanations for this, where the most important seems to be that marine proteins mirror the nutritional requirements of several aquaculture species. Moreover, lack of knowledge with respect to nutritional requirements limits more cost efficient use of different protein sources. Pork and poultry makes up most of the remaining use, jointly consuming 36%.

For most aquaculture species, fish meal only accounts for a small part of their diet. Other protein meals make up the major share, with soya as the largest component. This is also reflected in the aggregate picture as shown in Figure 5, since total fish meal production has not increased despite a strong growth in total aquaculture production. Fish meal production is minor compared with the total protein meal production, and is about 3% of soybean meal world disappearance. If fish meal continues to be part of the larger protein meal market, increased aquaculture production will not primarily lead to increased demand for fish meal, but rather increased demand for protein meals, of which vegetable meals account for the

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[7] The quality of the fishmeal produced in Thailand is normally inferior to the other major producers mentioned above and is mostly consumed domestically.

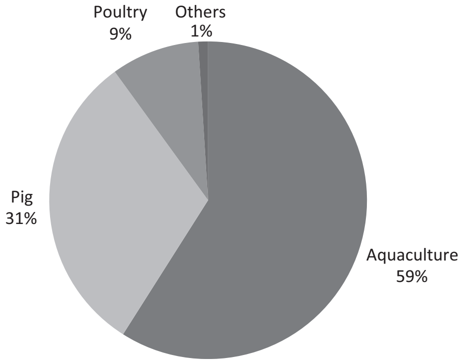

Figure 4. Fish meal use by sector, 2008.
Source: Jackson and Shepherd (2010).

majority. Furthermore, if fish meal is not an essential ingredient in aquafeed, increasing prices of fish meal relative to other protein sources will lead aquaculture to reduce its demand for marine proteins, by developing feed formulations based on vegetable proteins to satisfy the nutritional requirements of the fish. This is the direction of travel for the aquaculture industry, with the steady reduction in the use of fish meal in aquafeed (Tacon and Metian, 2008; Figure 5).

To analyse how the rapid changes in the marketplace for fish meal have affected the price relationship between fish meal and soybean meal, we use weekly FOB prices from the International Fishmeal and Fish Oil Organisation (IFFO) (Jackson and Shepherd, 2010) of fish meal from Peru and soybean meal from Brazil. The price data span January 1993 to September 2010. During this period there have been *El Niño* events in 1993, 1994, 1997–98, 2002–03, 2006–07 and 2009–10. These events are associated with an increase in sea surface water temperatures in the Southeast Pacific. When these events are strong, the warm sea surface water reduces upwelling of cold nutritious water of the Humboldt current, so that productivity of

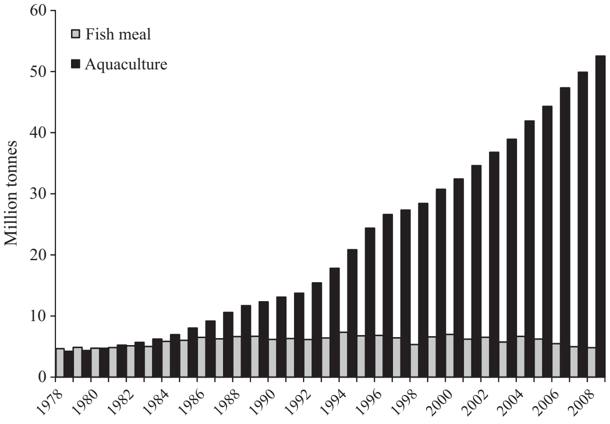

Figure 5. Fish meal and aquaculture production.
Source: FAO, Fishstat.

the world's largest fishery and most important source of fish meal – that of Peruvian anchovies – declines. The *El Niño* in 1997–98 was particularly strong and the volume of Peruvian anchovy catches in 1998 was less than one-fifth of the average catch level during the period 1993–2009. This greatly reduced the availability of fish meal in international markets. But less severe *El Niño* events have also affected fish meal supply. Prolonged fishing closures in Peru caused by, for example, *El Niño* tend to put many fishmeal producers in cash flow problems and force them to sell existing fishmeal stocks. Moreover, when fishing reopens there is initially limited stocking activity as companies need revenues to resolve cash flow constraints. This behaviour most likely exacerbates price volatility. It is the effects on the fish meal/soybean meal price ratio of these supply side shocks combined with the structural changes on the demand side that this paper sets out to analyse.

3. Method

Several different nonlinear alternatives to the standard linear model exist in the time series literature. In addition to the regime switching model, threshold autoregression (Tong, 1978) and smooth transition autoregression (Chan and Tong, 1986) models have become popular (Holt and Craig, 2006; Balcombe and Rapsomankis, 2008; Balagtas and Holt, 2009). These models allow parameters to change dependent on some function of underlying variables, such as powers of lagged dependent variables. To investigate the dynamics of the relative price of fish meal and soybean meal, we apply the two-state regime shifting model of Hamilton (1989, 1994). The choice of this model over alternatives is motivated by our empirical interest. The hypothesis that relative prices move between different pricing regimes, characterised by differences in mean and variances, dependent on some unobservable latent variables (such as *El Niño* effects) suggests that a regime switching model is an appropriate nonlinear alternative. Regime changes are determined endogenously by the relative fit of each regime to the data. The STAR model, in contrast, restricts nonlinear dynamics to be dependent on predetermined variable(s). In lieu of obvious predetermined variables driving regime changes, we opt for the data driven regimes implied by the regime switching model. It is worth noting that in a comparison of smooth transition and regime switching autoregressive models on US unemployment data, Deschamps (2008) finds that both models provide similar descriptions of the data.

Our model specification allows the intercept and trend of relative prices to shift between two different regimes. To accommodate volatility shifts, we also model the constant term in an ARCH (AutoRegressive Conditional Heteroskedasticity) representation of conditional variance to be regime dependent.[8] To accommodate the high degree of persistence, a non-regime dependent autoregressive component is added to relative prices. Let y denote the $[T \times 1]$ vector of relative prices (fishmeal divided by soymeal price). The unobservable state S_t is a $[2 \times 1]$ vector $[S_{1t}, S_{2t}]$ where element j of S_t takes value one if state j is realised and zero otherwise. The conditional mean and variance is then modelled as:

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[8] We also tried a specification with a state dependent autoregressive parameter. However, little evidence was found for regime shifts in this parameter.

$$y_t = \mu(S_t, t) + \beta y_{t-1} + \varepsilon_t, \quad \varepsilon_t \sim IID(0, \sigma^2_{S_t}), \tag{1}$$

$$\mu(S_t, t) = S_{1t}(\mu_1 + \gamma_1 t) + S_{2t}(\mu_2 + \gamma_2 t) \tag{2}$$

$$\sigma^2_{s_t} = \alpha_0 \theta(S_t) + \alpha_1 \varepsilon^2_{t-1}, \quad \theta(S_t) = 1, \tag{3}$$

where t denotes time. If the process is in state 1, $S_{1t} = 1$ and $S_{2t} = 0$, the intercept and time trend for the process is $\mu_1 + \gamma_1 t$. This is allowed to change to $\mu_2 + \gamma_2 t$ when the process changes to state 2, $S_{1t} = 0$ and $S_{2t} = 1$. In equation (3) $\theta(S_2) = \theta_0$ if $S_{2t} = \theta_0$ if $S_{2t} = 1$, and unity if $S_{1t} = 1$. This setup for the variance implies that variance is scaled by θ_0 if the process is in state 2. We follow Hamilton (1989) and model S_t as the result of a discrete time, discrete state first-order Markov process. Defining P as a $[2 \times 2]$ matrix of transition probabilities, the state process S_t evolves as a first-order autoregressive process:

$$S_t = PS_{t-1} + v_t, \tag{4}$$

where v_t is the error in predicting the state at time t from using information available at $t - 1$. Conditional on the set of parameters to be estimated: $\Phi = [\mu_1, \mu_2, \gamma_1, \gamma_2, \beta, \alpha_0, \theta, \alpha_1, P]$, inference on the unobservable state variables can be achieved by an iterative approach similar to the Kalman-filter.

Let $\bar{S}_{t|t}$ denote our current inference on states conditional on the parameters and data. The states can then be integrated out of the joint likelihood of observing the data and states. This provides us with the log-likelihood of observing the data conditional on parameters alone. The parameters Φ that maximise the log-likelihood function can then be found by conventional maximum likelihood methods. To start the procedure we fix the starting state, $\bar{S}_{1|0}$, at its unconditional, or ergodic, level. Since states $S_{t|t}$ are only evaluated using information available at time t we perform a sweep through the data at the end of the estimation to use the full sample to evaluate each time-dependent state (Hamilton, 1994, pp. 694). This gives us the smoothed states $\bar{S}_{t|T}$. Note that the state series can be interpreted as the probability of existing in either regime at a given time. If regimes evolve according to unobserved market drivers, we can use this measure in an intervention-style analysis to evaluate if dynamics change in a manner consistent with our hypothesis.

4. Empirical Results

The model is estimated using the maximum likelihood approach of Broyden, Fletcher, Goldfarb and Shannon (BFGS). Initial values are traced using the Simplex method. To highlight the relevance of the regime shifts, we compare the model to its linear counterpart. Although the linear alternative is non-nested, precluding Likelihood Ratio tests, we can compare the models by how well they encompass the data.

Prior to estimation we also check if the series contain significant seasonality, as seasonality might potentially affect regime changes. To investigate seasonality we augment the linear AR(1) model with annual cycle length trigonometric functions to represent seasonality. We do not find evidence of any significant seasonality in relative prices. The test statistics for exclusion of seasonality in the linear model gives a P-value of 0.80.

From Table 1, it is evident that the regime shifting model provides a significant improvement in fit over the linear model. This is perhaps not surprising given the regime switching model contains five more parameters than the linear model. A likeli-

Table 1

Estimation Results

	Linear model		Regime shift model	
	Coef.	SD	Coef.	SD
P_{11}	–	–	0.974**	0.0079
P_{12}	–	–	0.032**	0.0097
μ_1	0.043**	0.0048	0.045**	0.0042
μ_2	–	–	0.054**	0.011
γ_1	4.7e-05**	9.0e-06	8.0e-06	9.2e-06
γ_2	–	–	4.1e-05**	1.8e-05
β	0.975**	0.0013	0.979**	0.0014
α_0	0.017**	0.0005	0.007**	0.0003
α_1	0.296**	0.0292	0.045	0.0397
θ_0	–	–	6.068**	0.3829
Log-likelihood	411.10		508.71	
Ljung–Box: $\chi^2(24)$	42.46**		25.56	
Normality: $\chi^2(2)$	156.73**		35.394**	
Instability	1.4726**		0.3107	

Note: **indicates statistically significant at the 1% level

hood ratio test gives a test statistic of 195.22. Although the likelihood ratio test is not valid asymptotically, the high test statistic implies that even a very conservative critical value would reject equality of the models. The instability test of Hansen (1992) indicates the parameters of the linear model are non-constant. We cannot reject the null of stability for the regime shift model. The Ljung–Box statistic rejects serial correlation for the regime shift model, indicating that a first-order lag for the autoregressive component is sufficient. The AR(1) structure of the model is also selected using the Akaike Information Criteria. Moreover, re-estimating the model with additional lags up to 6 months does not significantly improve model fit over the AR(1) specification. Both models reject normality according to the Jarque–Bera test, although some residual skewness and kurtosis is accounted for by the regime shift model.

The regime shift model identifies two persistent regimes, where the probability of changing regime is less than 0.04 for both regimes. We further note that all coefficients are significant, except the trend component in regime 1 and the ARCH component. Since the ARCH component is significant in the linear model, the regime shift in conditional variance accounts for the apparent ARCH effect.

For relative prices, our results imply that when pricing changes from regime 1 to 2, the mean relative price increases. Regime 1 is low-price regime while regime 2 is associated with higher prices. Moreover, a positive significant trend is present in the high-price regime, while there is no trend in regime 1. Regime 2 is also accompanied by a six-fold increase in variance. Hence, in regime 1 the relative price is constant.

Figure 6 plots the relative price of fish meal and soybean meal in addition to the smoothed probability of being in regime 2. We observe that regime 2 mostly occurs in periods where the price ratio bubbles. The first prolonged period of regime 2 during the 1990s can be explained by the arrival of the *El Niño* of 1997–98. During this period, global supply of fish meal reduced substantially and led to scarcity and inflation of fish meal prices. The next *El Niño* in 2002–03 also seems to have influenced the relative price level, but to a milder extent. However, the subsequent *El Niño* events in

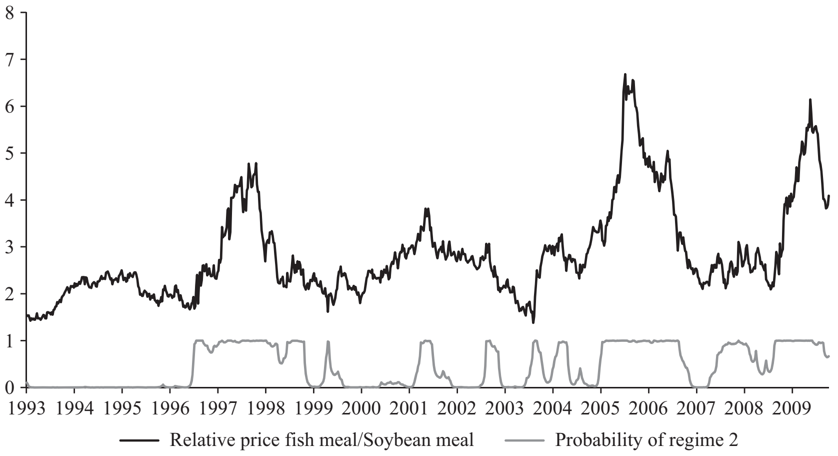

Figure 6. Relative price of fish meal and soybean meal and smoothed probability of existing in regime 2

2006–07 and 2009–10 appear to have had a significant impact on relative prices and generated prolonged periods of price regime 2. The financial crisis probably contributed to a muting of the effect of the 2009–10 *El Niño*. Moreover, it should be noted that forecasts of upcoming *El Niño* can influence price expectations and, as such, climatic events can affect price before they actually occur. This can explain why the fish meal/soybean meal price ratio appears to be affected prior to *El Niño* events. Other market factors also influence price. An important factor is the growth in commodity imports to China, which includes both fish meal and soybean meal. While it is not obvious how this has affected the historic volatility it is clear that the increased import demand from China has exacerbated temporary market shortages and therefore contributed to increased volatility in the fish meal price.

From the above discussion, it seems that when fish meal supply is sufficiently abundant relatively to soybean meal, the traditional stable relationship with a relative price of two between the prices found in Vukina and Anderson (1993) and Asche and Tveteras (2004) prevails. However, when this relationship breaks down – typically during *El Niño* events where supply contracts – the relative price changes significantly, and the trend indicates an increased relative premium for fish meal. Hence, the high-price regime seems to be associated with periods where fish meal is demanded for its unique properties. This kind of price behaviour is consistent with a kinked demand curve, where low supply corresponds to price being determined on the vertical part of the demand curve, which can lead to very high prices (i.e. Regime 2). However, when fishmeal production is high, the supply curve intersects with the horizontal part of the demand curve that determines the price floor associated with the low-price regime, Regime 1.

Since the autoregressive component is close to unity it is necessary to test for the presence of a unit root in the relative price. In terms of economics, we will not expect a relative price to exhibit a unit root, as a known substitution relationship exists between the commodities. However, since substitution is not instant it is not surprising that relative prices display significant persistence. If one combines this persistence with structural changes in mean and trend, the price is likely to display a near unit-root dynamic not easily statistically distinguishable from a true unit-root process. However, when we account for regime changes we would expect the evidence for a unit-root to be significantly reduced. When the alternative hypothesis is a stationary regime-switching model, one can question if the distribution of standard unit root tests is appropriate. Nelson *et al.* (2001) investigate the power of unit root tests under the Markov regime switching alternative. They find that power distortions are dependent on the persistence of the regime-shift effects. When effects are transitory, traditional unit root tests in general have good power. As the effects of regime shifts in our model are transitory we use conventional critical values when evaluating the test results.

To evaluate unit-roots in the series, we apply the approach suggested by Ng and Perron (2001). Ng and Perron illustrate that using GLS detrended data in collaboration with a Modified Information Criteria (MIC) yields improved size and power properties. The MIC takes into account that the bias in the sum of autoregressive coefficients depends on the truncation lag chosen. In the testing procedure, we first apply GLS detrending to both series. Next, we choose an appropriate lag level using the Modified Akaike Information Criteria (MAIC). We then apply five unit root tests, allowing both a constant and a constant and trend component, to the implied trend adjusted fish meal, soybean meal, relative prices and regime adjusted relative prices. The specific statistics for the test can be found in Ng and Perron (2001). The five unit root tests are the Augmented DickeyFuller (ADF) test, modified versions of the Phillips (1987) and Phillips and Perron (1988) tests, MZ_α and MSB, the Elliot *et al.* (1996) feasible point optimal test P_{GLS}, and the modified version of the feasible point optimal test MP_{GLS}.

Table 2 shows the results of the unit root test, where the tests are conducted with and without a trend. For fish meal, the evidence is clearly in favour of a unit root. The evidence for soybean meal price is less clear, as four of the five tests with a constant reject the null hypothesis of a unit root while all the tests with a trend cannot reject the null. As the tests with the trend are the most reliable when there are conflicting results (Banerjee *et al.*, 1993), we conclude that the series contains a unit root. The results are similar for the unit root tests for the relative price when the regime shifts are not accounted for. However, the null hypothesis of a unit root can be rejected in all cases when the structural breaks are accounted for. Asche and Tveteras (2004) found the fish meal and soybean meal prices to be cointegrated using data up to 1999 that showed no evidence of structural breaks. Our results confirm this relationship also in a period with a less stable relationship when the structural breaks are accounted for.

Table 2

Unit root tests

	Fish meal		Soybean meal		Relative price		Regime adjusted relative price	
	C	T+C	C	T+C	C	T+C	C	T+
Lag	2	2	3	3	0	0	1	1
ADF	−0.325	−0.324	−2.721*	−2.721	−2.582**	−2.582	−3.749**	−3.749**
MZ_a	−0.425	−0.425	−13.70*	−13.70	−12.71*	−12.71	−27.99**	−27.99**
MSB	0.645	0.645	0.191*	0.191	0.198*	0.19	0.133**	0.133**
P_{GLS}	36.15	8.86	4.20	8.42	2.45*	7.21	0.904**	3.30**
MP_{GLS}	24.93	85.30	1.78*	6.66	1.93*	7.17	0.897**	3.29**

Note: * and ** indicate statistically significant at a 5% and a 1% level, respectively.

5. Discussion and Conclusion

Kristofersson and Anderson (2006) and Tveteras (2010) conclude that there was a regime shift in the relative price of fish meal and soybean meal in 1998. This may be interpreted as a sign of the increased demand pressure on scarce wild fish resources leading to a decoupling of the fishmeal market from the soybean meal market. In this article, we follow Hamilton (1989) to test the hypothesis that rather than being a permanent shift in the relative price, recent history exhibits a market switching between two regimes. Our empirical results indicate that the conclusion that there has been a permanent regime shift in the market is too strong. Rather, the market appears to shift between two persistent price regimes. Regime 1 is characterised by relative low variance and a constant unconditional mean price ratio between fish and soy meals, while in regime 2 volatility increases, the mean price ratio increases (fish being more expensive) and exhibits a significant positive trend. We also show that the evidence of unit roots in relative prices is significantly reduced when we account for regime shifts. Hence, the regime shifting model of Hamilton (1989, 1994) can provide a good specification for the dynamics in some commodity markets, accounting for significant volatility but with a weaker alternative hypothesis than a structural break.

For this particular case, evidence of two regime shifts can be interpreted as reflecting reducing demand pressure from a growing aquaculture sector. In the fishmeal market, increasing volatility as represented by the positive trend in the high-price regime can be linked to structural changes on the demand side. All the main sectors that consume fish meal are growing, which implies that the market for fish meal is expanding. Moreover, many of these sectors tend to prefer fish meal over alternative vegetable protein sources such as soybean meal, *ceteris paribus*. Since the supply of fish meal remains relatively stable, it is clear that competition for marine proteins has become more intense (Tveteras and Tveteras, 2010).

The periods with very strong competition for fish meal caused by shortages of fishmeal stocks not only push buyers to re-examine their use of marine proteins, but high prices also lead feed producers in aquaculture, chicken and pork production to make an explicit effort to reduce technological dependency on fish meal through innovation. As aquaculture is now the largest user of fish meal, and chicken and pork producers reduced their use significantly in the 1990s, it is currently in aquafeed production that the largest changes are taking place. However, in all types of aquaculture feed the average inclusion rate of fish meal has been reduced. The fish meal inclusion per kg of fish produced in aquaculture can be computed using FAO and IFFO data (FAO, 2010; Jackson and Shepherd, 2010). This is shown as indices for different species in Figure 7. The average usage of fish meal per kilo of fish produced has been reduced by 29% from 2000 to 2008. As a result, innovation has provided new feeds with more flexible ingredient formulation. In the long run, the development of new feed formulations is likely to break the positive price trend associated with regime 2, as the main factor in the increased aquaculture production has been productivity growth through innovation and lower production cost (Asche, 2008; Smith *et al.*, 2010a,b).

The explosive growth of aquaculture production has not been able to decouple the fish meal market from the soybean meal market and therefore we can reject the 'fish meal trap' hypothesis which claims that stagnant supply of fish meal limits the growth of aquaculture. There is no such evidence. However, the fact

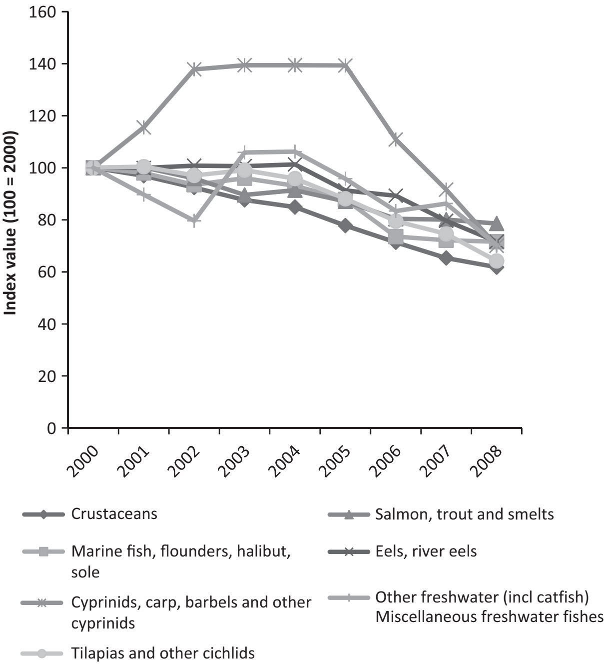

Figure 7. Indices of fish meal inclusion rates in different types of aquaculture production

that fish meal prices are becoming more volatile means that growth of aquaculture will favour those species that are less dependent on proteins from wild-caught fish. Thus, species that have traditionally been highly dependent on fish meal (shrimp and salmon) must reduce their dependency on fish meal, as is already happening.

This should be good news for those who believe that more of the world's small pelagic fish resources should be used to feed the world's poor instead of ending up as ingredients in animal and fish feeds. If demand from the aquaculture sector is abating then other uses for these pelagic fish should become more attractive in economic terms. However, there are reasons to believe that such a shift will take time. First, falling demand due to lower fish meal inclusion rates in aquaculture is offset by continued growth of the aquaculture sector. Second, while a part of these fish resources has a market as food fish, consumers' willingness to pay for most of these species is low. Due to the relatively high marketing costs involved, the growth in food-grade products of small pelagic fish is more likely to be pushed by developing countries' emerging middle class rather than their poor.[9]

Thus, small pelagic fish and consequently fish meal will continue to be sold to those with the highest willingness to pay. The fact that the chicken sector, which is highly sensitive to price changes, consumed 9.1% of the fish meal output in 2008 underscores that fish meal is still part of the vegetable protein market. This limits

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[9] For example, for canned Peruvian anchovies the fish account for around 10% of the production costs, as the majority of raw material cost is made up of metal and vegetable oil.

the long-term price spread between these two commodities and opens up options for new uses of small pelagic fish if so desired and, more important, if sufficient value is created by alternative uses.

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