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Giovanni Guastella, Daniele Moro, Paolo Sckokai and Mario Veneziani

Investment behaviour of EU arable crop farms in selected EU countries and the impact of policy reforms

ABSTRACT

This deliverable provides a comparative analysis, among selected EU member states, of the investment demand of a sample of specialised field crop farms for farm buildings, machinery and equipment as determined by different types and levels of Common Agricultural Policy support. It allows for the existence of uncertainty in the price of output farmers receive and for both long- and short-run determinants of investment levels, as well as for the presence of irregularities in the cost adjustment function due to the existence of threshold-type behaviours. The empirical estimation reveals that three investment regimes are consistently identified in Germany and Hungary, across asset and support types, and in France for machinery and equipment. More traditional disinvestment-investment type behaviours characterise investment in farm building in France and the UK, across support types, and Italy for both asset classes under coupled payments. The long-run dynamic adjustment of capital stocks is consistently and significantly estimated to be towards a – mostly non-stationary – lower level of capitalisation of the farm analysed. By contrast, the expected largely positive short-run effects of an increase in output prices are often not significant. The effect of CAP support on both types of investment is positive, although seldom significant, while the proxy for uncertainty employed fails to be significant yet, in most cases, has the expected effect of reducing the investment levels.

Keywords: Farm Investment, Threshold Models, FADN Data, Common Agricultural Policy.

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FACTOR MARKETS Coordination: Centre for European Policy Studies (CEPS), 1 Place du Congrès, 1000 Brussels, Belgium Tel: +32 (0)2 229 3911 • Fax: +32 (0)2 229 4151 • E-mail: info@factormarkets.eu **• web:** www.factormarkets.eu

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Giovanni Guastella, Daniele Moro, Paolo Sckokai and Mario Veneziani*

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1. Introduction

A farmer's decision to invest in physical capital (i.e., farm buildings, machinery and equipment) might be the result of economic considerations regarding the likely difference between the purchase and resale price of an asset (Johnson, 1956) as well as the uncertain nature of farm output price and government support (i.e., Common Agricultural Policy (CAP) provisions in the European Union (EU)) (Serra et al. 2009; Boetel et al., 2007). The latter might either influence relative prices through coupled support and/or increase the contribution of non-output related income to total farm income through decoupled subsidies. Both types of subsidy might relax existing budget or credit constraints (Sckokai, 2005) and/or diminish price/revenue uncertainty, resulting in higher physical investment. Nonetheless, the decision to avoid investing may still be optimal if irregularities in the adjustment cost function arise.

Drawing on Serra et al. (2009), the present contribution estimates a reduced form of investment demand function for two asset classes allowing for threshold-type behaviours compatible with a number of capital market imperfections (i.e., differences between an asset's purchase and resale price (Johnson, 1956); asymmetries in fixed capital adjustment costs (Abel and Eberly, 1994); real option (Huttel et al., 2010)) in an attempt to explain the frequent occurrence – in farm level data – of zero and negative gross investment levels. It does so by carefully implementing the threshold regression model developed by Hansen (1996, 1999, 2000) to endogenously and consistently determine and test whether the investment model is characterised by multiple, rather than single, equilibria. This would in turn highlight the optimal nature of the recorded investment values rather than suggesting that realised investment values strictly depend on the presence of imperfections in a number of connected markets. Since conditions and constraints are likely to vary significantly across the EU, also due to the different implementation of decoupled subsidies between old and new member state (MS), a comparative analysis of agricultural investment might be worth pursuing. In fact, while the existing literature adopting these theoretical and empirical tools has mainly focused on only one geographic region at a time (Boetel et al., 2007; Serra et al., 2009) this work provides simultaneous evidence of the effect of CAP provisions on the asset classes of interest for selected EU MS. Moreover, relying on the long-term time span covered by the Farm Accountancy Data Network (FADN) dataset, we are also able to highlight the different effects on agricultural investment of coupled and decoupled subsidies, the former implemented until the 2005 Fischler reform of the CAP, the latter currently in place in the form of the Single Farm Payment (SFP).

In what follows, the theoretical model is laid out; section 3 details the data management to construct the variables of interest while section 4 describes the estimation procedure. Section

* Giovanni Guastella and Mario Veneziani are at the Dipartimento di Economia Agroalimentare e DiSES, Università Cattolica del Sacro Cuore, Piacenza, Italy. Daniele Moro and Paolo Sckokai are at the Dipartimento di Economia Agroalimentare, Piacenza, Italy.

5 presents the econometric estimates for the two asset classes and types of CAP support of interest while the last section provides some concluding remarks.

2. Theoretical model

Borrowing from Sckokai and Moro (2009) the problem of a farmer's decision on the optimal quantity of investment to undertake originates from the objective to maximise – over an infinite horizon – the farm's discounted value by solving the problem:

$$\begin{gathered} J(\cdot) = \max_{I} \int_0^\infty e^{-rt} u(A, \sigma_A^2) \\ s.t. \dot{k} = (I - \delta k) \end{gathered} \tag{1}$$

where $u(A, \sigma_A^2)$ is a farmer's utility function defined on A, the farm's wealth, and σ_A^2, the variance of the farm's wealth. In turn, A may be specified as $A = A_0 + \bar{p}y - wx - ck + S$ where A_0 is a farm's initial wealth, y is a farm's aggregate output - deriving from the production function $y = f(x, k, I)$ defined over the quantity of an aggregate variable input x, the units of capital stock k and the level of gross investment I – priced at price p. Assuming the market output price p is a random variable, its mean value $\bar{p}$ affects the farmer's decision through its effect on the value of farm's assets while the price variance σ_p^2 is the source of uncertainty regarding the level of farm's assets since $\sigma_A^2 = f(\sigma_p^2)$. Moreover, the variable input x is priced at its market price w while c is the capital rental price and r is the discounting factor which can be well approximated by the interest rate. The specification of farm's wealth A includes the amount of subsidies received by the farm S. Given our interest in evaluating and comparing the effects of different types of CAP support (i.e., coupled and decoupled payments), the specification of S will vary according to the years for which the data are analysed.

The Hamilton-Jacobi-Bellman equation associated to (1) can be stated as:

$$rJ = \max_{I} \{u + J_k(I - \delta k)\} \tag{2}$$

where J_k is the first derivative of $J(\cdot)$ with respect to capital (Sckokai and Moro, 2009; Serra et al. (2009)). Assuming the existence of an interior solution, the shadow value of capital equals the marginal adjustment cost (Sckokai and Moro, 2009). In turn, the first derivative of (2) with respect to capital yields the investment demand equation represented in implicit form as:

$$\dot{k}(r, A_0, \bar{p}, w, c, S, \sigma_p^2, k) \tag{3}$$

While the capital adjustment cost function arising from (3) is strictly positive and increasing, the literature on investment under uncertainty has been increasingly interested in modelling more realistic behaviours of investment dynamics. In particular, Abel and Eberly (1994) have proposed a theoretical model able to account for differences between the asset's purchase and resale price, asymmetries in the fixed capital adjustment costs and a kinked adjustment costs function at the origin. While capital investment remains characterised by a non-decreasing relationship of the asset's shadow price J_k, a threshold-type behaviour emerges. In fact, optimal investment is expected to be negative (positive) for a value of J_k smaller (larger) than a lower (upper) threshold level while it is expected to be zero for values of the asset's shadow price between the two thresholds (Serra et al., 2009).

Applications of the Abel and Eberly (1994) extension of the traditional investment model under uncertainty appear to be quite rare in the agricultural economics literature. The first contribution could be traced back to Boetel et al. (2007), who investigated the effect of asset fixity, investment asymmetry and the possible existence of a sluggish regime in the demand for a quasi-fixed input in US hog production. It is interesting to note that Boetel et al. (2007) allow the constant and the coefficient for the capital stock at the previous period to vary across the three expected regimes while they circumvent the problem of the shadow price of capital being unobservable, assuming there exists a mapping function into farm output prices. Moreover, Boetel et al. (2007) are the first to employ the methodology suggested by Hansen (1996, 1999, 2000) to estimate an investment function exhibiting a threshold-type behaviour. Huttel et al. (2010) rely on the model in Abel and Eberly (1994) to account for additional costs relating to the financial structure of the farm, due to the existence of imperfections in the capital markets, and their possibility of yielding zero investment rates. The latter constitute an optimal reaction to the existence of investment irreversibility and uncertainty about future prospects. Serra et al. (2009) extend the contribution in Boetel et al. (2007) allowing for both the capital stock and output price variable to determine regime-specific speeds of adjustment to a long and short run capital endowment, respectively. Serra et al. (2009) are interested in studying the impact of a decoupled-type transfer on investment decisions of a sample of Kansas farms. Since output prices determine a short-run adjustment in investment levels, Serra et al. (2009) assume the existence of a function mapping the unobservable shadow price of capital J_k into the lagged value of per hectare net farm income. Although the specification of the structural value function $J(\cdot)$ would allow us to obtain the final form of the model in (3), Serra et al. (2009) avoid the many difficulties associated to its estimation subject to threshold effects by specifying its reduced-form counterpart as:

$$\dot{k} = \beta_1' x I\left(J_k \le J_k^l\right) + \beta_2' x I\left(J_k^l < J_k \le J_k^u\right) + \beta_3' x I\left(J_k^u < J_k\right) \tag{4}$$

where the βs are the vectors of parameters' estimates, the one for the exogenous explanatory variables is defined as $x = \left(r, A_0, \bar{p}, w, c, S, \sigma_p^2, k\right)$, with all the variables in parentheses defined as above, and $I(\cdot)$ is an indicator function taking the value of 1 if the condition in parentheses is met.

3. Data

This paper estimates the investment model, subject to threshold effects, on a sample of specialised arable crop farms drawn from those subject, every year, to the survey each MS carries out – on behalf of the European Commission (EC) – as part of the FADN initiative intended to collect relevant economic information from agricultural holdings in the EU. Specialised field crop farms are those classified, according to their main output, as “specialist cereals, oilseeds and protein crops (COP)”, “general field cropping” and “mixed cropping” (TF8=13 or TF8=14 or TF8=60). Although this paper aimed to provide a consistent evaluation of the impact of CAP subsidies on the level of investment in France (FR), Germany (DE), Hungary (HU), Ireland (IE), Italy (IT), Poland (PO), the Netherlands (NL) and the United Kingdom (UK), we were unable to obtain a sample size for IE and the NL allowing for profitable estimation, given the very limited number of farms in the specialisation of our interest. Moreover, we failed to secure a long enough time-series for the cost of buildings in PO which prevented us from calculating an appropriate rental price of capital. Therefore, the present analysis investigates investment behaviours for FR, DE, HU, IT and the UK only.

Since the empirical model employed (Hansen, 1999) has been developed only for balanced panel data sets and, with the exception of the new member states (NMS), the coupled and decoupled CAP subsidy regimes have been in place before and after 2005, respectively. Due to the heavily imbalanced nature of the FADN panel dataset, we have been forced to construct two 4-year datasets (2001-2004 and 2005-2008) for every country of interest.[1] A four-year time span appears sufficient to justify the adoption of the methodology while ensuring a reasonable operating sample size. Throughout the procedure developed to prepare the dataset for estimation, consistency and quality checks on the FADN data are carried out. To begin with, we delete observations that report zero total Utilised Agricultural Area (UAA), namely SE025=0.

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[1] The UK is an exception to this rule since, for this country, the year 2001 appears problematic, providing very few observations. Therefore, for the UK we construct the 1997-2000 and 2005-2008 datasets.

Following Serra et al. (2009), we allow investment levels to adjust – in a regime-dependent manner – due to short-run variations in a synthetic measure of the farm's output price. Since specialised field crop farms are multi-output enterprises both in crop production and in other complementary outputs (i.e., livestock and milk production), a synthetic output price index is constructed out of the data available in the FADN.

First of all, total output value – for the productions listed in Table 1 – is defined as the sum of sales (suffix SA to the K### product of interest), farmhouse consumption (suffix FC to the K### product of interest) and farm use (suffix FU to the K### product of interest). Since the FADN dataset provides information on the related amount of output quantity (suffix QQ to the K### product of interest), unit-value prices are calculated dividing the total output value previously constructed by the relevant K###QQ. The ensuing unit values suffer from two major problems: very small (and large) minimum (and maximum) values and a high incidence of missing values due to the division by zero K###QQ.[2] The first issue is tackled trimming each yearly price series at the bottom and top 5% to exclude outliers. The second issue has been tackled replacing, for every year, missing or zero prices with their respective averages at two different geographical levels. We exploit the FADN's representativeness at the division level (A1) and perform the substitution employing average prices calculated at the A1 level.[3] In case instances of zero and/or missing prices persisted after this first round of substitutions, they have been replaced with national yearly averages. We have built the list of outputs to be aggregated into a more complex index working on the FADN data for IT, since it is the country we know more about and for which arable crop farms should have a large output variety. Later on, we have adapted it for other countries, removing from the output set those that did not achieve a complete price time series after the two substitutions described above. Table 1 summarises the output prices considered for every country of interest. The rationale for this procedure is to ensure that we can aggregate the largest amount of price information in a single output price and exploit the longest time series available in further elaborations.

Table 1. Composition of the single-output farm price, by country

Output	Heading in FADN's Table K	FR	DE	HU	IT	UK
Common wheat and spelt	120	√	√	√	√	√
Durum wheat	121	√		√	√	
Barley	123	√	√	√	√	√
Oats	124	√	√	√	√	√
Grain maize	126	√	√	√	√	
Other cereals	128	√		√	√	
Potatoes	130	√	√	√	√	√
Sugar beets	131	√	√	√	√	√
Permanent pasture	150			√	√	
Rough grazing	151			√	√	
Cow's milk	162	√	√	√	√	
Wool	166	√		√	√	√
Sunflower	332	√	√	√	√	
Soya	333	√		√	√	

Source: Own elaboration on EU-FADN - DG AGRI data for the relevant countries.

[2] This occurs since sales might originate from the stocks of products carried over from the previous accounting year rather than out of current year production.

[3] The exclusion of IE and the NL from the set of countries for which the model can be estimated is further justified by the impossibility of performing this procedure while preserving significant price variability. In fact, for these two countries the A1 geographical level corresponds to the whole country.

Second, the newly filled-out and trimmed series of unit-value prices are multiplied by the respective output quantities (K###QQ) to re-construct output values and circumvent the problem of inconsistent reporting of output values and quantities. In fact, once the re-constructed output values have been added up, across outputs, observations with zero total output are dropped.[4]

The next step builds the share of the newly re-calculated total output value for the single output of interest to the farm's total newly re-calculated output value across the outputs considered. The latter constitute the weights employed to construct the weighted average synthetic price of farm output. Once the average farm output price is calculated, it is turned into a price index divided by the variable's value for the first farm in the sample for the base year.[5]

The price index is employed to derive farm-level expected output price index (*expoutpi*) and its variance (*varoutpi*) drawing on Chavas and Holt (1990). The series of the price index is collapsed at the regional (A1) level, for every year in the sample. To minimise the loss of useful information, we recode the missing regional identifiers (A1) to the closest neighbouring one. We adapt the original Chavas and Holt (1990) methodology to allow for only partially adaptive expectations. In fact, the expectation of the prediction error is obtained averaging the prediction errors for the preceding five years. Due to data limitations in HU, we restrict this informative base to only three years and do not calculate the *varoutpi* since, otherwise, we would be left with only a two-years sample.

The model's explanatory variables include an aggregate measure of the farm's input prices calculated employing a procedure similar to the one outlined for obtaining the output price index. The basic information employed in the procedure is drawn from the Farm Accountancy Data Network's (FADN) Table F, collecting the farm's expenditure for different inputs, and the EUROSTAT input price indexes. In fact, the former does not provide any useful information to calculate input unit-value prices, as suggested above. To improve on the consistency of the estimating dataset, base 2000 indexes (code apri_pi00_ina) have been used as a reference for extracting the price indexes to be associated to the relevant FADN information, as presented in Table 2.

Table 2 shows that seeds, seedlings and food stuff are assumed to have the same price, irrespective of whether they are purchased or produced on the farm. The procedure to obtain a synthetic price for the farm's input requires calculating the share of each of the farm's cost expenditure items, aggregated out of FADN's Table F, to total farm input expenditure and using them to weigh the relevant EUROSTAT price index. Note that the occurrence of negative shares leads to the associated farm dropping out of the sample. The ensuing values divided by the one obtained for the reference farm, already identified upon calculating the output price index, yield the series of the aggregate price index (*inppi*).

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[4] This procedure implies that more importance is attributed to the reporting of output quantities (K###QQ) than the 'original' output value(s) (for instance, K###SA).

[5] The first farm in the sample is selected to ensure that a long-time series of price indexes is obtained while the same farm is still in the sample at a later stage of the dataset manipulation. In turn, the base year for the output price index in FR is 1999, in DE is 2000, in HU is 2004, in IT and in the UK is 1996.

Table 2. EUROSTAT input price indices and related FADN input expenditure headings

EUROSTAT price index for	**EUROSTAT code**	**FADN heading(s)**	**FADN code(s)**
Seeds and planting stock	201000	Seeds and seedlings purchased Seeds and seedlings farm produced	F72+F73
Energy, lubricants	202000	Motor fuels and lubricants Electricity Heating fuels	F62+F79+F80
Fertilisers and soil improvers	203000	Fertilisers and soil improvers	F74
Plant protection products and pesticides	204000	Crop protection products	F75
Veterinary expenses	205000	Other specific livestock costs	F71
Animal food stuff	206000	Purchased food stuff Food stuff produced on the farm	F64+F65+F66+ F67 F68+F69+F70
Maintenance of buildings	208000		G98BV
Other goods and services	209000	Water Insurance Other farming overheads	F81+F82+F84
Machinery and other equipment	211100		G101BV
Buildings	212000		G98BV

Source: Own elaboration on EUROSTAT and EU-FADN - DG AGRI data.

Due to the focus of this paper, we pay particular attention to the determination of the 'quantity' of capital net investment (*inv*) and stock (K_{-1}) pertaining to each farm in the sample.[6] To do so we shall focus on the FADN's Table G, which collects information on farm assets. In this paper we differentiate between the farm net investment and stock ownership of farm buildings (FB) (Heading G98) and machinery and equipment (ME) (Heading G101). The FADN records information for these headings, differentiating between Opening (i.e., beginning of the year) Valuation (suffix BV), Investment before subsidies (suffix IG), Subsidies on Investment (suffix SU), Sales (suffix SA), Depreciation (suffix DP) and Closing (i.e., end of the year) Valuation (suffix CV). The subsidies that might shape a farm's investment in the capital items of interest are the direct payments (DPs) related to Measure Article 20(b)(i) Measure Article 26 "Modernisation of agricultural holdings".[7] Since these subsidies are disbursed conditional on the investment being carried out (i.e., a sort of rebate on investment costs), we model this receipt as a reduction in the cost of capital.

First of all, we calculate a unique measure for the farm net value of investment (uninv) in the relevant asset subtracting DP and SA from IG in FADN's Table G (i.e., G###IG- G###SA-G###DP). Upon calculating the depreciation rate as the ratio of G###DP to G###BV, we drop any observation featuring higher values of the former compared to the latter since, otherwise, depreciation rates larger than one would occur. Missing values originating from the ratio of zero DP to zero BV of each capital item are recoded to zero for economic consistency and to avoid further reductions in sample size. Since the value of DP does not appear – at least in IT – to be recorded reflecting standard accounting practices and displays quite extreme values, we attribute to each farm the sample's median value of the calculated depreciation rates in any given year for which we have data (*meddepr*). The share of investment subsidies (*shinvsub*), out of the value of gross investment, is calculated dividing

[6] The reader should be able to verify that this 'quantity' of capital is, in fact, the real value of capital items.

[7] We thank Prof. Ewa Rabinowics for pointing this out to us.

G###SU by G###IG and is cleaned up from unrealistic values (i.e., greater than or equal to 0.9) and is recoded to zero in case the farm has not reported any IG and, hence, was not entitled to any SU. The rental price of the concerned type of capital (*rprc*) is calculated as:

$$rprc = [pi * (1 - shinvsub)] * \left(\frac{rirate}{100} + meddepr\right) \quad (5)$$

where *rirate* is a measure of the national real interest rate, typically calculated as the difference between the yield of a government bond of a reasonable length and a measure of the nation's inflation rate (growth rate of a consumer price index). Table 3 collects country-specific proxies for both variables.

Table 3. Variables' proxies employed in the calculation of the national rental prices of capital

Country	Interest rate proxy	Inflation rate proxy
FR	10 years benchmark rate on loans to non-financial corporations	Year-on-year change in the harmonised consumer price index
DE	Yields on debt securities outstanding issued by residents, public debt securities	Year-on-year change of the consumer price index, seasonally adjusted
HU	Yearly average of the monthly average agreed interest rate on loans to non-financial corporations weighted by the amount of outstandings	Year-on-year change of the yearly average consumer price index
IT	Average gross yield of long-term public debt securities (BTP) listed on the Italian Stock Exchange	Year-on-year change of the yearly average of the general consumer price index for the whole community
UK	Annual average of four UK banks' base rates	Year-on-year change of the consumer price index

Source: authors' elaboration on national Central Banks and Statistical offices information.

Note that *pi*, the EUROSTAT price index for the cost of FB is the simple average of the one for maintenance of buildings and buildings from Table 2. Once the *rprc* has been calculated, the 'lagged' stock of capital (K_{-1}) is obtained as the ratio of G###BV to the relevant *rprc*.[8] Similarly, the *inv* in the relevant asset class is obtained as the ratio of *uniinv* to its *rprc*. The *rprc* of every asset class we are interested in is employed to standardise the value of the *expoutpi*, *inpi*, coupled and decoupled subsidies, farm wealth and lagged income employed in the empirical models for the investment in FB and ME. The square of the relevant *rprc* is employed to standardise the *varoutpi* in each model.

Following Serra et al. (2009) we model investment demand to depend upon a measure of lagged wealth in thousands of euros (*wealth*) and per hectare income (*income*).[9] The former is a measure of total assets (G103), excluding FB and ME, minus total debt (H106). This value is obtained relying on BVs to avoid additional reductions in sample size. Income is defined as the difference between the value of total output (SE131), the re-calculated expenditure on inputs and the reported value of the depreciation on the concerned assets (G98DP and G101DP). The latter is turned into its per hectare counterpart by dividing by the total UAA (SE025) and then the resulting series is lagged one year. The amount of subsidies the farm receives, in thousands of euros, is the sum of total subsidies on crops (SE610) and on livestock (SE615) for coupled support (*coupsub*) and decoupled payments (SE630) for decoupled support (*decsub*).

[8] We exploit the availability of both the beginning and end-of-year values of capital and use the former to provide the basis for obtaining K_{-1}. While we are aware that this is not a truly 'lagged' value, this helps to preserve sample size in the very unbalanced FADN dataset.

[9] We thank Laure Latruffe and Laurent Piet for the suggestion of accounting for farm size standardising the amount of farm income by total UAA (SE025).

The mean and standard deviations for the variables employed to estimate the models for the investment demands for FB and ME are presented in Table 4 and 5, respectively.

Table 4. Summary statistics for the model of investment demand for FB

	2001–2004¥				2005-2008				
Variable	**FR**	**DE**	**IT**	**UK†**	**FR**	**DE**	**HU**	**IT**	**UK**
inv	-16.09 (958.93)	8.31 (4435.26)	-279.74 (366.82)	-96.81 (1064.73)	-80.28 (1030.59)	411.67 (6074.28)	437.34 (6981.55)	-97.20 (5339.84)	40.04 (1444.48)
income	17.74 (30.572)	66.67 (77.15)	177.17 (277.97)	18.80 (25.33)	29.39 (52.64)	79.75 (102.64)	37.95 (70.52)	144.38 (333.20)	21.140 (29.71)
K_{-1}	1389.75 (2839)	9105.00 (18785)	6788.76 (9524)	2054.94 (3498)	1817.75 (5578)	7410.71 (14074)	6850.77 (23760)	9335.36 (33574)	1749.01 (2912)
expoutpi	0.0732 (0.0503)	0.111 (0.016)	0.1313 (0.0166)	0.0330 (0.0068)	0.0895 (0.0370)	0.1353 (0.0524)	0.1773 (0.0778)	0.1473 (0.0457)	0.0415 (0.0103)
inpi	6.1010 (3.6121)	11.980 (0.655)	13.9636 (0.9531)	4.5212 (0.5080)	6.7500 (2.3755)	13.7662 (2.5000)	12.9972 (3.5125)	14.9314 (1.7147)	5.9491 (1.2712)
varoutpi	0.9940 (22.7930)	1.088 (1.140)	1.2620 (0.7010)	0.1990 (0.2730)	2.0060 (2.9160)	5.3100 (6.3160)	§	7.7200 (12.9920)	0.5140 (0.3500)
coupsub	2.6655 (2.4225)	7.424 (12.550)	3.1291 (5.6061)	3.3945 (2.9097)	§	§	1.2947 (3.1297)	§	§
decsub	§	§	§	§	1.6938 (2.4008)	6.0378 (8.6204)	2.6892 (5.1027)	2.7190 (7.7950)	2.9112 (2.4886)
wealth	3.0135 (9.0275)	71.990 (93.502)	80.1332 (164.1416)	58.3491 (61.2950)	3.8531 (9.2979)	84.2990 (95.8574)	9.9175 (30.0271)	122.0474 (326.4338)	62.1945 (70.9741)
n	1636	1180	280	404	1400	1212	828	2280	384

Notes: *inv*, *income*, K_{-1}, *coupsub*, *decsub* and *wealth* are measured in constant Euros, *inppi*, *expoutpi* and *varoutpi* are index numbers, the summary statistics for the series *inppi* and *varoutpi* are presented multiplied by 100 and 10,000, respectively, to avoid computational issues when estimating the model in Matlab;; § denotes Not Applicable; ¥ the countries analysed over the period 2001-2004 do not include HU since, back then, it was not part of the EU; † the UK sample to analyse the role of *coupsub* covers the period 1997-2000.

Source: authors' calculations based on manipulated EU-FADN - DG AGRI data.

Table 5. Summary statistics for the model of investment demand for ME

	2001–2004¥				2005-2008				
Variable	**FR**	**DE**	**IT**	**UK†**	**FR**	**DE**	**HU**	**IT**	**UK**
inv	-164.06 (1240.19)	-259.73 (1304.44)	-361.37 (893.26)	-394.90 (1548.50)	-44.79 (1157.67)	17.77 (1487.67)	-42.56 (1782.96)	-358.04 (873.16)	236.52 (1661.81)
income	10.2005 (18.6680)	20.6730 (23.9800)	76.3971 (119.9770)	16.5257 (23.1052)	15.2264 (27.7006)	25.3539 (32.7532)	10.9587 (19.9063)	58.7533 (135.2087)	19.4698 (27.5666)
K_{-1}	3162.74 (3486.74)	4403.044 (4550.81)	3295.404 (3075.22)	7493.17 (6459.61)	2953.22 (3000.45)	4265.42 (5069.59)	3598.49 (6100.80)	3642.09 (5151.64)	7417.64 (9291.71)
expoutpi	0.0418 (0.0233)	0.0344 (0.0046)	0.0567 (0.0086)	0.0288 (0.0041)	0.0464 (0.0118)	0.0429 (0.0161)	0.0507 (0.0168)	0.0601 (0.0175)	0.0383 (0.0076)
inpi	3.5119 (2.5031)	3.7124 (0.1226)	6.0274 (0.6568)	3.9783 (0.1959)	3.5236 (0.2615)	4.3613 (0.6768)	3.8056 (0.4905)	6.0989 (0.5111)	5.4913 (0.8767)
varoutpi	0.2050 (2.1320)	0.1050 (0.1110)	0.2370 (0.1360)	0.1440 (0.1840)	0.4880 (0.4640)	0.5350 (0.6290)	§	1.2380 (2.0220)	0.4340 (0.2770)
coupsub	1.5336 (1.6030)	2.2969 (3.8759)	1.3665 (2.5443)	2.9874 (2.5275)	§	§	0.3923 (0.9500)	§	§
decsub	§	§	§	§	0.8538 (0.8176)	1.8932 (2.5908)	0.7898 (1.4555)	1.1143 (3.2463)	2.7054 (2.2978)
wealth	1.6221 (4.7398)	22.3253 (28.9849)	34.9133 (72.5745)	51.5233 (53.8386)	2.0072 (4.8646)	26.7501 (30.3583)	2.8917 (9.0746)	50.0360 (133.8025)	57.6573 (65.2601)
n	1636	1180	280	404	1400	1212	828	2280	384

Notes: *inv*, *income*, K_{-1}, *coupsub*, *decsub* and *wealth* are measured in constant Euros, *inppi*, *expoutpi* and *varoutpi* are index numbers, the summary statistics for the series *inppi* and *varoutpi* are presented multiplied by 100 and 10,000, respectively, to avoid computational issues when estimating the model in Matlab;; § denotes Not Applicable; ¥ the countries analysed over the period 2001-2004 do not include HU since, back then, it was not part of the EU; † the UK sample to analyse the role of *coupsub* covers the period 1997-2000.

Source: authors' calculations based on manipulated EU-FADN - DG AGRI data.

4. The empirical model

Following Serra et al. (2009), empirical estimation is based on the threshold regression framework proposed, in three seminal papers, by Hansen (1996, 1999, 2000), which develop the statistical and asymptotic distribution theory and clarify the procedure to implement the model on a balanced panel dataset. According to Hansen (2000), threshold models can be employed to estimate models of separating and multiple equilibria, to investigate the opportunity to empirically split the estimating sample on the basis of continuous variables and to parsimoniously estimate functions in a non-parametric fashion. Moreover, the same framework can accommodate, as special cases, more complex models such as mixture, switching, Markov switching and smooth transition models. Due to the possibility of different behaviours, both in the short and long run, of capital investment, this methodology appears particularly suited to test the existence of different investment regimes in the countries of our interest.

Since the methodology does not exist for unbalanced panel data, we rely on the balanced panel formulation in Hansen (1999). The structural equation for the farm's demand for the relevant type of investment (y), in the presence of two regimes, can be represented as:

$$y_{it} = \mu_i + \beta_1' x_{it} I(q_{it} \leq \gamma) + \beta_2' x_{it} I(q_{it} > \gamma) + e_{it} \tag{6}$$

where the indexes i and t represent farm and time, respectively, x_{it} is the vector of the model's non time-invariant covariates and β' the related regime-dependent coefficients' vector, μ_i is a farm-specific effect and e_{it} an independent identically distributed (iid) error term with zero mean and finite variance σ^2. The model's coefficients are regime-dependent because of the indicator function $I(\cdot)$ being one if the threshold variable q_{it} is smaller than or equal to its threshold value γ, and zero otherwise. The equality between the slopes β_1' and β_2' is a testable proposition which, if violated, supports the existence of a significant threshold effect identifying two regimes and, in turn, two different investment behaviours. Note that, as in Hansen (1999) and Serra et al. (2009), some of the coefficients in the β' vector might be restricted to be equal across the regimes such that the role of the related covariates x_{it} is regime-independent. In the present application, the only regime-dependent variables are K_{-1} and *expoutpi* expressing the pattern of long and short-run adjustment of capital quantities, respectively. Moreover, and contrary to Serra et al. (2009), we follow Hansen (1999) and allow the threshold variable q_{it} – in our case *income* – to be included in the vector of model's covariates.

If a threshold, hence two regimes, existed (i.e., the model in (6) has a higher explanatory power than a simple 'pooled' or zero-threshold model), multiple thresholds may exist, leading to the identification of more than two regimes/behaviours/equilibria. The empirical identification of the 'real' number of thresholds appears to be a matter of iteratively estimating and testing for the presence of one more threshold than in the previous iteration of the model. Nonetheless, we restrict the maximum number of allowed thresholds to two such that we can associate to this empirical evidence the existence of, respectively, the disinvestment (Dis), zero-investment (ZInv) and investment (Inv) farm behaviour. Therefore, the double-threshold model can be written as:

$$y_{it} = \mu_i + \beta_1' x_{it} I(q_{it} \leq \gamma_1) + \beta_2' x_{it} I(\gamma_1 < q_{it} \leq \gamma_2) + \beta_3' x_{it} I(q_{it} > \gamma_2) + e_{it} \tag{7}$$

with $\gamma_1 < \gamma_2$. The following sub-section summarises the estimation and testing procedures for the single-threshold model (6) first and then extends them, where necessary, to the double-threshold one in (7).

4.1 Estimation and testing

Hansen (1999) suggests removing the farm-specific effect μ_i in (6) taking the difference between each variable in the specification and its respective average over the time index t

(i.e., de-meaning with unit specific time means). The resulting variables will be denoted by the superscript *:

$$y_{it}^{*} = \beta_1' x_{it}^{*} I(q_{it}^{*} \leq \gamma) + \beta_2' x_{it}^{*} I(q_{it}^{*} > \gamma) + e_{it}^{*} \tag{8}$$

The formulation in (8) can be compacted rewriting x_{it} and β as:

$$x_{it}^{*}(\gamma) = \begin{pmatrix} x_{it}^{*} I(q_{it}^{*} \leq \gamma) \\ x_{it}^{*} I(q_{it}^{*} > \gamma) \end{pmatrix}$$

$$\beta = \begin{pmatrix} \beta_1 \\ \beta_2 \end{pmatrix}$$

such that (8) reads as:

$$y_{it}^{*} = \beta' x_{it}^{*}(\gamma) + e_{it}^{*} \tag{9}$$

Stacking the observations for individual i over t=2,...T and then over all individuals, the notation can be simplified to obtain:

$$Y^{*} = X^{*}(\gamma)\beta + e^{*} \tag{10}$$

If γ were known, the vector of slope coefficients β could be estimated by ordinary least squares (OLS) as:

$$\hat{\beta}(\gamma) = \left(X^{*}(\gamma)' X^{*}(\gamma)\right)^{-1} X^{*}(\gamma)' Y^{*} \tag{11}$$

but, since it is unknown, it needs estimating before moving on to determine $\hat{\beta}$. Hansen (2000) recommends estimating γ by least squares (LS) minimising the concentrated sum of squared errors $S_1(\gamma)$ from (11):

$$\hat{\gamma} = argmin\, S_1(\gamma) \tag{12}$$

where $S_1(\gamma) = \widehat{e^{*}}(\gamma)' \widehat{e^{*}}(\gamma)$ and the vector of regression residuals $\widehat{e^{*}}(\gamma)$ is obtained from (10) or, equivalently and more conveniently, $S_1(\gamma) = Y^{*\prime}\left(I - X^{*}(\gamma)'\left(X^{*}(\gamma)'X^{*}(\gamma)\right)^{-1} X^{*}(\gamma)'\right) Y^{*}$. Hansen (1999) remarks that it is undesirable for the optimal threshold value $\hat{\gamma}$ to separate too few observations in the two regimes, while Hansen (2000) notes that a monotonic transformation on the threshold parameter would render it unit free and constrain it in the [0,1] interval. Nonetheless, the pointwise test statistics appear to be very sensitive to extreme values of the transformed threshold. To overcome both problems, borrowing from the changepoint literature,[10] Hansen (2000) seems to endorse Andrews (1993) suggestion to restrict the range of the transformed threshold parameter to [0.15, 0.85]. This is achieved eliminating the smallest and largest 15% of the values of the distinct (unique) values assumed by the threshold variable q_{it}. This appears to be a well-established choice of the trimming parameter (Hansen (1999), Dang et al. (2012)) we comply with. The search for $\hat{\gamma}$ would entail the calculation of $S_1(\gamma)$ for every unique value of the threshold variable q_{it}, after the trimming percentage has been applied. To avoid the ensuing computational burden, especially when dealing with very large datasets, Hansen (1999) suggests obtaining equivalent results implementing a grid search over specific quantiles (400 in total) of the distribution of q_{it}.

Once the value of the threshold has been estimated, it is necessary to determine whether its effect is statistically significant, i.e., whether the sets of coefficients' estimates are different across regimes. Therefore, the following test for the significance of one threshold can be designed:

$$\begin{cases} H_0: \beta_1 = \beta_2 \\ H_a: \beta_1 \neq \beta_2 \end{cases} \tag{13}$$

[10] It should be noted that the threshold regression literature builds on the findings and results obtained by the changepoint literature.

The likelihood ratio statistic to evaluate the hypotheses (13), LR_1 is:

$$LR_1 = \frac{S_0 - S_1(\hat{\gamma})}{\hat{\sigma}^2} \tag{14}$$

where S_o is the concentrated sum of squared errors associated to the zero threshold model (i.e., pooled OLS), $S_1(\hat{\gamma})$ is the model's concentrated sum of squared errors at the estimated threshold value $\hat{\gamma}$ and $\hat{\sigma}^2 = \frac{1}{n(T-1)} S_1(\hat{\gamma})$, with n and T being the number of observations per time period and the number of time periods, respectively. The testing procedure in (13) is complicated by the evidence that, under H_o, the threshold γ is not identified; the asymptotic distribution of LR_1 is non-standard, strictly dominates the χ_k^2 one, but depends on the sample moments, such that it cannot be tabulated. Hansen (1996) proves that a bootstrap procedure obtains a first-order approximation of the asymptotic distribution of LR_1 such that the p-values constructed from the bootstrap itself are asymptotically valid. Hansen (1999) is the reference for a recommended implementation of this bootstrap procedure. Once the bootstrap procedure has been written, the researcher is left with choosing the number of repetitions to carry out. Andrews and Buchinsky (2001; 2002) provide Monte Carlo simulations for a three-step procedure to select the 'optimal' number of bootstrap repetitions to obtain a statistic of interest (i.e., standard error, confidence intervals, tests and p-values) – bootstrapped from a finite number of repetitions – which is close to the ideal one that would be obtained running an infinite number of repetitions. Drawing on this applied and simulation work and given our interest in both bootstrapped confidence intervals and p-values, we decided to run 5,000 repetitions.[11]

Having verified that the estimated threshold has significant effects, $\hat{\gamma}$ is a consistent estimate of the threshold's (γ) true, and unknown, value γ_0 (Hansen, 2000). Nonetheless, a test for the precision of such an estimate can be specified as:

$$\begin{cases} H_0: \gamma = \gamma_0 \\ H_a: \gamma \neq \gamma_0 \end{cases} \tag{15}$$

Hansen (2000) suggests carrying out the test forming a 'no-rejection region' for the LR_2 defined as:

$$LR_2(\gamma) = \frac{S_1(\gamma) - S_1(\hat{\gamma})}{\hat{\sigma}^2} \tag{16}$$

where the terms are as previously defined. The LR_2 statistic evaluates whether the true value of γ is indeed at $\hat{\gamma}$ or in its neighbourhood such that a confidence interval for the threshold's estimate, at a specified confidence level, is constructed.[12] The size of this confidence interval complements the significance of the threshold's effect in suggesting the 'strength' of separation between the identified regimes. The LR_2 statistic converges, in distribution, to a value ξ (i.e., a random variable) with non-standard – yet free of nuisance parameters – cumulative distribution function (CDF) $P(\xi \leq x) = \left(1 - exp\left(-\frac{x}{2}\right)\right)^2$. Since this CDF is pivotal it can be used to form valid asymptotic confidence intervals, based on – for every confidence level α – the associated critical values: $c(\alpha) = -2log\left(1 - \sqrt{1-\alpha}\right)$. The H_o in (11) is rejected, at the asymptotic level α, if $LR_2(\gamma)$ exceeds $c(\alpha)$. Since the asymptotic confidence interval for γ corresponds to the 'no-rejection region' of confidence level *1-α* (i.e., the values

[11] Andrews and Buchinsky (2001) express the disconcerting finding that the number of bootstrap repetitions carried out in most econometric applications is much smaller than the one required to achieve a sufficiently accurate simulated statistic of interest. Most likely due to a large dataset employed and more limited computing power than available today Hansen (1999) in his Matlab code, which we modify to suit our needs, selects 300 repetitions for his bootstraps, which might be deemed small by Andrews and Buchinsky (2001; 2002) standards.

[12] Hansen (1999), in his Matlab code, employs a 95% confidence interval, which we also adopt in our application.

of γ such that $LR_2(\gamma) \leq c(\alpha)$), Hansen (1999) suggests identifying them by plotting the values of $LR_2(\gamma)$ against γ and a flat line drawn at $c(\alpha)$. Note that Hansen (2000) remarks that the construction of the threshold's confidence interval rests on the assumption that e^* is homoschedastic, or at least that conditional heteroschedasticity is not regime dependent (Hansen, 1997).[13] Otherwise, the critical values $c(\alpha)$ need to be corrected by multiplying for an estimable – although with some difficulties – parameter.

The estimates of β, according to (11), appear to depend on the threshold's estimate $\hat{\gamma}$. Yet, Hansen (2000) demonstrates that this link is not of first-order asymptotic importance, resulting in $\hat{\beta}$ being asymptotically normal with an estimable – and possibly heteroschedastic – covariance matrix. Hansen (1999) suggests that large differences between conventional and heteroschedasticity robust OLS standard errors provide evidence in favour of a correction. Nonetheless, we believe a more formal test for the heteroschedastic behaviour of the OLS remainder error term should be run and the ensuing evidence should determine the nature of the presented standard errors. Contrary to much of the extant literature which adheres to Hansen's (1999) 'qualitative' recommendation, of which Savvides and Stengos (2000) constitute an exception, before presenting the model's estimates we run a version of the White (1980) test for heteroschedasticity of the remainder error. The test entails an ancillary regression of the squares of the OLS residuals on the squares of the explanatory variables. In case the null hypothesis of homoschedasticity is rejected, heteroschedasticity robust standard errors are obtained, and presented, employing the Eicker (1967)-Huber (1967)-White (1980) estimator.

Hansen (1999) posits that the procedure to estimate the single threshold model in (6) can be employed also to estimate the double (and higher order) threshold model in (7). This is due to the (multiple) changepoint literature's result that the sequential estimation of a single threshold model yields consistent estimates for the multiple threshold one. In the first stage, a single threshold model is estimated yielding the ensuing sum of squared errors $S_1(\gamma)$, which is minimum at $\widehat{\gamma_1}$. This estimate is consistent also in the case of the double threshold model in (7) for either γ_1 or γ_2, depending on which effect is stronger. Fixing the first threshold's estimate $\widehat{\gamma_1}$, the second threshold's estimate is:

$$\widehat{\gamma_2^r} = argmin\, S_2^r(\gamma_2) \tag{17}$$

where

$$S_2^r(\gamma_2) = \begin{cases} S(\widehat{\gamma_1}, \gamma_2) \;\; if \;\; \widehat{\gamma_1} < \gamma_2 \\ S(\gamma_2, \widehat{\gamma_1}) \;\; if \;\; \gamma_2 < \widehat{\gamma_1} \end{cases} \tag{18}$$

While $\widehat{\gamma_2^r}$ is asymptotically efficient, $\widehat{\gamma_1}$ is not since it has been estimated minimising a sum of squared error calculated in presence of a neglected regime. In turn, $\widehat{\gamma_1}$ can be improved by refining its estimate, holding the second threshold's estimate $\widehat{\gamma_2^r}$ constant, to be:

$$\widehat{\gamma_1^r} = argmin\, S_1^r(\gamma_1) \tag{19}$$

where

$$S_1^r(\gamma_1) = \begin{cases} S\left(\gamma_1, \widehat{\gamma_2^r}\right) \;\; if \;\; \gamma_1 < \widehat{\gamma_2^r} \\ S\left(\widehat{\gamma_2^r}, \gamma_1\right) \;\; if\, \widehat{\gamma_2^r} < \gamma_1 \end{cases} \tag{20}$$

[13] While this appears to be one of the model's stringent assumptions, the empirical literature applying threshold regression models seems to have largely overlooked the issue of regime-dependent heteroschedasticity. Notable exceptions are the empirical application in Hansen (1997) and the promising theoretical model accounting for both regime-dependent heteroschedasticity and endogeneity of the threshold variable in Kourtellos et al (2013). At the moment we have been unable to comply with the recommendation in Hansen (2000), but future extensions will tackle both the existence of regime-dependent heteroschedasticity and endogeneity of the threshold variable.

To test the significance, hence the existence, of two thresholds – against the null hypothesis of only one – a test similar to the one in (13) can be carried out (Hansen, 1999). Being $S_2^r(\widehat{\gamma_2^r})$ the minimising sum of squared error from the estimation of the second threshold and $\hat{\sigma}^2 = \frac{1}{n(T-1)} S_2^r(\widehat{\gamma_2^r})$ its estimated variance, the test statistic $LR_3 = \frac{S_1(\widehat{\gamma_1})-S_2^r(\widehat{\gamma_2^r})}{\hat{\sigma}^2}$ rejects the hypothesis of one threshold, in favour of two, if LR_3 is large such that its bootstrapped p-value achieves statistical significance. Hansen (1999) notes that the null sampling distribution of LR_3 depends on several nuisance parameters that are likely to reduce the precision of the statistic's bootstrapped p-values (compared to those obtained with the same method for LR_1), somewhat limiting the confidence in our findings.

Since the refinement estimators in (17)-(20) have the same asymptotic distribution as the threshold estimate in a single threshold model in (15) and (16), the procedure described for LR_2 can be applied. Therefore, the $(1-\alpha)\%$ confidence intervals for γ_1 and γ_2 include the values of γ such that $LR_2^r \leq c(\alpha)$ and $LR_3^r \leq c(\alpha)$ where

$$LR_2^r(\gamma) = \frac{S_1^r(\gamma)-S_1^r(\widehat{\gamma_1^r})}{\hat{\sigma}^2} \text{ and } LR_3^r(\gamma) = \frac{S_2^r(\gamma)-S_2^r(\widehat{\gamma_2^r})}{\hat{\sigma}^2}$$

5. Results

5.1 Models' estimates

In what follows we present the empirical findings of estimating a two-threshold model, for the countries and type of investment of interest, aiming to distinguish the Dis, ZInv and Inv behaviours. Table 6 provides a synoptic representation of the results, to elicit cross-country comparisons by nature of investment and CAP support.

Table 6. Synoptic table of the models' results

		N° Regimes	K_{-1}	NS/S Equilibrium	*expoutpi*	*coupsub*	*decsub*	*varoutpi*
FR	FB	2	- (***)	NS/S	+ (¥/**)	-	§	- (***)
	ME	3	- (***)	S	+ (***)	+	§	+
	FB	2	- (***)	NS	- (***)	§	+ (*)	+
	ME	3	- (***)	S/NS	- (¥/*/¥)	§	-	-
DE	FB	3	- (***)	S/NS	+ (***)	+ (***)	§	-
	ME	3	- (***)	S/NS	-	+ (***)	§	- (**)
	FB	3	- (**/***/¥)	NS	+ (*/¥/¥)	§	+ (***)	- (***)
	ME	3	- (***)	S/NS	+ (¥/*/¥)	§	+ (**)	+ (*)
HU	FB	3	- (***)	S/NS	-	+	+ (***)	§
	ME	3	- (¥/***/***)	NS	- (***/**/***)	+	+	§
IT	FB	2	- (***)	NS	-	- (***)	§	-
	ME	2	- (¥/***)	NS	-	+	§	-
	FB	3	- (¥/**/**)	NS	- (¥/***/***)	§	+	-
	ME	0	- (***)	NS	- (*)	§	-	- (**)
UK	FB	2	- (***)	S	-	+ (**)	§	-
	ME	0	- (***)	S	+	+	§	+
	FB	2	- (***)	NS/S	+	§	+	-
	ME	0	- (***)	S	+	§	+ (*)	+

Notes: *** denotes statistical significance at the 1% level, ** at the 5% level, * at the 10% level based on two tailed tests; ¥ denotes Not Significant in estimation; § denotes Not Applicable in estimation; regime dependent coefficients (K_{-1} and *expoutpi*) are presented with the sign which occurs more often while significance (their superscript) is presented multiple times if it differs across regimes (Dis, ZInv, Inv).

Source: authors' elaboration on models' estimates.

The evidence in Table 6 suggests that investment demand is clearly characterised by a three-regime behaviour only in DE and HU, irrespective of the actual type of investment and agricultural support under which it occurs. In FR and in the UK, the number of regimes appears to be investment specific with the one in FB being subject to two different behaviours, while the one in ME features all three of them in the former country but none in the latter. Investment behaviour in IT seems to emerge as differentiated in two regimes only when farms benefit from coupled CAP support. Here, the introduction of DPs develops an additional distinctive behaviour for the investment in farm buildings (FB) while it seems to equalise differences when farms decide their investment in ME. To shed light on the nature of the investment behaviour when only one threshold is significant, we shall refer to Tables 7 and 8 which report the models' estimates for every country of interest and for both asset classes. In particular, associating the Dis regime with threshold values smaller or equal to γ_1, the ZInv with values in the $(\gamma_1; \gamma_2)$ range and the Inv one with values greater or equal to γ_2, the variability in the thresholds' significance highlights which regime can be separated from the others.

Therefore, since only the larger threshold is significant in FR for the investment in FB over the period 2001-2004, we posit that it is not possible to distinguish the Dis from the Zinv regimes while a marked separation exists with the Inv one. In the UK, for that period, the Dis regime is distinguishable from a combined ZInv and Inv behaviour for FB while in IT the same holds for both asset types. While the 2005-2008 period appears to be mainly characterised by three distinctive behaviours, investment in FB in both FR and the UK is separated only into two. In particular, in the former, the Inv behaviour is significantly different from the Dis-ZInv combined while in the latter the regime setting itself apart from the other indistinguishable ones is the Dis one.[14]

The coefficient for K_{-1} is expected to range in the [-1; 1] interval determining the rate of adjustment of current capital stock to its long-run equilibrium. In turn, negative values would suggest farms disinvest to reach a lower long-run stock of capital (Sckokai and Moro, 2009) while positive ones should suggest farms are under-capitalised and are required to invest in capital assets to reach their long-run equilibrium. Column 2 of Table 6 provides the, somewhat surprising, and statistically precise testimony that the farms analysed in this study are over-capitalised both in FB and ME as well as over time and across all the possible regimes. Nonetheless, this finding seems to be confirmed by other recent studies employing different methodologies (Petrick and Kloss, 2012; Sckokai and Moro, 2009). Column 3 of Table 6 analyses, similarly to Abel and Eberly (1994) and Serra et al. (2009), the dynamic features of the concerned investment models highlighting whether the estimates suggest a path towards a non-stationary (NS) or stationary (S) long-run equilibrium. This is determined according to whether the coefficients for the regime-dependent long-run adjustment variable (K_{-1}) are closer to 0 or 1 in absolute value, respectively.[15] The most stable and concurring evidence suggests that investment in IT is on the path to NS long-run equilibria, both across asset and support types. Establishing a clear path towards either a NS or a S equilibrium is more difficult in FR, DE and HU, while the investment behaviours in the UK appear mostly geared towards an S long-run equilibrium.

The model allows for short-run adjustments in the stock of capital following fluctuations in *expoutpi*. Contrary to the consistency in both the sign and significance displayed by K_{-1}, the one for short-run changes in investment quantity varies significantly, both in sign and significance. The dependence of investment from *expoutpi* is largely statistically insignificant, it seems to be surprisingly negative in presence of decoupled support in FR and HU while it

[14] Tables 9 and 10 further clarify the nature of the separated regimes and the number of farms which, every year, belong to each of them.

[15] Note that coefficients close to (and larger than) 0.5, in absolute value, pose for a S equilibrium while a smaller (or very small ones) for a NS equilibrium. Moreover, in the case of mixed evidence across the regimes, an uncertain judgment is provided.

also achieves statistical significance – across the types of support – in IT. DE and the UK are the countries featuring a largely positive short-run relationship between investment and *expoutpi*, although seldom significant in the former.

While K_{-1} and *expoutpi* express a behavioural and a market-based reaction of farm investment, respectively, the remaining columns in Table 6 highlight the relationship between investment and a few policy variables. Among them, those regarding the nature of CAP support are the most relevant ones since they proxy the historical and existing forms of agricultural support. Moreover, the dependence of investment on *varoutpi* is of interest given a change in the nature of support might have altered farmers' perception of risk. While the effect of CAP subsidies on investment in both asset types is largely positive, negative dependence arises for investment in FB in FR and IT, with the latter effect being statistically significant at 1%. Decoupled subsidies have a negative, although statistically insignificant, effect on investment in ME in FR and IT. Agricultural support of both types has a consistently positive effect on investment in DE, HU, and the UK with only the one in DE being precisely determined in statistical terms. In turn, *decsub* appear to have a more significant impact on capital investment since its estimated coefficients achieve statistical significance – at conventional levels – in half of the models with a 25% increase compared to *coupsub*. The transition to a decoupled system of agricultural support does not seem to have induced any dramatic change in farmers' attitudes towards capital investment. Exceptions may include the effect of support becoming positive and significant for FB investment in FR, HU and for ME investment in the UK, while significance is lost for FB in the UK. While remaining insignificant, the effect of CAP support turns from positive to negative for ME in FR and IT. In this deliverable we focus on the qualitative dependence of capital investment from CAP policy variables while in deliverable D7.2 we investigate more thoroughly the quantitative dependence calculating the elasticity of investment to the support and simulating *ceteris paribus* changes due to the implementation of the existing and debated CAP reform hypotheses.

The last column in Table 6 contains the estimates of the effect of a measure of the risk of fluctuating output prices farmers experience in each nation, each year of interest. Roughly 70% (11 out of 16) of the estimated models are characterised by a negative coefficient for the effect of risk on both types of investments and across two types of support schemes. Nonetheless, only in roughly a third of these cases (4 out of 11) is the effect statistically significant at conventional levels. In DE and IT the estimated coefficient has the theoretically consistent and expected negative sign, with the ones for the former country being mostly statistically significant. ME investment under a regime of decoupled payment in DE is the only type of investment to be positively and statistically significant – although at 10% – affected by a rise in the *varoutpi*. FR and the UK appear be evenly characterised by positive and negative relationships between risk and investment. In turn, in the UK the signs of this relationship appear to vary by investment type with FB displaying a negative and ME a positive one, respectively.

Tables 7 and 8 (see Annexe) provide results, also for the dependence of the different types of farm investment on the regime-independent variables *inpi*, *wealth* and *income* according to the different types of support received. In particular, while *inpi* was expected to have a negative and statistically significant effect on the level of both FB and ME investment, since a higher expenditure for inputs is likely to result in more limited financial resources available for investment spending. This expectation is realised only for FB investment in DE under coupled support. In fact, the majority of the estimated models features a positive coefficient for the impact of *inpi* on investment with statistical significance – at conventional levels – being achieved more in the presence of decoupled, rather than coupled, subsidies.

A similar finding emerges for the role of *income*, above and beyond the one related to the determination of the regimes, whereby the majority of the related estimated beta have unexpected signs – with occasionally minimal statistical significance – under coupled support while conventional wisdom is restored for all countries, except FR, and investment types over the period 2005-2008. Moreover, estimated coefficients are statistically significant at the 1% level in DE and at the 5% level in IT while in the UK the one for FB achieves statistical significance of the 10% level, while for the ME the investment coefficient is insignificant at conventional levels.

The positive impact of *wealth* on investment levels appears very consistent across countries and asset classes, especially when coupled support was in place. In fact, it is statistically significant – although at 10% and 5% – in FR and DE for both investment types and for FB in the UK. On the contrary, *wealth* is highly statistically significant, while positive, only when investment in FB in HU is investigated over the 2005-2008 period. In the other cases, it remains largely positive while it loses any statistical precision.

The estimated models' explanatory power appears somewhat modest, especially over the period 2005-2008, since the R^2 measure only seldom exceeds 0.5. In fact, a few estimated coefficients are close to conventional statistical significance. These findings, associated with the almost ubiquitous heteroschedasticity of the remainder error, suggest that there still might be significant variability unaccounted for. Future extensions should attempt to control for additional sources of heterogeneity to limit the incidence of heteroschedasticity and, hopefully, to improve the models' explanatory power.

6. Conclusions

The peculiar characteristics of the agricultural sector across the EU MS and over time call for a cross-country analysis of the developmental effects of agricultural policies. Among the different implementations of this analysis, two main declinations can be identified: focusing on specific inputs (i.e., different types of policy instruments) and/or on specific outcomes (i.e., *inter alia* the farmers' choice to specialise in selected productions, the farmers' choice of the amount and type of labour employed and investment decisions). Agricultural investment decisions might be deemed very important since they are targeted at combating the obsolescence of capital assets, which are subject to significant yearly depreciation, as well as advancing the technological dimension of the production technique whenever there is a major innovation in terms of commercial capital goods. Adaptation of the capital stock to increasing farm size or differentiation, as well as to a more modern technological base, is likely to further increase the profitability of the agricultural sector to enhancing countries' economic prospects.

The present deliverable has investigated the role of first coupled and later decoupled CAP subsidies – in a sample of selected MS – in determining the investment demand for farm buildings and the machinery and equipment of specialised arable crop farms, allowing for an uncertainty. Moreover, the reduced-form investment equations have been specified to account for several of the determinants considered in the literature. Applying a theoretical model of investment choice featuring irregularities in the adjustment cost function and an econometric technique capable of identifying the existence of separating equilibria, the deliverable has been able to investigate whether disinvestment, zero investment and positive investment in both asset classes are optimal reactions to the existence of ranges – rather than point values – of shadow asset prices determining actual investment behaviour.

Empirical estimates suggest that the range of zero investment is clearly and consistently identified, on top of the more common disinvestment and investment ones, for Germany across asset classes and CAP support scheme and asset classes for Hungary, since the latter has only been an EU member since the implementation of decoupled support. Three regimes also appear to characterise machinery and equipment investment in France, irrespective of the type of support received. The more common differentiation between disinvestment and positive investment pertains to farm building investment in France and the UK, across support types; to Italy for both asset classes under coupled support. An investment behaviour based on point values of the shadow price of capital is proper to the UK for machinery and equipment, irrespective of the type of support. This evidence might help in devising new regime-specific policy interventions where clear separations are detected.

Since empirical evidence suggests that specialised arable crop farms are disinvesting towards lower levels of long-run capital endowments, frequent anecdotal evidence that agriculture is undercapitalised is not supported by the data. Nonetheless, except for the UK and Germany, these trajectories should lead towards non stationary long-run equilibria implying the possibility of further and different future dynamics. While empirical evidence suggests that the trajectories towards the long-run equilibria appear highly statistically significant, short-run adjustments in capital stocks due to changes in output prices appear largely insignificant and often display a counterintuitive sign.

The association between both types of investment and both types of CAP subsidies is mostly positive, with the dependence on decoupled support barely more significant. The expectation of a negative association between the variance of the expected output price index – accounting for some the uncertainty farmers face in commercialising their outputs – is largely met, although very rarely in a statistically significant manner.

The estimates presented in this deliverable provide the raw data to carry out a more detailed quantitative appraisal of existing levels of investment and support for every identified regime in deliverable 7.4. Moreover, deliverable 7.4 provides the comparative static analysis of the changes in investment levels due to the implementation of the different CAP reform scenarios currently being envisaged and discussed by the European Commission.

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Table 7. Estimates for the threshold investment models during the existence of coupled support (2001-2004)

	FR		DE		IT		UK	
	FB	**ME**	**FB**	**ME**	**FB**	**ME**	**FB**	**ME**
Regime-indep.								
inpi	358.0707***	-107.8006	-411.4871**	904.1538***	4.6058	328.0619	424.0656*	25.3139
	(101.8039)	(95.5483)	(187.1292)	(298.3131)	(16.9884)	(220.1820)	(255.5108)	(535.6026)
varoutpi	-44.4970***	67.6293	-121.6622	-755.2638**	-2.7284	-222.0963	-62.5688	279.5438
	(10.6224)	(59.6204)	(81.6572)	(368.1210)	(22.3605)	(413.0873)	(99.1689)	(334.5365)
coupsub	-4.3330	54.4573	595.9798***	541.5990***	-32.3735***	146.9623	199.3998**	320.4714
	(102.0790)	(170.9632)	(99.3995)	(55.5250)	(9.0984)	(96.6772)	(81.8289)	(269.5284)
wealth	31.4534*	77.5576**	18.2594*	25.4877*	0.1943	-0.9786	20.1043*	-12.8530
	(16.1728)	(30.4643)	(10.0692)	(14.8948)	(0.2605)	(3.0314)	(10.3869)	(19.1931)
income	-4.4661*	-1.4252	3.7162	-6.3046*	0.0086	-0.6148	-4.4220*	4.2557
	(2.4538)	(3.5815)	(2.8390)	(3.4143)	(0.0810)	(1.3441)	(2.2976)	(7.4252)
K_{-1}	§	§	§	§	§	§	§	-0.5704***
								(0.1804)
expoutpi	§	§	§	§	§	§	§	29,583.5858
								(29,265.4275)
Regime-dep.								
K_{-1} *Dis*	-0.2664***	-0.6193***	-0.4617***	-0.5000***	-0.0203***	-0.0499	-0.8853***	§
	(0.0949)	(0.0564)	(0.0995)	(0.0516)	(0.0047)	(0.0356)	(0.1369)	
K_{-1} *ZInv*	§	-0.4992***	-0.6851***	-0.4169***	§	§	§	§
		(0.0485)	(0.1062)	(0.0526)				
K_{-1} *Inv*	-0.4807***	-0.5627***	-0.3411***	-0.2829***	-0.0389***	-0.2417***	-0.6920***	§
	(0.0895)	(0.0640)	(0.0608)	(0.0708)	(0.0033)	(0.0813)	(0.1884)	
expoutpi Dis	5,632.1705	15,426.8277***	47,165.3703***	-7,057.3986	-847.4395	-4,107.6015	-14,759.9722	§
	(3,619.9064)	(4,764.1437)	(14,492.8949)	(15,815.2835)	(1,592.8180)	(13,161.5586)	(19,036.7487)	
expoutpi ZInv	§	12,514.4971**	56,311.1715***	-8,052.5394	§	§	§	§
		(5,186.1710)	(15,181.1422)	(15,525.6148)				
expoutpi Inv	10,500.8845**	23,283.1538***	39,995.3939***	-16,090.3710	-461.1140	-1,928.2455	-20,680.4681	§
	(4,581.6743)	(6,366.3156)	(14,724.8169)	(15,620.4561)	(1,529.5201)	(13,221.8401)	(20,137.3392)	
Thresholds¥								
γ_1	6.8161	2.1246***	54.5174***	6.9888**	11.9697***	37.0353**	14.6218***	§
	[6.5043; 19.9178]	[1.8557; 2.1246]	[53.6037; 54.5174]	[4.1944; 33.2024]	[11.9697; 32.2014]	[32.2584; 46.5037]	[13.5337; 16.5211]	
γ_2	31.5807***	14.5619**	106.0978***	21.8451***	334.1524	134.2202	29.0886	§
	[28.1491; 31.5807]	[11.7269; 14.9436]	[106.0978; 108.1543]	[21.3678; 22.1395]	[48.5538; 355.2406]	[4.8418; 154.5264]	[6.8238; 29.2584]	
Diagnostics								
R^2	0.2153	0.6144	0.4138	0.3252	0.5496	0.1962	0.5707	0.2482
Het Test	15.0235*	119.9096***	44.5959***	30.8945***	6.0287	98.0160***	35.2768***	64.5537***

Notes: estimation carried out in Matlab R2011b; Eicker (1967)-Huber (1967)-White (1980) heteroschedasticity robust standard errors in parentheses when the White test for heteroschedasticity is statistically significant at conventional levels; *** denotes statistical significance at the 1% level, ** at the 5% level, * at the 10% level based on two tailed tests; ¥ estimates are based on results from the refinement estimators (14) and (16) when two thresholds are statistically significant, 95% confidence intervals in square brackets.

Source: authors' estimation on EU-FADN - DG AGRI data employing a modified version of the Matlab code provided by Hansen on his website.

Table 8. Estimates for the threshold investment models during the existence of decoupled support (2005-2008)

	FR		DE		HU		IT		UK	
	FB	**ME**	**FB**	**ME**	**FB**	**ME**	**FB**	**ME**	**FB**	**ME**
Regime-indep.										
inpi	307.3357***	172.5433	414.2898***	-210.1775	-0.3557	1,281.0744***	689.5875**	286.1506**	-114.8194	-133.1749
	(53.0005)	(132.2762)	(135.7757)	(148.9530)	(104.2873)	(258.3974)	(281.5016)	(145.4518)	(172.5395)	(224.7502)
varoutpi	40.8033	-86.2871	-189.6246***	269.1842*	§	§	-21.4179	-26.0962**	-88.1014	190.6704
	(27.0404)	(81.2080)	(57.4738)	(161.9029)			(17.9789)	(11.1152)	(335.7510)	(337.0412)
coupsub	§	§	§	§	341.4439	374.3918	§	§	§	§
					(280.0737)	(231.5646)				
decsub	51.2279*	-121.6025	435.5835***	622.7084**	673.5537***	274.1298	21.8279	-12.7120	49.1979	687.1385*
	(27.2202)	(76.4121)	(104.9752)	(305.5804)	(188.1409)	(329.9545)	(21.3745)	(8.9094)	(179.9548)	(404.1334)
wealth	-6.9621	8.6418	-1.5339	-11.7505	96.9014***	-9.6685	0.3856	0.6714	9.2861	2.4298
	(11.3222)	(37.2889)	(14.6257)	(12.2459)	(31.0677)	(13.7904)	(0.3670)	(0.4877)	(7.3405)	(10.7986)
income	-0.9156	-1.8921	26.3529***	6.0514***	0.5739	4.9588	11.2992**	0.5173**	13.9523*	8.6444
	(1.6169)	(1.5755)	(8.6986)	(2.3207)	(1.5497)	(3.8006)	(5.1967)	(0.2423)	(8.1833)	(5.8430)
K_{-1}	§	§	§	§	§	§	§	-0.0563***	§	-0.6290***
								(0.0123)		(0.1122)
expoutpi	§	§	§	§	§	§	§	-4,400.7930*	§	14,065.1954
								(2,627.8104)		(16,394.0412)
Regime-dep.										
K_{-1} *Dis*	-0.3591***	-0.5355***	-0.2410**	-0.5803***	-0.5748***	-0.1080	0.0066	§	-0.2431***	§
	(0.0301)	(0.0703)	(0.1011)	(0.0654)	(0.0708)	(0.0828)	(0.0432)		(0.0896)	
K_{-1} *ZInv*	§	-0.3883***	-0.2904***	-0.5210***	-0.4583***	-0.4068***	-0.0711**	§	§	§
		(0.0673)	(0.0894)	(0.0601)	(0.0710)	(0.0722)	(0.0281)			
K_{-1} *Inv*	-0.2635***	-0.2838***	-0.0772	-0.3012***	-0.3611***	-0.3120***	-0.0837**	§	-0.4253***	§
	(0.0216)	(0.0714)	(0.1494)	(0.0473)	(0.0517)	(0.0752)	(0.0327)		(0.0635)	
expoutpi Dis	-12,650.4607***	-4,204.5030	26,506.6541*	7,248.6988	-2,696.3963	-21,558.2012***	-7,093.6948	§	17,293.6708	§
	(3,019.5694)	(5,172.6417)	(15,488.2088)	(7,318.9936)	(4,765.7337)	(5,234.0849)	(4,322.3438)		(23,180.7877)	
expoutpi ZInv	§	-7,707.3729*	8,170.6373	12,272.5012*	-4,748.8684	-12,116.5536**	-11,486.0585***	§	§	§
		(4,671.5545)	(8,252.5656)	(6,588.7262)	(4,611.7548)	(5,159.6178)	(3,820.5349)			
expoutpi Inv	-11,980.0311***	-6,196.4155	-10,675.5064	-1,653.1539	-6,052.2831	-15,420.1869***	-22,855.8909***	§	7,624.1822	§
	(3,182.4731)	(4,680.5041)	(7,455.8190)	(6,895.1521)	(4,430.6624)	(5,047.0552)	(6,412.0364)		(20,939.3009)	
Thresholds¥										
γ_1	10.2249	4.3943***	16.3660**	14.7580**	6.1187***	1.8091***	40.2202**	§	3.6914*	§
	[0.7690; 29.8313]	[3.5701; 4.6632]	[13.9281; 21.9305]	[10.0888; 19.5993]	[1.6986; 6.2218]	[1.7795; 1.8091]	[39.2797; 40.2202]		[2.7451; 19.0181]	
γ_2	47.0530**	16.3243**	112.2358*	43.5636***	34.4125***	10.5331**	114.8387**	§	18.4681	§
	[45.6733; 47.6036]	[7.8313; 17.0944]	[105.0486; 112.7820]	[43.5636; 43.6545]	[31.2791; 34.7213]	[5.5490; 16.7051]	[73.1232; 118.6542]		[14.8694; 33.9366]	
Diagnostics										
R^2	0.3136	0.2543	0.0896	0.2743	0.8190	0.3153	0.1639	0.0634	0.2286	0.3336
Het Test	11.3892	170.4870***	42.3277***	149.6563***	305.4955***	61.5894***	242.5125***	399.5303***	3.4143	76.9485***

Notes: estimation carried out in Matlab R2011b; Eicker (1967)-Huber (1967)-White (1980) heteroschedasticity robust standard errors in parentheses in case the White test for heteroschedasticity is statistically significant at conventional levels; *** denotes statistical significance at the 1% level, ** at the 5% level, * at the 10% level based on two tailed tests; ¥ estimates are based on results from the refinement estimators (14) and (16) when two thresholds are statistically significant, 95% confidence intervals in square brackets.

Source: authors' estimation on EU-FADN - DG AGRI data employing a modified version of the Matlab code provided by Hansen on his website.

FACTOR MARKETS

Comparative Analysis of Factor Markets for Agriculture across the Member States

245123-FP7-KBBE-2009-3

The Factor Markets project in a nutshell

Title	Comparative Analysis of Factor Markets for Agriculture across the Member States
Funding scheme	Collaborative Project (CP) / Small or medium scale focused research project
Coordinator	CEPS, Prof. Johan F.M. Swinnen
Duration	01/09/2010 – 31/08/2013 (36 months)
Short description	Well functioning factor markets are a crucial condition for the competitiveness and growth of agriculture and for rural development. At the same time, the functioning of the factor markets themselves are influenced by changes in agriculture and the rural economy, and in EU policies. Member state regulations and institutions affecting land, labour, and capital markets may cause important heterogeneity in the factor markets, which may have important effects on the functioning of the factor markets and on the interactions between factor markets and EU policies. The general objective of the FACTOR MARKETS project is to analyse the functioning of factor markets for agriculture in the EU-27, including the Candidate Countries. The FACTOR MARKETS project will compare the different markets, their institutional framework and their impact on agricultural development and structural change, as well as their impact on rural economies, for the Member States, Candidate Countries and the EU as a whole. The FACTOR MARKETS project will focus on capital, labour and land markets. The results of this study will contribute to a better understanding of the fundamental economic factors affecting EU agriculture, thus allowing better targeting of policies to improve the competitiveness of the sector.
Contact e-mail	info@factormarkets.eu
Website	www.factormarkets.eu
Partners	17 (13 countries)
EU funding	1,979,023 €
EC Scientific officer	Dr. Hans-Jörg Lutzeyer

SEVENTH FRAMEWORK PROGRAMME