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# Does the Common Agricultural Policy Reduce Farm Labour Migration?

## Panel data analysis across EU regions

**ABSTRACT**

This paper deals with the determinants of labour out-migration from agriculture across 149 EU regions over the 1990–2008 period. The central aim is to shed light on the role played by payments from the common agricultural policy (CAP) on this important adjustment process. Using static and dynamic panel data estimators, we show that standard neoclassical drivers, like relative income and the relative labour share, represent significant determinants of the intersectoral migration of agricultural labour. Overall, CAP payments contributed significantly to job creation in agriculture, although the magnitude of the economic effect was rather moderate. We also find that pillar I subsidies exerted an effect approximately two times greater than that of pillar II payments.

**JEL codes**: Q12, Q18, O13, J21, J43, J60.

**Keywords**: Out-farm Migration, Labour Markets, CAP Payments, Panel Data Analysis



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### Alessandro Olper, Valentina Raimondi, Daniele Cavicchioli and Mauro Vigani*



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## 1. Introduction

Over the last 50 years, EU countries have experienced dramatic adjustments in the agricultural labour market, showing an impressive out-farm migration of the labour force. Surprisingly, in the most recent decades, we do not find any substantial reduction in the migration rate, a stylised fact that is at odds with €50 billion per year of income subsidies spent by the common agricultural policy (CAP).[1]

Indeed, the empirical evidence on the effect of agricultural subsidies on out-farm migration is quite inconclusive. Among the literature, there are papers that find a negative impact of subsidies on out-farm migration (e.g. Breustedt & Glauben, 2007; D'Antoni & Mishra, 2010), papers that find no effect (e.g. Barkley, 1990; Glauben et al., 2006) and even papers that find a positive effect of subsidies on farm out-migration (e.g. Dewbre & Mishra, 2007; Petrick & Zier, 2011).

One interpretation of this counterintuitive pattern is that the agricultural subsidies have been ineffective as an income support policy,[2] especially because of the imperfections in both input and output markets. Different channels have been emphasised through which farm subsidies may affect agricultural employment, accounting for the mixed evidence summarised above. For example, Barkley (1990) stressed that an indirect impact of subsidies may have occurred through increased land values.[3] Land value appreciation slows down the rate of labour migration out of agriculture. By contrast, Goetz & Debertin (1996) argue that capital–labour substitution effects may represent a driver of the positive correlation between farm subsidies and out-farm migration. More recently, Berlinschi et al. (2011) found evidence of another


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[1] The rate of out-farm migration in the EU-15 has been equal to about 2.5-3% per annum over the last two decades.

[2] For example, an important OECD (2001) study emphasised that only 20% of all agricultural support policies resulted in the net growth of farm income in OECD countries, with the bulk of the aid being dissipated elsewhere, for example to the owners of production factors.

[3] Yet while studies from the US show that landowners capture a substantial share of subsidies (e.g. Goodwin et al., 2005; Kirwan, 2009; Lence & Mishra, 2003), recent evidence from the EU shows that CAP subsidies are only marginally capitalised into land values (see Ciaian et al., 2011; Michalek et al., 2011; Ciaian & Kancs, 2012).

indirect channel, the effect of subsidies on the educational level of farmers' children and the resulting impact on long-term labour supply.

These and other indirect effects of farm subsidies clearly deserve attention to better understand the key mechanisms responsible for the puzzling effect on agricultural labour summarised above. In this paper, however, we argue that when the direct effect of farm subsidies is properly estimated, at least in the context of the EU regions investigated here, we consistently find that on average the farm support programme has had a negative effect on out-farm labour participation. Thus, the CAP contributes to job creation in agriculture.

The creation and maintenance of jobs in agriculture and in rural areas has been a traditional CAP objective – one recently re-stated and underlined in several EU official documents (e.g. European Commission, 2010a; European Parliament, 2010).[4] Yet, especially owing to data limitations, evidence concerning the effect of CAP subsidies on off-farm labour migration has been rather inconclusive. Moreover, it is mostly confined to specific country or regional case studies, only rarely providing a European-wide perspective (Shucksmith et al., 2005; Petrick & Zier, 2011). Therefore, although interesting and often rich in detailed interpretations, such studies measure the CAP effects only within a single country or region, an approach that has the advantage of keeping factors like institutions fixed and circumventing problems associated with cross-country/regional analyses. Still, one of the shortcomings of these studies is that the findings are difficult to generalise to other countries and regions where there are wide differences in development, labour market institutions and farming structures. Until now the lack of comparable and consistent estimates of CAP payments at the EU regional level has prevented the adoption of an approach that takes into account both cross-country and cross-regional observable and unobservable heterogeneities.[5]

The main objective of this paper is to offer a preliminary contribution that moves in that direction. Specifically, the paper investigates the effect of CAP payments on intersectoral labour reallocation, extending earlier studies in three main directions. First, our analysis has broad coverage, considering 149 EU regions over the period from 1990 to 2008. Second, the effects of CAP instruments are analysed, concentrating on both pillar I payments (coupled and decoupled subsidies) and on several pillar II instruments for rural development. Notably, with the exception of Petrick & Zier (2011), who studied the entire portfolio of CAP measures, previous analyses have normally considered only one instrument at a time or an aggregate of various policies. This approach is problematic, however, as different policies can have different effects on job creation in agriculture. Third, we rely on modern panel data methods, estimating both static and dynamic migration equations in order to account for several identification issues, such as unobserved heterogeneity, dynamics and endogeneity. Finally, we undertake a back-of-the-envelope calculation of the net benefits of the CAP in terms of farm job creation.

The remainder of the paper is organised as follows. The next section provides a short review of the empirical literature to date. Section 3 presents our conceptual framework and the empirical strategy for investigating the effects of the CAP on labour migration. Section 4 describes the data and how we measure the CAP payments at the EU regional level. In section 5, the results are presented and discussed. Finally, section 6 concludes.

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[4] The European Commission's (2010b) reflection about the future of the CAP (Communication on the CAP towards 2020, COM(2010) 672) explicitly addressed agricultural and rural labour issues in several sections of the document. Labour and rural employment issues are also well represented in the recent European Parliament (2010) document on CAP reforms, *On the Future of the CAP after 2013* (EP 439.972).

[5] A notable exception is the paper by Esposti (2007), who investigated the effect of CAP pillar I payments on economic growth and convergence across EU regions over the 1989–2000 period.

## 2. Previous evidence

In this section we summarise the literature on the effect of agricultural and rural subsidies on the labour market. Theoretically, the studies can be divided into two main approaches. The first looks at agricultural household models to analyse the impact of subsidies on the allocation of household labour (Lee, 1965; Becker, 1965).[6] In this framework, subsidies directed at farm income support may affect farmers' decisions on labour allocation in a number of ways: increasing the marginal value of farm labour, increasing household wealth and reducing income variability. While a coupled payment increases the marginal value of farm work, decoupled payments are considered a source of non-labour income.

The other approach focuses on the change in labour markets resulting from the entry and exit processes from one sector to another. The decision to exit or enter farming is normally analysed using models of occupational choice that have their roots in the Todaro (1969) and Harris & Todaro (1970) two-sector model. In this framework, the choice of occupation is determined by comparing the discounted utility derived from each alternative job over the career of the individual, taking into account the net costs of changing occupation and the probability of obtaining a job in the other sector (see Mundlak, 1979).

The above distinction is also reflected in empirical works, with studies at the farm-household level largely based on microdata at the farm level, and studies on farm labour (re)allocation conducted at the aggregate (country or regional) level. Microdata allow us to address the individual adjustment behaviour in response to changes in factors affecting the household utility, such as alternative revenue sources. For example, Mishra & Goodwin (1997), using a Tobit model on farm households located in Kansas, found that policy changes reducing farm income support can increase the off-farm employment of farmers and their spouses. Similarly, El-Osta et al. (2004) investigated the effect of payments under the US Agricultural Market Transition Act on agricultural labour supply using 2001 data. Results indicate that government payments tend to increase the hours that operators work on-farm and vice versa. There are also important examples of microdata analysis using panel data (e.g. Pietola et al., 2003; Gullstrand & Tezic, 2008) and semi-parametric approaches (e.g. Pufahl & Weiss, 2009; Esposti, 2011a).[7]

Nevertheless, available farm-level data are often time constrained. This can impede an in-depth analysis of general economic conditions and agricultural policy, as these factors concern all the farmers in a specific region, and any existing time dimension of the studies is typically very short (Glauben et al., 2006). Moreover, it is not always clear how much the results from farm-level studies, mostly based on survey data, are representative of the entire population.

The analysis at the aggregate level is, in principle, less data constrained, enhancing panel data methods and providing results with broader coverage. The process of labour migration from one sector to another is assessed by controlling for such structural variables as country or regional relative income, unemployment, population densities, and for institutional and policy variables. Econometric approaches of aggregate studies range from cross-sectional to

[6] In the basic household model, individuals behave in accordance with a well-defined utility function accounting for household production, consumption and leisure. To maximise their utility, farm households choose to allocate time between leisure and on- and off-farm labour.

[7] In recent years, there has been a growing tendency to use semi-parametric approaches, like propensity-score matching, to study the economic effects of EU policy in general (e.g. Becker et al., 2010; 2012) and of CAP subsidies in particular (e.g. Pufahl & Weiss, 2009; Esposti, 2011a; Salvioni & Sciulli, 2011; Ciaian et al., 2011). Generally speaking, these quasi-experimental methods have several advantages with respect to standard regression tools, but they also have some drawbacks. For example, when applied to the CAP pillar I subsidies – a quasi-horizontal measure – finding suitable counterfactuals (controls) tends to be a challenge (see Esposti, 2011a; 2011b).

time-series analyses, and more recently panel data methods and also quasi-experimental approaches.

The seminal work of Barkley (1990) used a two-sector, occupational choice model on a large time series (from 1940 to 1985) to analyse the labour migration out of agriculture in the US, using government payments as a key variable. Results show that the effect of farm support on agricultural labour is negative but insignificant. The author interprets this result by arguing that it might stem from two offsetting effects of different government payments. Indeed, income subsidies, like price support and target price, are expected to reduce the rate of out-farm migration, while other farm policies, like acreage set-asides, inducing land diversion, can reduce the need for inputs that complement the land, resulting in increased out-farm migration. This interpretation, however, is at odds with the findings of D'Antoni & Mishra (2010). They extended Barkley's sample to 2007, also accounting for dynamics, through an autoregressive distributed lag model. By taking dynamics into account, the farm support effect on out-farm labour migration became significantly negative.

At the EU level, many studies have investigated the effect of national support policies (other than CAP payments) at the single country level (e.g. Pietola et al., 2003; Goodwin & Holt, 2002; Benjamin & Kimhi, 2006; Glauben et al., 2006), while only a few studies have investigated the effect of CAP subsidies on out-farm migration. For empirical works at both the household and aggregate level, actual evidence of the direct effect of CAP subsidies on the off-farm labour participation/migration is quite inconclusive. Results are often confined to specific countries or regions (Pufahl & Weiss, 2009; Hennessy & Rehman, 2008; Gullstrand & Tezic, 2008), mainly as a consequence of data limitations at the EU regional level. Most of the authors have used a cross-sectional approach (Breustedt & Glauben, 2007; Hennessy & Rehman, 2008; Van Herck, 2009), while those who have performed a panel data analysis have considered only a single country or specific policy (or both), such as Objective 1 or agri-environmental measures (Gullstrand & Tezic, 2008; Pufahl & Weiss, 2009; Salvioni & Sciulli, 2011).

Only a few studies have worked at the overall EU level. Breustedt & Glauben (2007) investigated the effect of total farm subsidies on out-farm labour migration in 110 EU NUTS 2 regions, finding that CAP payments slowed down structural change in the 1993–97 period. Van Herck (2009) used a multinomial logit approach to investigate the main destination of households exiting the agricultural sector. Coupled, decoupled and total subsidies showed a positive effect on out-farm migration for 144 EU NUTS 2 regions, mainly as a consequence of secondary-order effects. Becker et al. (2010) used a regression-discontinuity design approach to study the total employment effect on 285 EU NUTS 2 regions for the 1989–2006 period. The results showed no significant effect on total employment of the structural funds programme (Objective 1). Finally, Petrick & Zier (2011), using a difference-in-difference (DID) estimator on three East German Länder, found a positive effect of coupled, decoupled and rural development payments respectively (but not agri-environment payments) on out-farm labour migration. Their DID approach represents a relevant improvement, despite the results from three German Länder hardly being extendible to the EU as a whole.

To sum up, the actual evidence concerning the effect of CAP payments on out-farm migration is not only quite inconclusive but it also suffers several drawbacks. First, the evidence comes mostly from cross-sectional inference. Second, it is often derived from country or regional case studies. Third, it rarely takes into account the entire portfolio of CAP payments. Last but not least, no particular effort has been given to accounting for the potential problems of endogeneity bias. Our paper takes advantage of a large sample of 149 European regions observed over 18 years to assess the direct effect of subsidies on out-farm labour migration, overcoming some of the drawbacks of previous literature.

# 3. Conceptual model and empirical strategy

## 3.1 Out-farm migration equation

This paper is empirical in nature. Still, to rationalise our work we sketch the theory of occupational choice and labour migration decisions, which has its roots in the Todaro (1969) and Harris & Todaro (1970) two-sector model, subsequently developed by Mundlak (1979).[8]

Following Barkley (1990), let us consider individuals facing a given return in two mutually exclusive occupations $i$, say agricultural ($i$=1) and non-agricultural employment ($i$=2). The choice of occupation is determined by comparing the discounted utility derived from the job throughout their careers. A worker aged $g$ who retires at time $T$ will face an optimisation problem as described in equation (1), where $r$ is the discount rate:

$$H_{ik} = \int_g^T e^{-rt} V(X_{it}, L_{it})dt - \int_g^T e^{-rt} V[(X_{jt}, L_{jt}) - C_{ijt}]dt \tag{1}$$

with $X_{it} = q_{it}\, w_{it}\, L_{it}$.

Utility in the period $t$ is a function of both consumption ($X_{it}$) and hours of work spent on the job ($L_{it}$). Migration of an individual from one occupation to another occurs when the expected utility derived from a potential profession rises above the utility expected from the current job, net of the costs incurred in changing profession ($C_{ijt}$). We assume that agriculture $i$ is the current occupation and $j$ is some other non-agricultural occupation. Migration from $i$ to $j$ will occur when the net utility is negative ($H_{ik} < 0$).

Although the return to labour may be higher in a non-agricultural occupation than in farming, an agricultural worker involved in job search may discount the higher wage rate ($w_j$) by the probability ($q_j$) of obtaining employment in the non-agricultural sector. For that reason, migration from agriculture to other sectors does not occur instantaneously.[9]

To calculate $H_{ik}$, a potential migrant has to estimate the probability of obtaining a job in the industrial sector. Clearly, this probability is affected by macroeconomic conditions, such as the unemployment rate and the relative size of the sectoral labour forces. Other things being equal, the larger the non-agricultural labour market, the easier it should be to obtain a job there. Yet as most migrations are out of agriculture, migration will also increase with the size of the labour force in agriculture (Larson & Mundlak, 1997). Moreover, economic conditions in the agricultural sector, such as government payments or the structure of the family farm, are also expected to affect the migration rate out of agriculture.

The migration of individual $k$ occurs if $H_{ik} < 0$. As the empirical model considers the regional rates of net out-farm migration, an index function $f_{ik}$ is used to separate migrants from non-migrants. That is, $H_{ik}\, f_{ik} \leq 0$ where $f_{ik} = 1$ if $H_{ik} < 0$ (migration occurs), $f_{ik} = 0$ if $H_{ik} \geq 0$ (migration does not occur). This index function allows for the aggregation of individual migrants by the summation across $f_{ik}$. The gross migration rate $M_{ij}$ from occupation $i$ to occupation $j$ during one period can be written as

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[8] The Harris & Todaro (1970) model is a modification of the original Todaro (1969) model, which adds a two-sector, neoclassical trade model to the analysis. The model uses traditional neoclassical mechanisms and introduces a migration equation, which represents its innovative feature.

[9] Note that the return to labour in this model works as summary statistics, in the sense that structural parameters like the substitutability of capital for labour, the (low) income consumption elasticity of farm products and the productivity growth rate are supposed to affect the migration rate solely through their effect on the relative returns to labour in the farm and non-farm sectors. For a two-sector growth model with a farm–non-farm wage gap that explicitly considers these structural parameters, see Dennis & Iscan (2007).

$$M_{ij} = \sum_{k=1}^{I} f_{ik} \, , \qquad (2)$$

where $I$ is the people employed in occupation $i$.

Because of people flow from one sector to another and vice versa, the net out-farm migration can be defined as $m = M_{ij} - M_{ji}$, where $m$ represents our dependent variable in the empirical model.

### 3.2 Measurement issues

In practice, due to data limitations, migration flows in both directions are not observable. Previous empirical applications measured out-farm migration simply as the growth rate in agricultural employment from one year to the next, disregarding the dynamics in the total labour force (e.g. Barkley, 1990; D'Antoni & Mishra 2010). This approach can be a reasonable approximation when the exercise is conducted within a single country. Working across the EU regions, however, as in the present study, disregarding the differences in the total labour force dynamics at the regional level can introduce a systematic bias in the estimates of intersectoral labour migration.

To reduce this potential source of bias, the approach of Larson & Mundlak (1997) has been followed, assuming that without migration labour in agriculture and non-agriculture would grow at the same rate as the total labour force. Deviation from this rate is attributed to migration. Formally, the net migration rate is estimated using the following relation:

$$m = [L_{1t-1}(1+n) - L_{1t}] / L_{1t-1} \, , \qquad (3)$$

where $n = (L_t - L_{t-1}) / L_{t-1}$ is the growth rate of the total labour force.

It is important to point out that using equation (3) to estimate the out-farm migration of labour is not immune to other potential shortcomings. Indeed, a first drawback lies in the fact that it does not take into account part-time farming, which has become an important characteristic of the EU agricultural labour market. Hence, it potentially leads to a heterogeneous underestimation of the out migration of labour, as part-time farming differs significantly across EU regions. Thus, our empirical strategy has to be robust to this and other forms of regional heterogeneity. A second issue is that to measure the migration rate we should use data on labour. Yet, as better explained in the data section, the disposable regional sources do not report data on agricultural labour, but rather on agricultural employment. This introduces volatility into the series because we are introducing demand shocks in the migration estimates.[10]

### 3.3 Econometric approach

Armed with this simple theoretical logic and following previous works, the rate of out-farm migration $m$ is expected to be primarily a function of the relative per-capita income between non-farm and farm activities ($RI$), and all other factors affecting the costs incurred to change profession ($C$).

Our main goal is to isolate the effect of the CAP on the rate of out-farm migration. Following the model's logic, to the extent that CAP subsidies ($S$) are effective in transferring income to farmers, their effect should decrease the farmers' propensity to migrate to another sector, *ceteris paribus*. Empirically the rate of out-farm migration of the EU region $i$ at time $t$ can be represented by the following benchmark equation:

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[10] Note, however, that previous papers faced the same issue (see Barkley, 1990; Larson & Mundlak, 2003; D'Antoni & Mishra, 2010).

$$m_{it} = \beta_0 + \beta_1 RI_{it-1} + \beta_2 S_{it-1} + Z_{it-1} + \nu_{it}, \quad (4)$$

where $Z$ is a vector including all other observable factors, such as the relative labour share and the unemployment rate, which affect the migration costs, $C$, and $\nu_{it}$ is the error term. If the neoclassical drivers, $RI$ and $S$, have direct and independent effects on the migration rate $m$, then we should expect that $\beta_1 > 0$ and $\beta_2 < 0$, respectively.

The assumption about the error term is critical for our identification hypothesis. Our main concern in estimating equation (4) is omitted variable bias due to factors correlated with our key variables of interest. We assume that the error term $\nu_{it} = \alpha_t + \mu_i + \varepsilon_{it}$ comprises time fixed effects common to all regions $\alpha_t$, time-invariant, regional fixed effects $\mu_i$ and a time-varying component $\varepsilon_{it}$. Therefore, by including time and regional fixed effects, equation (4) is equivalent to a DID estimator. The fixed effects control for both observed and unobserved (regional) heterogeneity, rendering the assumption of exogeneity of our right-hand side variables more credible. This consideration is of vital importance for properly identifying the average effect of the CAP payments on regional out-farm migration. Indeed, the inclusion of fixed effects controls for (time-invariant) observable and unobservable differences in the unit of observations, such as the stock of human capital, the age structure of the farm population or the share of land under property. These are all variables that can affect a farmer's decision to migrate, but which change very slowly over time.

Nevertheless, the inclusion of fixed effects does a good job in resolving endogeneity bias due to regional heterogeneity or selection bias (or both). Hence, our key identification assumption is that the policy variable, $S_{it}$, is not simultaneously determined with the regional rate of out-farm migration, $m_{it}$. Different arguments may justify this assumption. First, because we work at the EU regional level, it appears plausible to assume that pillar I payments are exogenous to migration, given that these policies are decided at the EU level. In principle, this assumption may be more questionable when pillar II payments are considered. In this case, the policy-making process also falls under the responsibility of the EU regional institutions (Petrick & Zier, 2011), which may generate a potential problem of endogeneity bias owing to political economy motives (see Berlinschi et al., 2011). Still, the degree of freedom of regional governments to allocate money from pillar II affects only the equilibrium among different pillar II measures (and the axis), but not their aggregated level. The overall amount of pillar II expenditure is predetermined through a bargaining process at the EU and national levels.[11] Thus, in our basic model we treat the policy variable as exogenous. To be more precise, because it is plausible to assume that the farmer's choice to exit at time $t$ is affected by the level of CAP support at time $t$-1, in equation (4) the term $S$ as well as the other independent variables are always included as lagged by one year, thereby treated as predetermined variables.[12]

A potential concern of using equation (4) is its static nature. D'Antoni & Mishra (2010) for the US as well as Petrick & Zier (2011) for three East German Länder showed that considerations of 'dynamics' may be important in studying the effect of farm subsidies on out-farm migration.[13] To tackle this issue we also estimate a dynamic autoregressive specification:

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[11] Clearly this does not mean that some form of 'compensation rule' between pillar I and pillar II policies cannot work here. But to the extent to which this compensation is decided at the EU level, it should not then affect expenditure at the regional level, *ceteris paribus*.

[12] This approach also reduces the possible simultaneity in the model between farm migration and other right-hand side variables, especially relative income and unemployment.

[13] Specifically, D'Antoni & Mishra (2010) showed that moving from a static to a dynamic autoregressive specification matters for the final results. In a different vein, Petrick & Zier (2011) reported evidence of persistent lag structure for some CAP instruments.

$$m_{it} = \beta_0 + m_{it-1} + \beta_1 RI_{it-1} + \beta_2 S_{it-1} + Z_{it-1} + \mu_i + \alpha_t + \varepsilon_{it} \,. \tag{5}$$

Given the large cross-sections and the short time series of our data set, the correlation between the lagged dependent variable and the transformed error term renders the least squared within estimator inconsistent. To avoid this inconsistency, we use a first difference, generalised method of moments (DIFF-GMM) estimator as an alternative to the within estimator (Arellano & Bond, 1991). A DIFF-GMM estimator transforms the model into a two-step procedure based on first difference to eliminate the fixed effects, as a first step. Next, in a second step, the lagged difference of the dependent variable is instrumented using the lagged differences and levels of the dependent variable.

An important feature of the GMM estimator is the possibility of treating the key variable of interest – namely the CAP subsidies, S – as endogenous, using their lagged values as instruments and testing for their exogeneity. As a robustness check, we also follow this strategy. Because in the presence of variables that display high persistency – like policy variables – the DIFF-GMM estimator could also be biased, suffering weak instrument problems, we also use a system GMM (SYS-GMM) estimator that exploits the second moment condition of the level equation (Arellano & Bover, 1995).

## 4. Data

We start from an initial database of about 160 regions. Some regions are lost, however, due to the lack of data, while others are dropped, resulting in outliers after using specific measures of influence (i.e. DF-Beta). The final sample used for the empirical analysis covers 149 regions of 15 EU countries, over the period 1990–2008.[14] Table 1 shows the number of regions used for each country, according to EU NUTS regions, distinguishing between NUTS 1 and NUTS 2. The choice to utilise both NUTS 1 and NUTS 2 was motivated by the necessity of matching data from different sources. Indeed, the Farm Accountancy Data Network (FADN) regional classification does not always match the NUTS 2 level defined by Eurostat.[15]

[14] DF-Beta is a specific measure of influence that assesses how each coefficient is changed by deleting a specific observation. It measures the influence of each observation on the coefficient of a particular independent variable (i.e. relative labour and relative income). On the basis of this test, two regions – London in the UK and Ovre Norrland in Sweden – were dropped due to their high influence on the coefficients. Note, however, that all the results reported in the paper are robust to the inclusion of these additional regions. These additional results, as well as the DF-Beta tests, can be obtained from the authors upon request.

[15] An alternative solution is to apply the FADN information at NUTS 1 and also to those NUTS 2 regions where FADN data are lacking. Because our empirical strategy especially exploits the within-region variation of out-farm migration and CAP payments, following this approach does not add substantial 'new' information to the model structure.

*Table 1. Sample of country/regions considered*

| Country | NUTS | Number of regions |
|---|---|---|
| Belgium | (2) | 10 |
| Denmark | (2) | 5 |
| Greece | (2) | 11 |
| France | (2) | 22 |
| Germany | (1) | 14 |
| Ireland | (2) | 2 |
| Italy | (2) | 21 |
| The Netherlands | (2) | 12 |
| Austria | (2) | 9 |
| Portugal | (2) | 5 |
| Finland | (2) | 4 |
| Sweden | (2) | 7 |
| Spain | (2) | 17 |
| United Kingdom | (1) | 10 |
| *Total* | | *149* |

*Notes*: Information is missing on the four French overseas departments, the two Portuguese regions Madeira and Azores, the two Greek regions Voreio Aigaio and Notio Aigaio, the Aland region in Finland, Northern Ireland in the UK, the Luxembourg state region and the Bruxelles-Capitale region in Belgium, due to lack of data. The London region in the UK and the Ovre Norrland in Sweden were dropped, being outliers (see text).

## 4.1 Dependent variable

Our dependent variable is the net migration rate, obtained as described in equation (3). In theory, to calculate migration we should use data on labour starting from census data. Unfortunately, such data are available every ten years and can only be transformed into annual series through interpolations. Consequently, because of data limitations, we were forced to use employment data to measure annual migration at the EU regional level. As highlighted by Butzer et al. (2003), these data present two sets of problems: first, they bring the demand for workers into the migration series; second they tend to be more erratic. Nevertheless, the trend still prevails. The basic employment data used to measure the net migration rate comes from Cambridge Econometrics' Regional Database.

## 4.2 Policy data

Given our main objective, how we measure the policy variables at the regional level is a critical issue. Previous studies followed two main approaches. One entailed measuring a regionalised, producer subsidy equivalent, as in Anders et al. (2004), Tarditi & Zanias (2001) and more recently Hansen & Herrmann (2012). Another involved using the FADN, as in Shucksmith et al. (2005), and by combining the same source with Eurostat's Regio New Cronos database, assuring in the former a time variation as well, as in Esposti (2007).

In theory, the last approach is the most suited to our analysis where econometric identification is based on the within-region variation in CAP payments. Unfortunately it has two main shortcomings. First, Eurostat does not provide time series data at the regional level for all EU countries.[16] Second and more importantly, Eurostat data is based on agricultural sector series, and so do not incorporate decoupled subsidies after 2005. Thus, their use

[16] Esposti (2007) resolves this issue by applying the growth rate at the higher aggregation level (NUTS 1) to those regions (NUTS 2) whose Eurostat data are lacking.

would reduce the time coverage of the analysis, and would preclude the possibility of investigating the possible differentiated effect between coupled and decoupled payments, as well as the effect of pillar II subsidies.

To overcome these issues, we adopted a new strategy measuring CAP payments starting from the FADN data at the regional level. For every region covered by the FADN, we have the amount of payments received by the 'average farm' in each year over the period 1990–2008. To the extent that the average farm is representative of the farm population,[17] computation of the ratio between such a farm's CAP payments and the respective farm's net income (inclusive of subsidies) offers the possibility to measure a consistent regional level of farm protection stemming from different policy measures of the CAP.

Note that this approach is fully consistent with previous empirical exercises conducted on US out-farm migration (see Barkley, 1990; D'Antoni & Mishra, 2010), where the effect of government payments is measured using the ratio between farm subsidies to the farm value added at the aggregated (country) level.

A key advantage of our approach is the possibility of disentangling total CAP payments into their distinct components (pillar I and pillar II). Specifically, we can distinguish between the coupled and decoupled payments of pillar I, as well as agri-environmental payments, those for less favoured areas (LFA), investment aids and a residual category called 'other' subsidies of pillar II.[18] Note that some of the latter payments were introduced before Agenda 2000, and hence the 'pillar II' expression would not be fully correct. Nevertheless, we have chosen to use it to clearly and easily distinguish between CAP market subsidies and CAP structural policies.

Finally, a potential limitation of our policy variable is that it does not capture the 'price support' component of CAP transfers – a component that was in place at a decreasing rate until 2003. It is important to note, however, that in our empirical model the price component of CAP protection is implicitly controlled for by the relative income variable, *RI*.

### 4.3 Other covariates

The intersectoral income differential is measured by the ratio of income in non-agriculture to that in agriculture (*RI*). Income is measured as gross value added (GVA) per worker, at constant and basic prices. For the non-agricultural sector we used the difference between total GVA and GVA in agriculture, as well as for non-agricultural employment.[19] The data for GVAs and employment are from Cambridge Econometrics' Regional Database.

The other control variables included in the vectors *Z* are as follows. First, following Larson & Mundlak (1997) and others, we include the relative labour force (*RL*) calculated as the ratio

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[17] For each region, the FADN sample is stratified according to the type of farming (TF) and the economic size unit (ESU) class, while the same stratification is made for the regional farm population. Each stratum in the sample is then weighted to render its data representative of the underlying population. Such a procedure makes the FADN data representative at the regional level for TF and ESU, and indirectly for pillar I payments, while the same may not be said for pillar II payments.

[18] Pillar I includes 'total subsidies on crops', 'total subsidies on livestock' and 'decoupled payments'. Pillar II includes 'total support for rural development' and 'subsidies on investments'. Note, however, that for unknown reasons, in the FADN data the sum of the components of pillar II policies (agri-environmental payments, LFA payments, investment aids and the residual category of 'other' subsidies) is slightly lower than the 'aggregate' pillar II subsidies.

[19] Models of the Harris–Todaro type suggest wages as a measure of (relative) labour returns. Nevertheless, many papers investigating out-farm migration equations have found that more robust results are obtained when relative income or productivity, instead of relative wages, is used. Mundlak (1979) and Larson & Mundlak (1997) justify this finding, arguing that for a long-term decision involving expectations, such as migration out of agriculture, income is thought to be a more informative measure of future prospects than wages, since wages are not the only component of a farmer's income. They also note that measurement problems with wage data are another reason to use relative income rather than relative wages.

of employment in the non-agricultural sector to that in the agricultural sector. Relative labour, on the one hand, captures the absorption capacity of non-agricultural sectors. On the other hand, given the direction of structural change with economic development, having a high level of (relative) agricultural employment means more potential migrants coming out of the farm sector. Therefore, its estimated effect can be either positive or negative. Second, to control for search costs and the probability of finding a job in the non-agricultural sector, we include the overall rate of unemployment and a measure of population density, calculated as the total population over the regional area in $km^2$. This variable might account for several market conditions, in particular product and land markets (Glauben et al., 2006); furthermore, it represents a very rough proxy of the average 'distance' from urban areas. Third, we include a variable that measures the number of family workers. The underlying idea here is that a high number of family members working on the farm should lower the exit rate (Breustedt & Glauben, 2007).

Finally, we also include a variable measuring country differences in labour market institutions, which increases with rigidities in labour entry and exit. Specifically, we use the OECD employment-protection indicator called 'EP_v1' (see OECD, 2010). This index is the average of six different sub-indices of 'regular' and 'temporary' contracts with a scale from 0 (less restrictive) to 6 (most restrictive). The intuition is that higher labour rigidities should increase the costs of off-farm labour migration. A shortcoming of the index is that its time variation is obviously linked to labour market reforms, events that do not occur yearly, inducing a low time variation.

Information on population, regional area, unemployment rate, total and sectoral employment come from Cambridge Econometrics' Regional Database. Information on farm family workers comes from FADN, while the rigidity index of labour institutions is based on OECD data. Summary statistics of the variables explained above are reported in Table 2.

*Table 2. Descriptive statistics*

| Variable | | Mean | Std.Dev. | Min | Max |
|---|---|---|---|---|---|
| Out-farm migration | Growth rate | 0.026 | 0.075 | -0.939 | 0.375 |
| Relative Income | Ratio | 2.114 | 1.461 | 0.475 | 30.92 |
| Relative Labour | Ratio | 34.36 | 53.03 | 1.25 | 605.45 |
| Unemployment rate | % | 8.52 | 5.06 | 1.59 | 36.11 |
| Population density | Persons /Km2 | 263.31 | 513.75 | 3.01 | 4796.32 |
| Family Farm Labor Force | Annual work unit | 1.324 | 0.256 | 0.430 | 2.160 |
| Total payments/VA | Share | 0.374 | 0.316 | 0.000 | 3.097 |
| Pillar I payments/VA | Share | 0.276 | 0.217 | 0.000 | 1.982 |
| Coupled payments/VA | Share | 0.226 | 0.215 | 0.000 | 1.982 |
| Decoupled payments/VA | Share | 0.050 | 0.123 | 0.000 | 0.750 |
| Pillar II payments/VA | Share | 0.098 | 0.144 | 0.000 | 1.172 |

*Source*: See text.

## 5. Econometric results

Table 3 reports the static DID estimate of equation (4). The specifications differ with respect to how the policy variables are considered.[20] Following D'Antoni & Mishra (2010), Augmented Dickey Fuller (ADF) tests were used to determine whether the data were

[20] We test several potential non-linearities (i.e. square terms, threshold effects), focusing especially on the relative income and CAP variables. The data systematically reject these non-linearities, however. These additional results can be obtained from the authors upon request.

stationary.[21] All variables, with the exception of the relative labour and unemployment rates, were found stationary. Thus, these variables were introduced in first difference in the static DID specification.

*Table 3. Out-farm migration and the CAP: Static difference-in-differences results*

| Dependent variable: Out-farm migration | | | | | |
|---|---|---|---|---|---|
| | Difference-in-differences | | | | |
| Variables | (1) | (2) | (3) | (4) | (5) |
| Total payments | -0.0083*** | | | | |
| | (0.0024) | | | | |
| Pillar I payments | | -0.0148*** | | | |
| | | (0.0045) | | | |
| Coupled payments | | | -0.0142*** | | |
| | | | (0.0043) | | |
| Decoupled payments | | | -0.0472*** | | |
| | | | (0.0144) | | |
| Pillar II payments | | | | -0.0127* | |
| | | | | (0.0065) | |
| Agrienvironment | | | | | -0.0225*** |
| | | | | | (0.0083) |
| Less favoured areas | | | | | -0.0428 |
| | | | | | (0.0286) |
| Investment aids | | | | | 0.0291 |
| | | | | | (0.0320) |
| Other pillar II payments | | | | | -0.4274** |
| | | | | | (0.2085) |
| Relative income | 0.0132*** | 0.0133*** | 0.0134*** | 0.0132*** | 0.0129*** |
| | (0.0029) | (0.0029) | (0.0029) | (0.0029) | (0.0030) |
| Relative labour (diff) | 0.0113*** | 0.0113*** | 0.0113*** | 0.0113*** | 0.0113*** |
| | (0.0009) | (0.0009) | (0.0009) | (0.0009) | (0.0009) |
| Unemployment (diff) | 0.1068 | 0.1062 | 0.0937 | 0.1071 | 0.1113 |
| | (0.1188) | (0.1187) | (0.1181) | (0.1191) | (0.1200) |
| Population density | 0.0011 | -0.0030 | -0.0319 | 0.0082 | -0.0111 |
| | (0.1206) | (0.1208) | (0.1208) | (0.1204) | (0.1206) |
| Family work | -0.0108 | -0.0110 | -0.0121 | -0.0104 | -0.0124 |
| | (0.0082) | (0.0082) | (0.0082) | (0.0082) | (0.0084) |
| Labour protection | -0.0047 | -0.0046 | -0.0036 | -0.0051 | -0.0046 |
| | (0.0033) | (0.0033) | (0.0033) | (0.0033) | (0.0035) |
| Decoupling dummy | -0.0018 | -0.0119** | -0.0034 | 0.0058 | 0.0089 |
| | (0.0071) | (0.0054) | (0.0064) | (0.0069) | (0.0070) |
| Constant | -0.0391 | -0.0143 | 0.0944 | -0.0654 | 0.0116 |
| | (0.4577) | (0.4581) | (0.4583) | (0.4571) | (0.4578) |
| No. Groups | 149 | 149 | 149 | 149 | 149 |
| No. of obs. | 2548 | 2548 | 2548 | 2548 | 2548 |
| R-Sq | 0.53 | 0.53 | 0.53 | 0.53 | 0.53 |
| Adj. R-Sq | 0.49 | 0.49 | 0.49 | 0.49 | 0.49 |

*, ** and *** indicate statistical significance at the 10%, 5% and 1% levels, respectively.

*Notes*: Region and year fixed effects are included in each regression. Robust standard errors are clustered by region in parentheses.

[21] Specifically, given the unbalanced panel structure of our dataset, use was made of the Maddala & Wu (1999) ADF test for unbalanced panel data.

In line with the labour migration model, the relative income between the non-farm and farm sector exerts a positive and significant effect on the level of out-farm migration ($p$-value < 0.01). The estimated elasticity is around 1.1, therefore smaller than that estimated by Barkley (1990) for the US, and equal to 4.5. Notably, however, it lies in the same order of magnitude. Our lower estimated elasticity suggests that at the EU regional level, out-farm migration is less responsive to income differences. The changes in the relative labour force were also significant and positive. Thus, when there is a positive difference in the labour force ratio from one period to the next, farm labourers can increasingly be absorbed into the non-agricultural sector, resulting in greater migration of labour from agriculture, a result close to D'Antoni & Mishra's (2010) findings.

Concerning the other covariates, family workers and the restrictiveness of labour protection institutions have the expected negative sign. By contrast, the effects of changes in the unemployment rate and population density were (often) unexpectedly positive. Even so, all these additional controls are always insignificantly different from zero in this specification.

Moving to the CAP effects, column 1 of Table 3 considers the total level of CAP payments (pillar I plus pillar II). Its estimated coefficient is negative and significant at the 1% level. Overall, the CAP played a role in keeping labour in agriculture, *ceteris paribus*. This result confirms the finding by D'Antoni & Mishra (2010) for the US economy, but it goes in the opposite direction of that by Petrick & Zier (2011), who showed that excluding agri-environmental payments, CAP subsidies significantly increase out-farm migration in three East German Länder.

The subsequent regressions of Table 3 display results considering the CAP policy instruments separately. Note that we have been forced to conduct the analysis of pillar I and pillar II policies in isolation, as the two series are strongly collinear.[22] Taking first the pillar I payments, the estimated policy coefficient is again negative and strongly significant, both in isolation (column 2) and when the effect between coupled and decoupled subsidies is split (column 3).

Columns 4 and 5 show results for pillar II policies. Also this group of measures, taken as a whole, points to a significant, negative out-farm migration effect, although only at the 10% level. Still, this effect is heterogeneous across instruments. Splitting pillar II policies, we find that money directed at agri-environmental measures and the category of 'other pillar II' payments contributes significantly to job creation in agriculture. LFA payments exert an effect in the same direction, while investment aids, consistent with expectations, display a positive effect on out-farm migration, although both are statistically insignificant. Broadly speaking, the results of pillar II measures are more in line with the findings of Petrick & Zier (2011). Finally, in the DID specification we do not find any clear evidence of a policy shock due to the 2003 Fischler reform. Indeed, a dummy equal to 1, from 2005 onwards (0 otherwise) switches from negative to positive and is often insignificant.

Next, Table 4 introduces dynamics into the specification, by estimating an autoregressive model using the DIFF-GMM estimator. This strategy should shed further light on the robustness of our findings. First, the bottom of Table 4 reports standard tests to check for the consistency of the GMM estimator (see Roodman, 2009). The Arellano-Bond test for autocorrelation indicates that second-order correlation is not present. On the contrary, the presence of first-order serial correlation suggests that the OLS estimator is inconsistent. Moreover, the standard Hansen test confirms that in all cases our set of instruments is valid.

[22] The correlation coefficient between the two series is actually quite high, and equal to 0.50. Note, however, that by including pillar I and pillar II subsidies separately the resulting estimated coefficients could be biased owing to an omitted variable problem. To attenuate this bias, in the next section we also run regressions treating the CAP policy variables as endogenous.

*Table 4. Out-farm migration and the CAP: Dynamic GMM differences results*

| Dependent variable: Out-farm migration | | | | | |
|---|---|---|---|---|---|
| | DIFF-GMM | | | | |
| Variables | (1) | (2) | (3) | (4) | (5) |
| Lagged migration | -0.0803** | -0.0803** | -0.0802** | -0.0802** | -0.0817** |
| | (0.0367) | (0.0366) | (0.0367) | (0.0368) | (0.0367) |
| Total payments | -0.0094*** | | | | |
| | (0.0027) | | | | |
| Pillar I payments | | -0.0123*** | | | |
| | | (0.0041) | | | |
| Coupled payments | | | -0.0120*** | | |
| | | | (0.0040) | | |
| Decoupled payments | | | -0.0319** | | |
| | | | (0.0146) | | |
| Pillar II payments | | | | -0.0212*** | |
| | | | | (0.0077) | |
| Agrienvironment | | | | | -0.0316** |
| | | | | | (0.0155) |
| Less favoured areas | | | | | -0.0315 |
| | | | | | (0.0339) |
| Investment aids | | | | | 0.0442 |
| | | | | | (0.0308) |
| Other pillar II payments | | | | | -0.2597* |
| | | | | | (0.1396) |
| Relative income | 0.0150*** | 0.0151*** | 0.0152*** | 0.0148*** | 0.0146*** |
| | (0.0042) | (0.0042) | (0.0042) | (0.0042) | (0.0044) |
| Relative labour | -0.0028*** | -0.0028*** | -0.0028*** | -0.0028*** | -0.0028*** |
| | (0.0006) | (0.0006) | (0.0006) | (0.0006) | (0.0006) |
| Unemployment | -0.1900** | -0.1891** | -0.1791** | -0.1914** | -0.1963** |
| | (0.0832) | (0.0826) | (0.0827) | (0.0840) | (0.0769) |
| Population density | 0.7334*** | 0.7316*** | 0.7018*** | 0.7415*** | 0.7267*** |
| | (0.1902) | (0.1900) | (0.1847) | (0.1925) | (0.1860) |
| Family work | 0.0066 | 0.0068 | 0.0057 | 0.0067 | 0.0055 |
| | (0.0130) | (0.0130) | (0.0132) | (0.0131) | (0.0131) |
| Labour protection | -0.0043 | -0.0044 | -0.0038 | -0.0045 | -0.0043 |
| | (0.0051) | (0.0051) | (0.0051) | (0.0051) | (0.0053) |
| Decoupling dummy | -0.0163*** | -0.0163*** | -0.0104 | -0.0165*** | -0.0153*** |
| | (0.0055) | (0.0055) | (0.0063) | (0.0055) | (0.0055) |
| No. Instruments | 33 | 33 | 34 | 33 | 36 |
| No. Groups | 149 | 149 | 149 | 149 | 149 |
| No. of obs. | 2360 | 2360 | 2360 | 2360 | 2360 |
| Sargan | 0.34 | 0.36 | 0.37 | 0.32 | 0.34 |
| Hansen test (p-value) | 0.63 | 0.65 | 0.64 | 0.62 | 0.65 |
| AR1 test (p-value) | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| AR2 test (p-value) | 0.70 | 0.71 | 0.71 | 0.70 | 0.74 |

*, ** and *** indicate statistical significance at the 10%, 5% and 1% levels, respectively.

*Notes*: Year fixed effects are included in each regression. Windmeijer-corrected, cluster-robust standard errors are in parentheses. The DIFF-GMM estimator is implemented in STATA using the xtabond2 routine with option laglimits (9), a two-step procedure and orthogonal deviation equation.

The autocorrelation coefficient is significant and negative, although its magnitude is quite low, and around −0.08. A negative autocorrelation coefficient means that if the migration rate at the time *t*-1 is high, then it will be slightly lower at time *t*, a result consistent with the adjustment process under study.

The results of the DIFF-GMM estimator present some important differences. First, the relative labour ratio now negatively affects the migration rate and is strongly significant. Although the specification here is different from the static model (where the entered variable addresses non-stationarity), and therefore not fully comparable, this change in sign is surprising. At the same time, it is consistent with the idea that the larger the labour force in agriculture relative to the non-agricultural sector, the more out-farm migrants can be expected; in other words, regions tend to converge on a similar level of the relative labour ratio. Second, and in line with the *a priori* expectation, in the dynamic model the unemployment rate negatively affects the rate of out-farm migration and is significant at the 5% level. Third, consistent with the intuition, the population density is now positive and strongly significant. Finally, the Fischler reform dummy for the introduction of decoupling (equal to 1 from 2005 onwards) is now negative and significant in all the specifications, but not in the one where pillar I payments are split into coupled and decoupled subsidies (see column 3).” This result is puzzling. On the one hand it seems to suggest that the Fischler reform induced a policy shock that *decreased* the likelihood of the farmer's decision to exit the agricultural sector. On the other hand, as we better discuss in the next section, the migration elasticity of decoupled subsidies is significantly lower in absolute value than that of coupled payments. A possible explanation for this counterintuitive result could lie in the commodities price spike of 2007 and 2008. Indeed, commodity prices had already started to rise slowly in 2005–06, thus at least partially overlapping with the Fischler reform effect.

Moving to policy variables, their estimated coefficients are always negative and significant, giving broad confirmation of the DID results. Considering first the pillar I policies (columns 2 and 3), the magnitude of the estimated coefficients is slightly lower (in absolute value) than the corresponding static estimates reported in Table 3. This result is not surprising as their coefficients capture short-run effects. Yet the picture changes somewhat when the policy variables considered are those related to pillar II. In this case (see columns 4 and 5) the absolute magnitude of the coefficients often shows a slight increase on passing from the static to the dynamic specification, although the pattern of the effects remains quite similar.

## 5.1 Robustness checks

Table 5 reports some robustness checks. For practical reasons let us focus only on total CAP payments (*panel a*), pillar I payments (*panel b*) and pillar II payments (*panel c*).[23] Columns 1 and 2 display results from an OLS and a least square with dummy variables (LSDV) estimator applied to equation (5), with a specification identical to the regressions of Table 4. The CAP effects are still negative and always significant, with the exclusion of pillar II payments in the OLS specification.

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[23] This is because running GMM regressions with many left-hand side variables treated as endogenous induce the well-known problem of instrument proliferation, rendering identification of the CAP effects problematic.

*Table 5. Out-farm migration and the CAP: Robustness checks*

| Dependent variable: Out-farm migration | | | | | |
|---|---|---|---|---|---|
| | OLS | LSDV | DIFF-GMM | | SYS-GMM |
| | | | Exogenous | Endogenous | Endogenous |
| | (1) | (2) | (3) | (4) | (5) |
| *Panel (a) Total payments* | | | | | |
| Lagged migration | -0.0668** | -0.0904*** | -0.0803** | -0.0767** | -0.0847* |
| | (0.0275) | (0.0259) | (0.0367) | (0.0385) | (0.0430) |
| Total payments | -0.0068** | -0.0084** | -0.0094*** | -0.0122*** | -0.0132*** |
| | (0.0028) | (0.0034) | (0.0027) | (0.0035) | (0.0047) |
| No. of obs. | 2425 | 2425 | 2360 | 2360 | 2512 |
| No. Groups | 149 | 149 | 149 | 149 | 149 |
| No. Instruments | | | 33 | 38 | 41 |
| R-Sq | 0.490 | 0.540 | | | |
| Sargan | | | 0.340 | 0.480 | 0.200 |
| Hansen test (p-value) | | | 0.630 | 0.630 | 0.160 |
| Diff-in-Hansen test | | | | | 0.383 |
| AR1 test (p-value) | | | 0.000 | 0.000 | 0.000 |
| AR2 test (p-value) | | | 0.700 | 0.670 | 0.780 |
| *Panel (b) Pillar I payments* | | | | | |
| Lagged migration | -0.0675** | -0.0906*** | -0.0803** | -0.0721* | -0.0782* |
| | (0.0275) | (0.0258) | (0.0366) | (0.0383) | (0.0424) |
| Pillar I | -0.0136*** | -0.0150** | -0.0123*** | -0.0187*** | -0.0166** |
| | (0.0044) | (0.0061) | (0.0041) | (0.0067) | (0.0074) |
| No. of obs. | 2425 | 2425 | 2360 | 2360 | 2512 |
| No. Groups | 149 | 149 | 149 | 149 | 149 |
| No. Instruments | | | 33 | 38 | 41 |
| R-Sq | 0.490 | 0.540 | | | |
| Sargan | | | 0.360 | 0.580 | 0.300 |
| Hansen test (p-value) | | | 0.650 | 0.820 | 0.370 |
| Diff-in-Hansen test | | | | | 0.345 |
| AR1 test (p-value) | | | 0.000 | 0.000 | 0.000 |
| AR2 test (p-value) | | | 0.710 | 0.590 | 0.670 |
| *Panel (c) Pillar II payments* | | | | | |
| Lagged migration | -0.0670** | -0.0905*** | -0.0802** | -0.0691* | -0.0844* |
| | (0.0275) | (0.0259) | (0.0368) | (0.0407) | (0.0448) |
| Pillar II | -0.0044 | -0.0134* | -0.0212*** | -0.0282*** | -0.0318*** |
| | (0.0078) | (0.0076) | (0.0077) | (0.0063) | (0.0092) |
| No. of obs. | 2425 | 2425 | 2360 | 2360 | 2512 |
| No. Groups | 149 | 149 | 149 | 149 | 149 |
| No. Instruments | | | 33 | 38 | 41 |
| R-Sq | 0.490 | 0.540 | | | |
| Sargan | | | 0.320 | 0.230 | 0.110 |
| Hansen test (p-value) | | | 0.620 | 0.070 | 0.020 |
| Diff-in-Hansen test | | | | | 0.149 |
| AR1 test (p-value) | | | 0.000 | 0.000 | 0.000 |
| AR2 test (p-value) | | | 0.700 | 0.610 | 0.800 |

*, ** and *** indicate statistical significance at the 10%, 5% and 1% levels, respectively.

*Notes*: Year fixed effects are included in each regression. Windmeijer-corrected, cluster-robust standard errors are in parentheses. See the notes of Table 4 for details on the GMM estimator. In the GMM regressions of columns 4 and 5, the CAP policy variable is treated as endogenous, and instrumented by its lagged levels and differences.

As is well known, these two additional estimators suffer from dynamic panel bias, the lagged dependent variable being biased upwards and downwards, respectively. Good estimates of the true parameter should therefore lie in the range between the OLS and LSDV values – or at least near it given that these numbers are themselves point estimates with associated confidence intervals (Roodman, 2009). As Bond (2002) pointed out, this provides a useful check on results from theoretically superior estimators. As can be seen, all the GMM regressions reported in the table display a magnitude of the autocorrelation coefficient that systematically falls within the OLS and LSDV range, giving further credence to the robustness of GMM results.

Column 4 exploits one of the key properties of the difference GMM estimator, treating the CAP variable as endogenous and instrumenting it with its lagged values. For all the CAP payments considered, the estimated effect is still negative and strongly significant, also showing a slight increase in their (absolute) magnitude (compare columns 3 and 4). Again, all the specification tests reported at the bottom of each panel indicate well-specified models. Hence, our results are robust to possible endogeneity bias stemming from political economy motives or measurement errors in the CAP variables (or both). Moreover, note that by treating CAP payments as endogenous, we indirectly control for the potential, omitted variable bias induced by treating pillar I and pillar II payments separately in the regressions.[24] We return to this point later.

Finally, column 5 reports a further robustness check by running a system GMM regression. This estimator, exploiting the additional orthogonality condition of the level equation, should work better in the presence of strong persistency in dependent variables and in any other explanatory variable not treated as strictly exogenous. This is because lagged levels of the dependent (explanatory) variable tend to be weak instruments for actual first differences (see Arellano & Bover, 1995; Blundell & Bond, 1998). In our specific situation, we are especially worried about persistence in CAP subsidies when this variable is treated as endogenous. Nevertheless, when running the system GMM regression the results are qualitatively and quantitatively very close.

## 5.2 Discussion

Consistent comparison of the job creation effects of different CAP policies can be made on the basis of their respective elasticities (see Table 6). Several interesting patterns emerge. First, a 1% increase in total CAP payments decreases out-farm migration by about 0.117% when the effect is estimated using the static DID estimator, a value that rises to 0.144% when dynamics are accounted for and to 0.187% when CAP subsidies are treated as endogenous. Thus, the magnitude of the overall economic effect is rather moderate, but it increases when dynamics and endogeneity are taken into account.

[24] More generally, the increase in the absolute magnitude of the estimated effect when CAP payments are treated as endogenous is consistent with different forms of endogeneity bias. Indeed, if political economy motives are at work, then politicians tend to increase CAP payments in response to an increase in out-farm migration, inducing a positive correlation between these two variables and a bias towards zero of the (negative) effect of subsidies on migration. Similarly, measurement error in an explanatory variable suffers the well-known attenuation bias problem, which induces a bias towards zero of its estimated effect. Finally, if the endogeneity bias is the result of running regressions separately for pillar I and pillar II policies, then it could also be the case that this omitted variable problem can translate into a (absolute) downward bias in the estimated effect of the policy variables.

*Table 6. Out-farm migration elasticity to CAP payments*

| | DID | Difference GMM | | | |
|---|---|---|---|---|---|
| | | Long-run | | Short-run | |
| | | Exogen. | Endogen. | Exogen. | Endogen. |
| Total payments | -0.117 | -0.144 | -0.187 | -0.133 | -0.172 |
| Pillar I payments | -0.157 | -0.142 | -0.214 | -0.131 | -0.199 |
| Coupled payments | -0.123 | -0.113 | _ | -0.104 | _ |
| Decoupled payments | -0.091 | -0.067 | _ | -0.061 | _ |
| Pillar II payments | -0.045 | -0.083 | -0.108 | -0.076 | -0.101 |
| Agrienvironment | -0.033 | -0.051 | _ | -0.047 | _ |
| Less favoured areas | -0.030 | -0.024 | | -0.022 | |
| Investment | 0.019 | 0.031 | _ | 0.029 | _ |
| Other pillar II payments | -0.017 | -0.012 | | -0.011 | |

*Notes*: This table reports sample mean elasticities of CAP policy variables based on difference-in-difference, and the DIFF-GMM regression results of Tables 3, 4 and 5, respectively.

This average effect cancels out relevant differences across CAP instruments. The long-run elasticity of pillar I payments, equal to about 0.2% when dynamics and endogeneity are considered (see columns 3 and 5), is actually two times higher in absolute magnitude than the elasticity of pillar II policies. In pillar I, the coupled payments display higher absolute elasticity than decoupled payments, while across pillar II instruments, agri-environmental payments display the higher absolute elasticity to out-farm migration. Note, however, that the last results should be treated with caution, as they are obtained without taking into account the endogeneity of CAP payments, due to the difficulty of running GMM regressions with many instruments.

Interestingly, using the magnitude of the estimated elasticities of pillar I and pillar II payments can shed some light on the source of bias of the different estimators. Consider for example the estimated elasticities based on the DIFF-GMM reported in column 2 of Table 6. The weighted sum of pillar I and pillar II elasticities is significantly lower than the estimated elasticity of total payments.[25] If we exclude aggregation bias, one possibility is that this bias is the result of estimating the effect of pillar I and pillar II policies in isolation. If this is the case, then by treating the CAP payments as endogenous, this bias should be attenuated or eliminated. This is exactly what we find in the data. Indeed, the weighted sum of pillar I and pillar II elasticities from columns 3 and 5 give precisely the estimated elasticities of the total payments reported in the respective first row. This result gives some support to the idea that the endogeneity of CAP payments is, if anything, due to omitted variable bias more than to a simultaneity problem.

Finally, with our estimates at hand and based on a back-of-the-envelope calculation, we may quantify the job creation effect of the CAP. According to the parameter estimates from column 1 of Table 4, a marginal increase in the explanatory variable 'total payments' makes the dependent variable decrease by 0.0094 points. Using the average value across the panel (that is 0.374, see Table 2) and multiplying for the parameter estimates (–0.0094) we obtain –0.00352, that is the reducing average effect of CAP subsidies in terms of out-farm migration. Multiplying such a value for the average stock of agricultural workers (6.897 million/year), we can obtain a rough estimate of the flow of out-farm migration prevented by CAP payments, which is 24,247 agricultural workers per year. To render such a value in percentage terms, consider that without subsidies, the annual out-farm migration rate would

[25] Starting from the sample share of pillar I and pillar II payments, equal to 75% and 25% respectively, we can measure the resulting total payment elasticity as [(–0.142*0.75)+(–0.083*0.25)]= –0.127. This number is actually lower than the estimated total payment elasticity of –0.144.

increase from the actual rate of 0.0260 to 0.0295. The effect of CAP payments, then, reduces the rate of farm labour migration by around 11.9% – not an irrelevant number.[26] This back-of-the-envelope calculation is based on the point estimate of the benchmark specification in column 1 of Table 4. Thus, taking into account the confidence interval around that point estimate, the percentage reduction of the out-farm migration rate attributable to the CAP is still positive, ranging from a minimum of about 5% to maximum 17%. A conservative view is to interpret the back-of-the-envelope calculation as saying that the CAP subsidies might generate a reduction of out-farm migration, although the effect can be rather moderate.

## 6. Conclusions

Understanding the effects of CAP policies is important, as a deeper comprehension of their incidence would allow the design of better policies. This paper contributes in this direction by studying how different CAP instruments affected job creation in agriculture across 149 EU regions over the 1990–2008 period. In the neoclassical two-sector model, intersectoral labour migration is affected by income differences across sectors, *ceteris paribus*. Therefore, to the extent that CAP policies have been effective in transferring income to farmers, they should have contributed to a reduction in the rate of out-farm migration.

This paper has attempted to test these predictions by exploiting the within-region and cross-regional variations in out-farm migration and CAP policies. Using both static and dynamic panel data methods, allowing also for the possible endogeneity of the CAP, we find robust evidence that the CAP has played a role in keeping labour forces in agriculture, although the overall effect is rather small. Among CAP instruments, we show that so far pillar I payments are the most effective policy in reducing out-farm migration, with coupled subsidies showing an elasticity to out-farm migration significantly higher than decoupled ones. The effect of pillar II payments on job creation is significantly lower than that of pillar I payments, and conditional on the instruments considered.

With regard to the other conditioning variables, the results give broad confirmation that relative income is an important determinant of the decision to migrate from agriculture. Still, its elasticity to out-farm migration is quite low when compared with similar studies conducted in other countries, such as the US. This suggests that at the EU regional level, out-farm migration is less responsive to income differences, or put differently, other important forces are at work in affecting farmers' decisions to migrate. Moreover, we also find notable effects on the migration decision of standard structural variables, such as relative labour, the unemployment rate and population density – all factors that affect migration costs.

Our results confirm that the use of a dynamic panel specification is appropriate in this kind of exercise, and also that irrespective of the specification and estimator used, the CAP payments systematically exert a negative effect on the rate of farm labour migration. Thus, the comparison of these results with previous studies on the impact of EU policies on the labour market reveals the criticality of how the policy effect is measured and identified in the empirical model.

An interesting implication of the study, which comes from the structure of the conceptual model, is related to the 'efficiency' of CAP payments in transferring income to farmers. Indeed, although several previous works have documented an overall inefficiency of (coupled) agricultural payments (e.g. OECD, 2001), our results seem to contradict this conclusion, at least partially. This appears in line with the most recent evidence, reported in Michalek et al. (2011), which shows that farmers gain from 60% to 95% of the value of CAP coupled payments, and only a marginal fraction of such payments is capitalised into land rent. Clearly, future research is needed to better understand these aspects.

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[26] There are several caveats behind this calculation. For example, concerning the consequences, we are assuming that the effects are fairly homogeneous across regions. Yet relaxing this assumption would be beyond the scope of our analysis.

## References

Anders, S., J. Harsche, R. Herrmann and K. Salhofer (2004), "Regional Income Effects of Producer Support under the CAP", *Cahiers d'Economie et Sociologie Rurales*, 73, pp. 103–21.

Arellano, M. and O. Bover (1995), "Another look at the instrumental variable estimation of error-components models", *Journal of Econometrics*, Vol. 68, No. 1, pp. 29–51.

Arellano, M. and S. Bond (1991), "Some tests of specification for panel data: Monte Carlo evidence and application to employment equations", *Review of Economic Studies*, Vol. 58, No. 1, pp. 227–97.

Barkley, A.P. (1990), "The Determinants of the Migration of Labour out of Agriculture in the United States, 1940-85", *American Journal of Agricultural Economics*, Vol. 72, pp. 567–73.

Becker, G.A. (1965), "Theory of the allocation of time", *Economic Journal*, Vol. 75, No. 299, pp. 493–517.

Becker, S.O., P.H. Egger and M. von Ehrlich (2010), "Going NUTS: The effect of EU Structural Funds on regional performance", *Journal of Public Economics*, Vol. 94, Nos 9-10, pp. 578–90.

——— (2012), "Too much of a good thing? On the growth effects of the EU's regional policy", *European Economic Review*, Vol. 56, No. 4, pp. 648–68.

Benjamin, C. and A. Kimhi (2006), "Farm work, off-farm work, and hired farm labour: Estimating a discrete-choice model of French farm couples' labour decisions", *European Review of Agricultural Economics*, Vol. 33, No. 2, pp. 149–71.

Berlinschi, R., K. Van Herck and J.F.M. Swinnen (2011), "Farm subsidies and agricultural employment: The education channel", paper presented at the 122nd EAAE Seminar, Ancona, Italy, 17-18 February.

Blundell, R. and S. Bond (1998), "Initial conditions and moment restrictions in dynamic panel data models", *Journal of Econometrics*, Vol. 87, No. 1, pp. 115–43.

Bond, S. (2002), *Dynamic panel data models: A guide to micro data methods and practice*, Working Paper cw09/02, Cemmap, Institute for Fiscal Studies, London (http://cemmap.ifs.org.uk/wps/cwp0209.pdf).

Breustedt, G. and T. Glauben (2007), "Driving Forces behind Exiting from Farming in Western Europe", *Journal of Agricultural Economics*, Vol. 58, No. 1, pp. 115–27.

Butzer, R., D. Larson and Y. Mundlak (2003), "Intersectoral Migration in Southeast Asia: Evidence from Indonesia, Thailand, and the Philippines", *Journal of Agricultural and Applied Economics*, Vol. 35, pp. 105–17.

Ciaian, P. and d'A. Kancs (2012), "The Capitalization of Area Payments into Farmland Rents: Micro Evidence from the New EU Member States", *Canadian Journal of Agricultural Economics*, forthcoming.

Ciaian, P., d'A. Kancs and J. Michalek (2011), *SFP Capitalization into Land Value: Generalized Propensity Score Evidence from the EU*, LICOS Discussion Paper 293/2011, LICOS, Leuven.

D'Antoni, J. and A.K. Mishra (2010), *Agricultural Policy and Its Impact on Labour Migration from Agriculture*, selected paper prepared for presentation at the Southern Agricultural Economics Association Annual Meeting, Orlando, FL, 6-9 February.

Dennis, B.N. and T.B. Iscan (2007), "Productivity growth and agricultural out-migration in the United States", *Structural Change and Economic Dynamics*, Vol. 18, No. 1, pp. 52–74.

Dewbre, J. and A.K. Mishra (2007), “Impact of Program Payments on Time Allocation and Farm Household Income”, *Journal of Agricultural and Applied Economics*, Vol. 39, No. 3, pp. 489–505.

El-Osta, H.S., A.K. Mishra and M.C. Ahearn (2004), “Labor Supply by Farm Operators under ‘Decoupled’ Farm Program Payments”, *Review of Economics of the Household*, Vol. 2, No. 4, pp. 367–85.

Esposti, R. (2007), “Regional growth and policies in the European Union: Does the Common Agricultural Policy have a counter-treatment effect?”, *American Journal of Agricultural Economics*, Vol. 89, No. 1, pp. 116–34.

———— (2011a), *Evaluating the CAP reform as a multiple treatment effect: Evidence from Italian farms*, paper presented at the 122nd EAAE Seminar, Ancona, Italy, 17-18 February.

———— (2011b), “La chiave e la luce: perché valutare la riforma del primo pilastro della PAC è difficile”, *AgriRegioniEuropa*, Vol. 7, No. 25.

European Commission (2010a), *Europeans, Agriculture and the Common Agricultural Policy*, Special Eurobarometer 336, Directorate-General for Agriculture and Rural Development, European Commission, Brussels.

———— (2010b), Communication on the CAP towards 2020: Meeting the food, natural resources and territorial challenges of the future, COM(2010) 672 final, European Commission, Brussels, 18 November.

European Parliament (2010), *On the future of the Common Agricultural Policy after 2013*, EP 439.972, Committee on Agriculture and Rural Development, Rapporteur: George Lyon.

Glauben, T., H. Tietje and C. Weiss (2006), “Agriculture on the Move: Exploring Regional Differences in Farm Exit Rates in Western Germany”, *Jahrbuch fur Regionalwissenschaft*, 26, pp. 103–18.

Goetz, S.J. and D.L. Debertin (1996), “Rural Population Decline in the 1980s: Impacts of Farm Structure and Federal Farm Programs”, *American Journal of Agricultural Economics*, Vol. 78, No. 2, pp. 517–29.

Goodwin, B.K. and M.T. Holt (2002), “Parametric and semiparametric modelling of the off-farm labour sypply of agrarian household in transition Bulgaria”, *American Journal of Agricultural Economics*, Vol. 84, No. 1, pp. 184–209.

Goodwin, B.K., A.K. Mishra and F. Ortalo-Magné (2005), *Land owners' riches: The distribution of agricultural subsidies*, University of Wisconsin–Madison, School of Business, Department of Real Estate and Urban Land Economics, Madison, WI.

Gullstrand, J. and K. Tezic (2008), “Who leaves after entering the primary sector? Evidence from Swedish micro-level data”, *European Review of Agricultural Economics*, Vol. 35, No. 1, pp. 1–28.

Hansen, H. and R. Herrmann (2012), “The two dimensions of policy impacts on economic cohesion: Concept and illustration for the CAP”, *Food Policy*, Vol. 37, No. 4, pp. 483–91.

Harris, J.R. and M.P. Todaro (1970), “Migration, Unemployment and Development: A Two-Sector Analysis”, *American Economic Review*, Vol. 60, No. 1, pp. 126–42.

Hennessy, T.C. and T. Rehman (2008), “Assessing the Impact of the ‘Decoupling’ Reform of the Common Agricultural Policy on Irish Farmers’ Off-farm Labour Market Participation Decisions”, *Journal of Agricultural Economics*, Vol. 59, No. 1, pp. 41–56.

Kirwan, B. (2009), “The incidence of U.S. agricultural subsidies on farmland rental rates”, *Journal of Political Economy*, Vol. 117, No. 1, pp. 138–64.

Larson, D. and Y. Mundlak (1997), "On the Intersectoral Migration of Agricultural Labour", *Economic Development and Cultural Change,* Vol. 45, January, pp. 295–319.

Lee, J.E. (1965), "Allocating Farm Resources between Farm and Nonfarm Uses", *American Journal of Agricultural Economics,* Vol. 47, No. 1, pp. 83–92.

Lence, S. H. and A.K. Mishra (2003), "The impacts of different farm programs on cash rents", *American Journal of Agricultural Economics,* Vol. 85, No. 3, pp. 753–61.

Maddala, G.S. and S. Wu (1999), "A Comparative Study of Unit Root Tests with Panel Data and a New Simple Test", *Oxford Bulletin of Economics and Statistics,* Vol. 61, No. S1, pp. 631–52.

Michalek, J., P. Ciaian, d'A. Kancs.and S. Gomez y Paloma (2011), *Distributional Effects of CAP Subsidies: Micro Evidence from the EU,* paper presented at the Agricultural and Applied Economics Association Annual Meeting, Pittsburgh, PA, 24-26 July.

Mishra, A.K. and B.K. Goodwin (1997), "Farm Income Variability and the Supply of Off-Farm Labor", *American Journal of Agricultural Economics*, Vol. 79, No. 3, pp. 880–87.

Mundlak, Y. (1979), *Intersectoral factor mobility and agricultural growth,* International Food Policy Research Institute, Washington, D.C.

OECD (2001), *Decoupling: A Conceptual Overview*, OECD, Paris.

——— (2010), "Indicators on Employment Protection – Annual time series data 1985-2008", Version 2, last updated 24-09-2010, OECD, Paris (http://www.oecd.org).

Petrick, M. and P. Zier (2011), "Regional employment impacts of Common Agricultural Policy measures in Eastern Germany: A difference-in-differences approach", *Agricultural Economics*, Vol. 42, No. 2, pp. 183–93.

Pietola, K., M. Vare and A.O. Lansink (2003), "Timing and type of exit from farming: Farmers' early retirement programmes in Finland", *European Review of Agricultural Economics*, Vol. 30, No. 1, pp. 99–116.

Pufahl, A. and C.R. Weiss (2009), "Evaluating the effects of farm programmes: Results from propensity score matching", *European Review of Agricultural Economics*, Vol. 36, No. 1, pp. 79–101.

Roodman, D., (2009), "How to do xtabond2: An introduction to difference and system GMM in Stata", *Stata Journal*, Vol. 9, No. 1, pp. 86–136.

Salvioni, C. and D. Sciulli (2011), *Impact of Rural Development Policy and Less Favored Areas Scheme: A Difference in Difference Matching Approach*, paper prepared for presentation at the EAAE 2011 Congress, Zurich, 30 August–2 September.

Shucksmith, M., K.J. Thomson and D. Roberts (eds) (2005), *The CAP and the Regions: The Territorial Impact of the Common Agricultural Policy*, Wallingford, UK: CABI.

Tarditi, S. and G. Zanias (2001), "Common Agricultural Policy" in R. Hal, A. Smith and L. Tsoukalis (eds), *Competitiveness and Cohesion in EU Policies*, Oxford: Oxford University Press, pp. 179–16.

Todaro, M.P. (1969), "A Model of Labour Migration and Urban Unemployment in Less Developed Countries", *American Economic Review*, Vol. 59, No. 1, pp. 138–48.

Van Herck, K. (2009), *A comparative analysis of rural labour markets*, Deliverable 7.4, SCARLED Project, November.

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.<br><br>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. |
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