Reallocation of Agricultural Labor and Farm Subsidies:

Evidence From the EU Regions

Alessandro Olper

University of Milano & LICOS

Valentina Raimondi

University of Milano

Daniele Cavicchioli

University of Milano

Mauro Vigani

University of Milano

Selected Paper prepared for presentation at the International Association of Agricultural Economists (IAAE) Triennial Conference, Foz do Iguaçu, Brazil, 18-24 August, 2012.

Reallocation of Agricultural Labour and Farm Subsidies: Evidence from the EU Regions[*]

Alessandro Olper
University of Milano & LICOS

Valentina Raimondi
University of Milano

Daniele Cavicchioli
University of Milano

Mauro Vigani
University of Milano

Version: *June 2012*

Abstract

The 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 CAP payments on this important adjustment process. Using static and dynamic panel data estimators, we show that standard neo-classic drivers, like the relative income and the relative labour share, represent significant determinants of the inter-sectoral migration of agricultural labour. Overall, CAP payments contributed significantly to job creation in agriculture, although the magnitude of the economic effect was quite moderate. We also found 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.

1. Introduction

Over the last fifty years European Union (EU) countries have experienced dramatic adjustments in the agricultural labour market, showing an impressive out-farm migration

[*] A. Olper is associate professor at the Department of Economics, Management, and Quantitative Methods, Università degli Studi di Milano and associate researcher at LICOS – Centre for Institution and Economic Performance, Catholic University of Leuven; V. Raimondi and D. Cavicchioli are researchers at the Department of Economics, Management, and Quantitative methods, Università degli Studi di Milano; M. Vigani current position at European Commission, Institute of Prospective Technological Studies (IPTS), Joint Research Center (JRC) and Università degli Studi di Milano.
Corresponding author: Alessandro Olper, e-mail: alessandro.olper@unimi.it.
Financial support from the European Commission FP7 Project 'Factor Markets' is gratefully acknowledged.

of the labour force. However, in the most recent decades[1] no substantial reduction has been revealed, which is at odds with the income subsidies of €50 billion per year spent by the Common Agricultural Policy (CAP).

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

One interpretation for this counterintuitive pattern is that the agricultural subsidies were ineffective as income support policy,[2] especially because of the imperfections in both input and output markets. Different channels have been emphasized through which farm subsidies may affect agricultural employment, accounting for the mixed evidence summarized above. For example, Barkely (1990) stressed that an indirect impact of subsidies may have occurred through increased land values.[3] Indeed, land value appreciation slows down the rate of labour migration out of agriculture. Differently, Goetz and 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 a another 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, summarized above. In this paper we argue that when the direct (income) effect of farm subsidies is properly estimated, at least in the context of the EU regions investigated here,

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

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

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

we consistently find that, on average, the farm support program has a negative effect on out-farm labour participation. Thus, the CAP makes a contribution to job creation in agriculture.

The creation and maintenance of jobs in agriculture and in rural areas, has been a traditional CAP objective, an objective recently re-stated and emphasized by several EU official documents (e.g. European Commission, 2010; European Parliament, 2010).[4] However, especially due to data limitations, evidence concerning the effect of CAP subsidies on off-farm labour migration has been quite inconclusive. Moreover, it is mostly confined to specific country or regional case studies, only rarely focusing on the European-wide perspective (Shucksmith *et al.*, 2005; Petrick and Zier, 2011). Thus, 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/region analyses. However, one of the shortcomings of these studies is that the findings are difficult to generalize 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-region 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 inter-sectoral 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 analyzed focusing on both Pillar I payments (coupled and decoupled subsidies) and on several Pillar II rural development instruments. Indeed, with the exception of Petrick and Zier (2011) who studied the entire portfolio of CAP measures, previous analyses have normally considered only one

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

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

instrument at a time, or an aggregate of different policies. However, this is problematic 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 like unobserved heterogeneity, dynamics and endogeneity. Finally, we deliver 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 organized 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 to investigate the CAP effect 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.

2. Previous evidence

In this section we summarize the literature on the effect of agricultural and rural subsidies on the labour market. Theoretically, the studies can be divided into two main approaches. In one we look at agricultural household models to analyze the impact of subsidies on the allocation of household labour (Lee 1965; Becker, 1965).[6] In this framework, subsidies directed to farm income support may affect farmers' labour allocation decisions in a number of ways: increasing the marginal value of farm labour; increasing household wealth; 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 analyzed using models of occupational choice that have their roots in the Todaro (1969) and Harris and 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).

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

The above distinction is also reflected in empirical works, with studies at farm-household level largely based on micro farm-level data, and studies on the farm labour (re)allocation conducted at aggregate (country or regional) level. Micro-data allow to address the individual adjustment behavior in response to changes in factors affecting the household utility, such as different revenues sources. For example, Mishra and Goodwin (1997), using a Tobit model on farm households located in Kansas, found that policy changes that reduce farm income support can increase off-farm employment of farmers and their spouses. Similarly, El-Osta *et al*. (2004) investigated the effect of the US Agricultural Market Transition Act (AMTA) payments on agricultural labour supply using 2001 data. Results indicate that government payments tend to increase the hours operators work on-farm and *vice versa*. There are also important examples of micro-data analysis using panel data (e.g. Pietola *et al*., 2003; and Gullstrand and Tezic, 2008), and semi-parametric approaches (e.g. Pufahl and 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 structural variables such as country or regional relative income, unemployment, population densities, and institutional and policy variables. Econometric approaches of aggregate studies range

[7] In recent years, there has been a growing tendency to use semi-parametric approaches, like propensity-score matching, to study the economic effect of EU policy in general (e.g. Becker et al. 2010; 2012), and of CAP subsidies in particular (e.g. Pufahl and Weiss, 2009; Esposti, 2011a; Salvioni and Sculli, 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).

from cross-sectional to time-series analyses and, more recently, panel data methods and also quasi-experimental approaches.[8]

The seminal work of Barkley (1990) used a two-sector occupation choice model on a large time series (from 1940 to 1985) to analyze 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 be due to 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. However, this interpretation is at odds with the findings of D'Antoni and Mishra (2010) who extended the Barkley's sample to 2007, accounting also for dynamics, through an autoregressive distributed lag model. By taking dynamics into account, the farm support effect on out-farm labour migration becomes significantly negative.

At the EU level, many studies have investigated the effect of national public support policies (others than CAP payments) at the single country level (e.g. Pietola *et al*., 2003; Goodwin and Holt, 2002; Benjamin and Kimhi, 2006; Glauben *et al*., 2006;), while only a few studies have investigated the effect of CAP subsidies on out-farm migration. For both household and aggregate level empirical works, 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 and Weiss, 2009; Hennessy and Rehman, 2008; Gullstrand and Tezic, 2008), mainly as a consequence of data limitation at the EU regional level. Most of the authors used a cross-sectional approach (Breusted and Glauben, 2007; Hennessy and Rehman, 2008; Van Herck, 2009), while those who performed a panel data analysis considered only a single country and/or specific policy, such as Objective 1 or agri-environmental measures (Gullstrand and Tezic, 2008; Pufahl and Weiss, 2009; Salvioni and Sciulli, 2011).

[8] An interesting quasi-experimental approach is that proposed by Becker *et al*. (2010), who apply regression discontinuity design to investigate the effect of objective 1 fund, on both GDP *per-capita* and employment growth at EU regional level (Nuts 2).

Only a few studies have worked at the overall EU level. Breusted and 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 1993-1997 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 NUTS 2 EU 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 NUTS 2 EU regions for the 1989-2006 period. The results showed no significant effect on total employment of the Structural Funds Programme (Objective 1). Finally, Petrick and Zier (2011), using a difference-in-difference (DID) estimator on 3 East-Germany landers, found a positive effect of each coupled, decoupled and rural development payment (but not agrienvironment) on out-farm labour migration. Their DID approach represents a relevant improvement, despite the results focusing on 3 German counties are hardly 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 focused on country or regional case studies. Third, it only rarely takes into account the entire portfolio of CAP payments. Last, but not least, no particular effort was given to accounting for 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

The paper is empirical in nature. However, to rationalize our work we sketch the theory of occupational choice and labour migration decision, which has its roots in the Todaro

(1969) and Harris and Todaro (1970) two-sectors model, subsequently developed by Mundlak (1979).[9]

Following Barkley (1990), let us consider individuals facing a given return in two mutually exclusive occupations i, say agriculture (i=1), and non-agriculture 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 optimization problem as described in equation (1), where r is the discount rate

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

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 rise above the utility expected in 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 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 non-agricultural sector. For that reason, migration from agriculture to other sectors does not occur instantaneously.[10]

A potential migrant has to estimate the probability of obtaining a job in the industrial sector, to calculate H_{ik}. Clearly, this probability is affected by macroeconomic conditions, like 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. However, as most migrations are out of agriculture, migration will also

[9] The Harris and 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 that represents its innovative feature.

[10] 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 only through their effect on the relative returns to labour in the farm and non-farm sectors. For a two sectors growth model with farm-non-farm wage gap, that explicitly consider these structural parameters, see Dennis and Iscan (2007).

increase with the size of the labour force in agriculture (Larson and Mundlak, 1997). Moreover, economic conditions in the agricultural sector, like 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 farm out-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

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

where I are people employed in occupation i.

Because of people flow from one sector to another and *vice-versa*, the net out migration from agriculture 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 limitation, migration flow in both directions is 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. Barkely, 1990; D'Antoni and Mishra 2010). This approach can be a reasonable approximation when the exercise is conducted within a single country. However, working across the EU regions, 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 inter-sectoral labour migration estimates.

To reduce this potential source of bias, the approach of Larson and Mundlak (1997) was 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 farm labour out-migration 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 an heterogenous underestimation of the labour out-migration, as part-time farming differs significantly across EU regions. Thus, our empirical strategy has to be robust to this and others forms of regional heterogeneity. A second issue is that to measure migration rate we should use data on labour. However, as better explained in the data section, the disposable regional sources do not report data on agricultural labour, but rather agricultural employment. This introduces volatility into the series because we are introducing demand shocks in the migration estimates.[11]

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 to which 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:

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

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

where $\boldsymbol{Z}$ is a vector including all other observable factors like the relative labour share and the unemployment rate that affect the migration costs, *C*, and v_{it} is the error term. If the neo-classic 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 variables bias due to, and this is difficult to observe, factors correlated with our key variables of interest. We assume that the error term $v_{it} = \alpha_t + \mu_i + \varepsilon_{it}$, comprises time fixed effects common to all regions α_t, time-invariant regional fixed effects μ_i, and a time-varying component ε_{it}. Thus, by including time and regional fixed effects, equation (4) is equivalent to a difference-in-difference (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 to properly identify 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, like 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 that change very slowly over time.

Nevertheless, the inclusion of fixed effects does a good job in resolving endogeneity bias due to regional heterogeneity and/or selection bias. 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 centralized level. In principle, this assumption may be more questionable when Pillar II payments are considered. In fact, in this case the policy making process is also under the responsibility of the EU regional institutions (see Petrick and Zier, 2011), and this may generate a potential problem of endogeneity bias due to political economy motives (see Berlinschi *et al.*, 2011). However, the degree of freedom of regional governments to allocate money of Pillar II, affects only the equilibrium between different Pillar II measures (and axis), but not their aggregated level. Indeed, the overall amount of Pillar II expenditure is predetermined through a bargaining process at the EU and national level.[12] Thus, in our basic model we treat the policy variable as exogenous given. 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, thus treated as predetermined variables.[13]

A potential concern of using equation (4) is its static nature. Indeed, both D'Antoni and Mishra (2010), for the US, and Petrick and Zier (2011), for three East German Lander, showed that 'dynamics' considerations may be important in studying the effect of farm subsidies on out-farm migration.[14] To tackle this issue we estimate also a dynamic autoregressive specification

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

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 and Bond, 1991). DIFF-GMM estimator transforms the model into a two steps 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 lagged differences and levels of the dependent variable.

Moreover, because in presence of variables that display high persistency – like policy variables – the DIFF-GMM estimator could also be bias, suffering weak instrument problems, as a robustness check we also use a system GMM (SYS-GMM) estimator that exploit the second moment condition of the level equation (Arellano and

[12] Clearly this does not mean that some form of 'compensation rule', between Pillar I and Pillar II policies, cannot work here. However, to the extent to which this compensation is decided at the EU level, then it should not affect expenditure at the regional level, *ceteris paribus*.
[13] This approach also reduces the possible simultaneity in the model between farm migration and other right hand side variables, like especially relative income and unemployment.
[14] Specifically, D'Antoni and Mishra (2010) showed that moving from a static to a dynamic autoregressive specification matters for the final results. Differently, Petrick and Zier (2011) reported evidence of persistent lag structure for some CAP instruments.

Bover, 1995). Finally, an important feature of the GMM estimator is the possibility of treating the key variable of interest – namely the CAP subsidies, *S* – as endogenous, adding further credence to our findings.

4. Data

We start from an initial database of about 160 regions. However, some regions are lost 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 covered 149 regions of the 15 European Union countries, over the period 1990-2008.[15] Table 1 shows the number of regions used for each country, according to the Nomenclature of Statistical Units (NUTS) and distinguishing between the NUTS1 and NUTS2. The choice to utilize both NUTS1 and NUTS2 was motivated by the necessity to match data from different sources. Indeed the 'Farm Accountancy Data Network' (FADN) regional classification does not always match the NUTS2 level defined by Eurostat.[16]

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 be transformed as annual series only through interpolations. Thus, due to data limitation, 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

---

[15] 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, relative income). On the basis of this test, two regions, the London region in the UK and the Ovre Norrland region 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 authors upon request.

[16] An alternative solution is to apply the FADN information at NUTS1 and also for those NUTS2 regions where FADN data are lacking. However, because our empirical strategy especially exploits the within region variation in out-farm migration and CAP payments, following this approach does not add substantial 'new' information to the model structure.

the trend still prevails. The basic employment data used to measure the net migration rate comes from the Cambridge Econometric's 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 different approaches: measuring a regionalized producer subsidy equivalent (PSE) as in Anders et al. (2004), Tarditi and Zanias (2001) and, more recently, Hansen and Herrmann (2012); using the Farm Accountancy Data Network as in Shucksmith et al. (2005), and by combining the same source with Eurostat Regio-New Cronos database, assuring to the former also a time variation, 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.[17] Second, and more importantly, Eurostat data is based on agriculture sectoral series, and so do not incorporate decoupled subsidies after 2005. Thus, their use 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 to which the average farm is representative of the farm population,[18] then the computation of the ratio between such farm CAP payments and the respective farm net income (inclusive of subsidies), offers the possibility of measuring a consistent regional level of farm protection due to different CAP policy measures.

---

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

[18] 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 on 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.

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

A key advantage of our approach is the possibility of disentangling CAP total payments into their different, Pillar I and Pillar II, components. Specifically, we can distinguish between coupled and decoupled payments of Pillar I, as well as agri-environmental payments, less favoured areas (LFA), investment aids and a residual category called 'other' subsidies of Pillar II.[19] Note that some of the latter payments were introduced before Agenda 2000, thus the 'Pillar II' expression could be not fully correct. Nevertheless, we chose 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. However, it is important to note that the price component of CAP protection in our empirical model is implicitly controlled for by the relative income variable, *RI*.

4.3 Other covariates

The inter-sectoral 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 non-agriculture sector we used the difference between total GVA and GVA in agriculture, as well as for non-agricultural employment.[20] The data for GVAs and employment are from the Cambridge Econometric's Regional Database.

---

[19] 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 'other' subsidies) it is slight lower than the 'aggregate' Pillar II subsidies.

[20] Harris-Todaro type models suggest wages as a measure of (relative) labour returns. However, many papers investigating out-farm migration equation found that more robust results are obtained when relative income or productivity, instead of relative wages, is used. Mundlak (1979) and Larson and Mundlak 1997) justify this findings arguing that, for a long-run decision that involves expectations, such as the migration out of agriculture, income is thought to be a more informative measure of the future prospects than wages,

The other control variables included in the vectors *Z* are as follows. First, following Larson and Mundlak (1997) and others, we include the relative labour force (*RL*) calculated as the ratio of employments in the non-agricultural sector to that in 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 out from the farm sector. Thus, its estimated effect can be either positive or negative. Second, to control for search costs and the probability to find 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 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 amount of family workers. The underline idea is that a high number of family members working on the farm should lower exit rate (Breustedt and Glauben, 2007).

Finally, we also include a variable measuring country differences in labour market institutions, that is increasing in the rigidities of labour entry and exit. Specifically, we use the OECD employment protection indicator called 'EP_v1' (see OECD, 2010). This index is the average of 6 different sub-indices of 'regular' and 'temporary' contracts with a scale from 0 (less restriction) to 6 (most restrictions). 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 the Cambridge Econometric's Regional Database. Differently, information on farm family workers comes from FADN, while the labour institutions rigidity index is based on OECD data. Summary statistics of the variables explained above are reported in Table 2.

since wages are not the only component of farmer's income. He also notes that measurement problems with wage data provide another reason to use relative income rather than relative wages.

5. Econometric results

Table 3 reports the static DID estimate of equations (4). The specifications differ with respect to how the policy variables are considered.[21] Following D'Antoni and Mishra (2010), Augmented Dickey Fuller (ADF) tests were used to determine whether the data were stationary.[22] All variables, with the exception of relative labour and unemployment rate, were found stationary. Thus, these variables were introduced in first difference in the static DID specification.

In line with the labour migration model, the relative income between 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 of around 1.1, thus smaller than the one estimated by Barkely (1990) for the US, and equal to 4.5. 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 the findings of D'Antoni and Mishra (2010).

Considering the other covariates, family workers and the restrictiveness of labour protection institutions have the expected negative sign. Differently the effect of a change in the unemployment rate and population density were (often) unexpectedly positive. However all these additional controls are always insignificant 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 1% level. Thus, overall, the CAP played a role in keeping labour within agriculture, *ceteris paribus*. This result confirms the finding of D'Antoni and Mishra (2010) for the US economy, but it goes in the opposite direction of Petrick and Zier (2011), who

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

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

showed that with the exclusion of agrienvironmental payments, CAP subsidies significantly increase out-farm migration in three Eastern German Landers.

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.[23] Considering first 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 at only 10% level. However, this effect is heterogeneous across instruments. Splitting Pillar II policies, we find that money directed to agrienvironmental measures and the category ‘Other pillar II’ payments, contributes significantly to job creation in agriculture. Differently, LFA payments exert an effect in the same direction while investment aids, consistent with the expectation, 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 and 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 onward (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.

[23] The correlation coefficient between the two series is indeed 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 due to an omitting variable problem. To attenuate this bias, in the next section we also run regressions treating as endogenous the CAP policy variables.

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 thus not fully comparable, this change in sign is surprising. However, it is consistent with the idea that the larger the labour force in agriculture relative to the non-agriculture sector, the more out-farm migrants can be expected, namely regions tend to converge to a similar level of 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 the one where Pillar I payments are split in coupled and decoupled subsidies (see column 3). This result is puzzling. In fact, on the one hand it seems to suggest that the Fischler reform induced a policy shock that *decreased* the farmer's decision to exit the agricultural sector. On the other hand, as we will 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 of this counterintuitive result could lie in the commodities price spike of 2007 and 2008. Indeed commodity prices started to rise slowly already in 2005-2006, 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 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. However, 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*).[24] 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 regressions of Table 4. The CAP effects are still negative and always significant, with the exclusion of Pillar II payments in the OLS specification.

As is well known, these two additional estimators suffer from dynamic panel bias, lagged dependent variable being bias up- and down-ward, 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, 2006). 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 fall 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 their lagged values. For all the considered CAP payments, the estimated effect is still negative and strongly significant, showing also 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. Thus, our results are robust to possible endogeneity bias due to political economy motives and/or measurement errors in the CAP variables. Note moreover that, by treating CAP payments as endogenous we indirectly control for the

[24] This is because running GMM regressions with many left-hand side variables treated as endogenous, induce the well know problem of instruments proliferation, rendering the identifications of the CAP effects problematic.

potential omitted variables bias induced by treating Pillar I and Pillar II payments separately in the regressions.[25] We will return to this point later.

Finally, column 5 reports a further robustness check by running 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 and Bover, 1995; Blundell and Bond, 1998). In our specific situation, we especially worry about persistency in CAP subsidies when this variable are treated as endogenous However, running the system GMM regression the results are qualitatively and quantitatively very close.

5.2 Discussion

Consistent comparison between the job creation effect of different CAP policies can be made on the basis of their respective elasticities (see Table 6). Several interesting patterns emerged. 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 rise 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 accounted for.

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 indeed two times higher in absolute magnitude than the elasticity of Pillar II policies. Within Pillar I, the coupled payments display higher absolute elasticity than decoupled payments, while across Pillar II instruments, agrienvironmental payments display the higher absolute elasticity to out-

[25] More in general, 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 toward zero of the (negative) effect of subsidies on migration. Similarly, measurement error in an explanatory variable suffers the well know attenuation bias problem, which induces a bias toward zero of its estimated effect. Finally, if the endogeneity bias is the result of running regressions separately for Pillar I and II policies, then it could also be the case that this omitted variable problem can translate to an (absolute) down-ward bias in the estimated effect of the policy variables.

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.[26] Because, in principle, this incongruence cannot be attributed to 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, more than to a simultaneity problem is, if anything, due to omitted variable bias.

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 to 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 average reducing effect of CAP subsidies in terms of out-farm migration. Multiplying such value for the average stock of agricultural workers (6.897 millions/year) we can obtain a rough estimate of the flow of out-farm migration prevented by CAP payments, that is 24,247 agricultural workers per year. To render such value in percentage consider that, without subsidies, the annual out-farm migration rate would increase from the actual 0.0260 to 0.0295. The effect of CAP payments, then, reduces the rate of farm labour migration by around 11.9%, thus not an

[26] 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 indeed lower than the estimate total payment elasticity of –0.144.

irrelevant number.[27] 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 rate of out-farm migration attributable to the CAP is still positive, ranging from a minimum of about 5% to a maximum of 17%. A conservative view is thus to interpret the back-of-the envelope calculation as saying that the CAP subsidies might generate a reduction of farm out migration, although the effect can be rather moderate.

6. Conclusions

Understanding the effect 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. Within the neo-classic two sectors model, inter-sectoral labour migration is affected by across sectors income difference, *ceteris paribus*. Thus, to the extent to which 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 and across-region variation 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 CAP has played a role in keeping labour forces in agriculture, although the overall effect is rather small. Among CAP instruments, we show that Pillar I payments are, so far, the most effective policy in reducing out-farm migration, with coupled subsidies showing an elasticity to out-farm migration significantly higher than decoupled ones. Similarly, the effect of Pillar II payments on job creation is significantly lower than Pillar I payments, and conditional to 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. However, its elasticity to out-farm migration is quite low when compared to

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

similar studies conducted in other countries like the US. This suggests that at the EU regional level, out-farm migration is less responsive to income differences or, put differently, that other important forces are at work in affecting the farmers' decisions to migrate. Moreover, we also found important effects on the migration decision of standard structural variables like relative labour, 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 exert, systematically, 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 came 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, at least partially, seem to contradict this conclusion. This appears in line with the most recent evidence, reported in Michalek *et al*. (2011), that shows that farmers gain between 60% to 95% of the value of CAP coupled payments, and only a marginal fraction of such payments is capitalized in land rent. Clearly, future research is needed to better understand these aspects.

References

Anders, S., Harsche, J., Herrmann, R. and Salhofer, K. (2004). Regional Income Effects of Producer Support under the CAP. *Cahiers d'Economie et Sociologie Rurales*, 73,103-121.

Arellano, M. & Bover, O. (1995). Another look at the instrumental variable estimation of error-components models. *Journal of Econometrics*, 68(1): 29-51.

Arellano, M., Bond, S. (1991). Some tests of specification for panel data: Monte Carlo evidence and application to employment equations. *Review of Economic Studies* 58, 227–297.

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*, 567-573.

Becker, G. A. (1965). Theory of the allocation of time. *Economic Journal*, 75: 493–517.

Becker, S.O., Egger, P.H. and von Ehrlich, M. (2010). Going NUTS: The effect of EU Structural Funds on regional performance. *Journal of Public Economics*, 94: 578–590.

Becker, S.O., Egger, P.H., and von Ehrlich, M. (2012). Too much of a good thing? On the growth effects of the EU's regional policy, *European Economic Review*, 56: 648–668.

Benjamin, C. and Kimhi, A. (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*, 33(2): 149–171.

Berlinschi, R., Van Herck, K. and Swinnen, J. F.M. (2011). Farm subsidies and agricultural employment: The education channel. Paper presented at the 122nd European Association of Agricultural Economists Seminar, February 17-18, 2011, Ancona, Italy.

Blundell, R. & Bond, S. (1998). Initial conditions and moment restrictions in dynamic panel data models. *Journal of Econometrics*, 87(1): 115-143.

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

Breustedt, G. and Glauben, T. (2007). Driving Forces behind Exiting from Farming in Western Europe. *Journal of Agricultural Economics,* 58(1), 115-127.

Butzer, R., Larson, D. and Mundlak, Y. (2003), Intersectoral Migration in Southeast Asia. Evidence from Indonesia, Thailand, and the Philippines. *Journal of Agricultural and Applied Economics*, 35, 105-117.

Ciaian, P., Kancs, D., and Michalek, J. (2011). SFP Capitalization into Land Value: Generalized Propensity Score Evidence from the EU, LICOS discussion paper 293/2011.

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

D'Antoni, J. and Mishra, A. K. (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, February 6-9, 2010.

Dennis, B.N. and Iscan, T.B. (2007). Productivity growth and agricultural out-migration in the United States, *Structural Change and Economic Dynamics*, 18, 52-74.

Dewbre, J. and Mishra, A.K. (2007). Impact of Program Payments on Time Allocation and Farm Household Income. *Journal of Agricultural and Applied Economics*, 39(3): 489-505.

El-Osta, H.S., Mishra, A.K. and Ahearn, M.C. (2004). Labor Supply by Farm Operators Under ''Decoupled'' Farm Program Payments. *Review of Economics of the Household,* 2: 367–385.

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,* 89, 116–134.

Esposti, R. (2011a). Evaluating the CAP Reform as a multiple treatment effect: evidence from Italian farms. 122nd EAAE Seminar, February 17-18, 2011, Ancona, Italy.

Esposti, R. (2011b). La chiave e la luce: perché valutare la riforma del primo pilastro della PAC è difficile, *AgriRegioniEuropa*, 7 (25).
European Commission, Directorate-General for Agriculture and Rural Development, (2010). Europeans, Agriculture and the Common Agricultural Policy. Special Eurobarometer 336.
European Parliament, Committee on Agriculture and Rural Development, (2010). On the future of the Common Agricultural Policy after 2013. EP 439.972.
Glauben, T., Tietje, H. and Weiss, C. (2006). Agriculture on the Move: Exploring Regional Differences in Farm Exit Rates in Western Germany. *Jahrbuch fur Regionalwissenschaft* 26, 103-118.
Goetz, S.J. and Debertin, D.L. (1996). Rural Population Decline in the 1980s: Impacts of Farm Structure and Federal Farm Programs. *American Journal of. Agricultural Economics*, 78: 517-529.
Goodwin, B. K. and Holt, m. T. (2002). Parametric and semiparametric modelling of the off-farm labour sypply of agrarian household in transition Bulgaria. *American Journal of. Agricultural Economics,* 84(1): 184-209.
Goodwin, B. K., Mishra, A. K. and Ortalo-Magné, F. (2005). Land owners' riches: The distribution of agricultural subsidies. University of Wisconsin-Madison School of Business, Department of Real Estate and Urban Land Economics.
Gullstrand, J. and Tezic, K. (2008). Who leaves after entering the primary sector? Evidence from Swedish micro-level data. *European Review of Agricultural Economics*, 35(1): 1–28.
Hansen, H. and Herrmann, R. (2012). The two dimensions of policy impacts on economic cohesion: Concept and illustration for the CAP, *Food Policy*, 37(4): 483–491.
Harris, J.R. and Todaro, M.P. (1970). Migration, Unemployment and Development: A Two-Sector Analysis. *The American Economic Review*, 60(1): 126-142.
Hennessy, T.C. and Rehman, T. (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*, 59(1), 41-56.
Kirwan, B. (2009). The incidence of U.S. agricultural subsidies on farmland rental rates. *Journal of Political Economy*, 117 (1): 138–64.
Larson, D. and Y. Mundlak. (1997). On the Intersectoral Migration of Agricultural Labour. *Economic Development and Cultural Change*, 45 (January 1997), 295-319.
Lee, J.E. (1965). Allocating Farm Resources Between Farm and Nonfarm Uses. *American Journal of Agricultural Economics*, 47(1):83–92.
Lence, S. H. and Mishra, A. K. (2003). The impacts of different farm programs on cash rents. *American Journal of Agricultural Economics*, 85: 753–61.
Maddala, G.S. and Wu, S. (1999). A Comparative Study of Unit Root Tests With Panel Data and A New Simple Test. *Oxford Bulletin of Economics and Statistics*, 61, 631-652.
Michalek, J., Ciaian, P., Kancs, D. and Gomez y Paloma, S. (2011). Distributional Effects of CAP Subsidies: Micro Evidence from the EU. Agricultural and Applied Economics Association Annual Meeting, July 24-26, 2011, Pittsburgh, Pennsylvania.
Mishra, A.K. and Goodwin, B.K. (1997). Farm Income Variability and the Supply of Off-Farm Labor. *American Journal of Agricultural Economics*, 79(3): 880-887.
Mundlak, Y. (1979). *Intersectoral factor mobility and agricultural growth*. Washington: International Food Policy Research Institute.
OECD (2010). Indicators on Employment Protection - annual time series data 1985-2008, Version 2 - Last updated 24-09-2010. Available online at, http/www.oecd.org/.
OECD. 2001. *Decoupling: A Conceptual Overview*. Paris: OECD.
Petrick, M. and Zier, P. (2011). Regional employment impacts of Common Agricultural Policy measures in Eastern Germany: a difference-in-differences approach. *Agricultural Economics*, 42(2):183-193.

Pietola, K., Vare, M. and Lansink, A.O. (2003). Timing and type of exit from farming: farmers' early retirement programmes in Finland. *European Review of Agricultural Economics*, 30(1): 99-116.

Pufahl, A. and Weiss, C.R. (2009). Evaluating the effects of farm programmes: results from propensity score matching. *European Review of Agricultural Economics*, 36(1): 79–101.

Roodman, D., (2009). How to do xtabond2: an introduction to difference and system GMM in Stata, *Stata Journal*, 9(1), 86-136.

Salvioni, C. and Sciulli, D. (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, August 30 to September 2, 2011, ETH Zurich, Switzerland.

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

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

Todaro, M. P. (1969). A Model of Labour Migration and Urban Unemployment in Less Developed Countries. *American Economic Review,* 59, 138-48.

Van herck, K. (2009). A comparative analysis of rural labour markets. SCARLED Project deliverable 7.4, November 2009.

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: Missing is information 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 United Kingdom, the Luxembourg state-region and the Bruxelles-Capitale region in Belgium, due to lack of data. The London region in the United Kingdom and the Ovre Norrland in Sweden were dropped, being outliers (see text).

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

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

Notes: Region, and year fixed effects included in each regression. Robust standard errors clustered by regions in parentheses. *, ** and *** indicate statistically significance at 10%, 5% and 1% level, respectively.

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

Notes: year fixed effects included in each regression. Robust standard errors in parentheses. *, ** and *** indicate statistically significance at 10%, 5% and 1% level, respectively. DIFF-GMM estimator is implemented in STATA using the xtabond2 routine with option laglimits (10).

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

Notes: year fixed effects included in each regression. Robust standard errors in parentheses. *, ** and *** indicate statistically significance at 10%, 5% and 1% level, respectively. GMM estimators are implemented in STATA using the xtabond2 routine. In GMM regressions of columns 4 and 5, the CAP policy variable is treated as endogenous, and instrumented by its lagged level and differences.

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: The table reports sample mean elasticity of CAP policy variables based on difference-in-difference, and DIFF-GMM regression results of Table 3, 4 and 5, respectively.