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Common Agricultural Policy effects on dynamic labour use in agriculture

Martin Petrick and Patrick Zier

Herausgeber:
DFG-Forschergruppe 986, Humboldt-Universität zu Berlin
Philippstr. 13, Haus 12A, D-10099 Berlin
http://www.agrar.hu-berlin.de/struktur/institute/wisola/fowisola/siag

Redaktion:
Tel.: +49 (30) 2093 6340, E-Mail: k.oertel@agrar.hu-berlin.de

Common Agricultural Policy effects on dynamic labour use in agriculture

Martin Petrick[*] and Patrick Zier
Leibniz Institute of Agricultural Development in Central and Eastern Europe (IAMO), Halle (Saale), Germany

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Abstract

The aim of this study is to investigate the effects of direct payments and rural development measures of the EU's Common Agricultural Policy (CAP) on employment in agriculture. We work with a dynamic labour demand equation augmented by the full set of policy instruments of the CAP, which is estimated on a panel dataset of 69 East German regions. We present results for four estimators which differ in how they eliminate the fixed effects and how they address the endogeneity of the lagged dependent variable. The results suggest that there were few desirable effects on job maintenance in agriculture. While there is some indication that investment subsidies have halted labour shedding on farms, a rise in the general wage level reduced labour use in agriculture. Changes in direct payments had no employment effects. Generally, labour adjustment exhibits a strong path dependency.

Keywords: Agricultural employment; Dynamic panel data models; Common Agricultural Policy; East Germany.

JEL-codes: Q18; J43; C23

[*] Corresponding author.

Zusammenfassung

Das Ziel dieser Studie ist es, die Auswirkungen der im Rahmen der Gemeinsamen Agrarpolitik (GAP) gewährten Direktzahlungen und der Maßnahmen zur ländlichen Entwicklung auf die Beschäftigung im Agrarsektor zu untersuchen. Wir verwenden hierfür eine dynamische Arbeitsnachfragegleichung, welche um das vollständige Maßnahmenbündel der GAP erweitert wurde. Diese Gleichung wird für einen Paneldatensatz aus 69 ostdeutschen Landkreisen geschätzt. Wir stellen Ergebnisse für vier verschiedene Schätzer vor, die sich darin unter-scheiden, wie sie fixe Effekte und die Endogenität der verzögert abhängigen Variable kontrollieren. Die Ergebnisse legen nahe, dass es wenige wünschenswerte Effekte auf die Beschäftigungssicherung in der Landwirtschaft gegeben hat. Einige Ergebnisse sprechen dafür, dass Investitionsbeihilfen den Arbeitskräfteabbau verlangsamt haben. Ein Anstieg des allgemeinen Lohnniveaus hat den Arbeitseinsatz in der Landwirtschaft verringert. Änderungen in den Direktzahlungen hatten keinen Beschäftigungseffekt. Grundsätzlich zeigt die Anpassung des Arbeitseinsatzes eine starke Pfadabhängigkeit.

Schlüsselwörter: Landwirtschaftliche Beschäftigung; Dynamisches Paneldaten Modell; Gemeinsame Agrarpolitik; Ostdeutschland.

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1 Introduction

Agricultural employment poses a dilemma for policy makers in Europe. On the one hand, agriculture's share in employment of all West European economies has been constantly declining for decades (Tracy 1993). On the other hand, many citizens expect that safeguarding jobs should be the top priority of government. Following this logic, politicians and farm lobbyists regularly claim that a protective agricultural policy is indispensable for keeping jobs in the sector. Furthermore, it is argued that agriculture has much potential to also provide environmental services, contribute to quality of life in rural areas, and supply raw material for energy production. The "second pillar" instruments of the European Union's (EU) Common Agricultural Policy (CAP), such as investment aid, agri-environmental payments, and a broad range of rural development measures, are supposed to create employment via these additional functions (see, e.g., EC 2006). It is thus a question of high political relevance whether these policy measures can indeed halt or even reverse the persistent decline of agricultural employment.

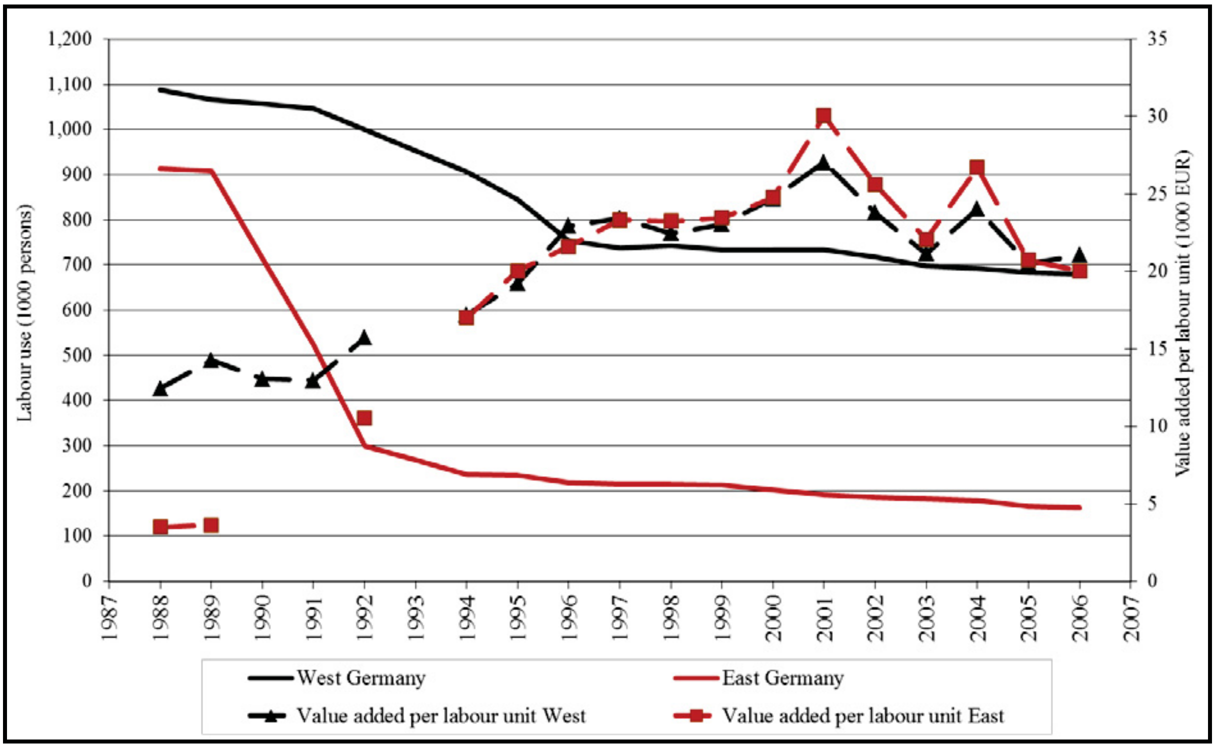
The contribution of this article is to econometrically analyse the effects of the full CAP portfolio in the framework of a dynamic labour adjustment model. The model is applied to East German county (*Landkreis*) data. Our methodological inspiration is taken from the econometric evaluation literature using panel data to address problems of unobserved heterogeneity and selectivity in programme participation (Blundell and Costa Dias 2009; Imbens and Wooldridge 2009). Such approaches have only recently been taken up in the field of agricultural policy analysis. Patton et al. (2008) as well as Kirwan (2009) used regression methods to analyse the effects of farm programs on land rental values, based on panel data. Pufahl and Weiss (2009) applied semi-parametric propensity score matching to evaluate the effects of the German agri-environmental programme on production decisions. Petrick and Zier (2011) used a difference-in-differences estimator to analyse the effects of various CAP measures on labour use in East German agriculture. While these studies shed light on interesting empirical patterns of policy treatments and economic outcomes in agriculture, their theoretical foundation tends to be weak. Their methodological choices imply that there is typically no linkage to the underlying structural models of economic behaviour (Heckman 2010).

The subsequent policy analysis aims to be both informed by theory and aware of the methodological issues important in quantitative impact evaluation. We motivate the dynamic structure of our empirical model theoretically and use fixed effects and instrumental variable regression techniques to address problems of latent heterogeneity and endogeneity. Given this specification, a key question is how consistent estimates of the adjustment coefficient can be obtained. We present several results that build on recent insights from the dynamic panel modelling literature. Our article is the first to utilise these methods for an analysis of dynamic factor adjustment in agriculture. While the dynamic structure is strongly supported by the empirical evidence, policy impacts on employment appear to be quite modest. As discussed in the conclusions section, the current set of agricultural policy instruments hardly serves the public goal of job maintenance in agriculture.

2 Labour adjustment and the Common Agricultural Policy in East Germany

In 1989, collectivised agriculture in the former German Democratic Republic (GDR) entered the transition process with a share of eight per cent in domestic value added and ten per cent in domestic employment (BMELF 1991). Eight years after the beginning of reforms, labour use had gone down by 76 per cent of the 1989 level (Figure 1), the strongest decline among all European transition countries. However, wages paid to agricultural workers unambiguously grew. Value added per labour unit increased almost six-fold between 1989 and 1995. At this time, it approached the level of the old *Länder*, which it outperformed regularly after 2000. Due to special government programmes and immediate CAP implementation, farms had more rapid and easy access to capital than other Central European transition countries. On the other hand, rising capital stocks, labour-saving technologies, the terms-of-trade shock due to unification, and a generous social safety net implying increasing reservation wages explain why labour cuts in agriculture were higher than anywhere else in the region (Forstner and Isermeyer 2000; Koester and Brooks 1997).

Figure 1: Labour use and labour productivity in German agriculture, 1988-2006



Notes: Labour use: Values for 1990 and 1993 are linear interpolations. East German data before 1990 represents stocks on September 30, all other figures are annual averages.
Value added: Labour costs not included. 4.4 DDR Mark are equivalent to 1 Deutsche Mark, following Jenkis (2005, 448). Missing data points are due to gaps in official statistics.

Sources: Authors' calculations based on Statistical Yearbooks and Destatis (2009).

Several indicators suggest that the transition process of East German agriculture had been accomplished by the mid 1990s. Labour productivity had reached the West German level. National structural policies in agriculture were made uniform in East and West after 1996. Legal and institutional structures were widely harmonised. However, twenty years after the fall of the Berlin wall, farm structures continue to differ widely between the old and new *Länder*. In the latter, land utilisation was and still is dominated by legal entities based on hired labour, often in the form of agricultural cooperatives. In 1998, the average farm size in the East was 175 ha compared to 24 ha in the West, and more than 50 per cent of East German agricultural area were cultivated by farms bigger than 1000 ha (Forstner and Isermeyer 2000, 77).

At the same time, rural unemployment rates of 25 per cent and more in conjunction with a significant outmigration of young persons pressed East German politicians to make safeguarding and creation of rural jobs their top priority. This priority was widely used as a justification for the continuing inflow of CAP transfers, which became the major political determinant of decision making in agriculture. As a result of the Agenda 2000 and the Mid-term review reforms of the CAP, East German *Länder* have been spending about two thirds of their CAP budget on direct payments, of which 75 per cent are co-financed by the EU (see Petrick and Zier, 2011, on the three *Länder* studied in the following). In 2005, the Single Farm Payment (SFP) decoupled payments from the specific cropping pattern of a given farm and from the number of animals kept. It linked them only to a certain reference area of land in agricultural use.

After 1999, structural and environmental measures were unified into the "rural development" regulation 1257/1999, which was implemented according to Objective 1 provisions in the new *Länder*. The emphasis was on instruments for the "development of rural areas" which included infrastructure investments, such as road construction and improvement. They were usually disbursed to local municipalities. Agri-environmental measures entailed payments for the maintenance of extensive grassland and the conversion to organic farming. About ten to twenty million euro were spent on compensatory allowances for less-favoured areas (LFA), as well as on investment aids and processing and marketing support. The former represented support for regions with unfavourable conditions for agriculture, while the latter two were credit subsidies for a wide range of capital investments on farms and in the downstream sector. After 2007, most of these measures were continued under the European Agricultural Fund for Rural Development regulations.

3 Theoretical considerations and hypotheses

3.1 A model of dynamic labour adjustment

The planning horizon of the farmer is assumed to start at time zero and last infinitely. In each period *t*, the farm produces a single current output as described by a production function $f(.)$ that has the current stock of labour, L_t, as its only argument. $f(.)$ is assumed to be concave, such that $f'>0, f''<0$. With static expectations, prices for output, *p*, and labour, *w*, are assumed constant over time. The farmer adjusts his plans and his targets every year as prices and technology change. However, adjustment of the labour stock is costly, as described by a convex adjustment cost function $C(\dot{L}_t)$, with $\dot{L}_t$ denoting the gross change of labour stock per period, and $C'>0, C''>0$. Furthermore, $C'\neq 0$ as $\dot{L}\neq 0$, and $C(0)=0$. Given current prices, technology and potentially other exogenous factors such as the policy environment, farmers project a desired level of employment, L^*, every period and adjust the current stock accordingly. However, as adjustment is subject to a convex cost schedule, it will be gradual over time, so that the equilibrium employment level is reached only asymptotically.

Formally, the decision problem of the agricultural firm faced at time zero is to maximise the present value, *PV*, of its earnings:

$$(1) \qquad \max_{L_t} PV = \int_0^\infty \left\{ pf(L_t) - wL_t - C(\dot{L}_t) \right\} e^{-rt} dt ,$$

subject to L_0 given, where *r* is a constant discount rate.

Using the calculus of variations to solve this problem, the first order condition for an optimal path of L_t governed by (1) is given by the following Euler equation (Nickell 1986, 482):

$$(2) \qquad pf'(L_t) = w + rC'(\dot{L}_t) - \ddot{L}_t C''(\dot{L}_t).$$

This problem is typically studied by assuming quadratic adjustment costs $C(\dot{L}) = a|\dot{L}| + b\dot{L}^2$, with $a,b>0$ (Hamermesh 1993, 210). Equilibrium labour demand in the long-run steady state, L^*, is characterised by $\dot{L} = \ddot{L} = 0$, and, as an implication of the quadratic adjustment cost function, obeys the following condition:

$$(3) \qquad pf'(L^*) = w + ra .$$

This is the familiar first-order condition from a static profit maximisation problem, except that the labour cost include both the current wage and a discounted once for all marginal adjustment cost of hiring or releasing one additional worker. Hence, in the presence of adjustment costs, it pays the firm to reduce employment as long as the foregone output is compensated by the saved adjustment costs.

A convenient implication of the assumption of quadratic adjustment costs is that it establishes a direct link to the flexible accelerator or partial adjustment model, which has been a widely used basis for empirical work on quasi-fixed factor demand (Bond and Van Reenen 2007, 4443).[1] Under quadratic adjustment costs, eq. (2) yields a general solution to the Euler

[1] In agricultural applications, also labour adjustment models based on dynamic duality have been popular, see Vasavada and Chambers (1986), Chang and Stefanou (1988), Stefanou et al. (1992), Pietola and Myers (2000).

equation in the form of a second-order linear differential equation which can be solved for its characteristic roots. As Chiang (1992, 110) shows, the characteristic roots yield a solution for the coefficient of adjustment, γ, in the following partial adjustment model:

(4) $$\dot{L}_t = \gamma(L^* - L_t).$$

Given the above theoretical framework, this equation describes how the firm partially adjusts its labour stock to the steady state through time. The speed of adjustment is determined by $0 \leq \gamma \leq 1$ and is decreasing in the level of adjustment costs (Nickell 1986, 504).

3.2 The cost of adjustment

Adjustment costs are largely determined by the specific organisational and institutional structure of the sector under study and are thus an empirical matter. Because employment protection legislation is relatively strict in Germany (OECD 2004), there will be significant firing costs due to government regulation. Moreover, in agricultural cooperatives, dismissing workers may have the consequence that these workers withdraw their share in the cooperative and hence reduce equity. Dismissed workers who are also land owners may no longer be willing to rent their plot to the large farm. These types of costs increase linearly with the number of fired workers. However, considering that typical farms in the region employ about 30 workers, releasing more than one or two workers per year may lead to significant internal disruption and reorganisation costs that increase at the margin. Other important firing costs will be of a social nature, in the sense that farm managers fear a negative reputation in the local public if they fire too many at a time (Wolz et al., 2009).

On the other hand, there is now widespread evidence that it is increasingly difficult to find trained and motivated workers in cases where they are to be hired. Recent years have seen significantly decreasing numbers of students leaving secondary schools in East Germany, thus threatening the availability of trainees for “green” jobs (Agra Europe 2010). As shown by Uhlig (2008), unemployment levels in the age class below 25 years have recently not been higher in the East German states than elsewhere in Germany. For this reason also hiring costs can be assumed to be substantial and marginally increasing.

We thus maintain the standard assumption of a convex cost schedule in the following. With regard to the empirical application, this has the advantage of motivating a simple specification of the partial adjustment model.

3.3 Hypotheses about CAP effects on labour use

In order to analyse policy effects on long-term labour demand given our theoretical model, it is crucial to identify how changes in exogenous conditions affect L^*. The standard model implies that higher output prices and less productive technology tend to increase optimal labour use, while higher wages reduce it. In addition, the following hypotheses about the impact of measures can be generated:

1. Direct payments coupled to certain production activities, such as field crops or livestock rearing, will induce additional employment if more workers are required to maintain these activities. However, as payments were no longer coupled to the level of output already in the beginning of the period observed here, allocation effects will be small. Direct payments and compensations for LFA will have no effect on labour use if they are fully decoupled.[2] A shift from a coupled to a decoupled policy regime, as implied by the CAP reform implemented in 2005, will therefore tend to release workers.

2. Most of the public goods investments, both for "rural development" or "processing and marketing", can be assumed to generate higher output prices (for example by reducing transaction or transport cost) and thus tend to increase labour use. Some may also reduce adjustment costs by making it easier to hire or release labour. For example, search costs may be lower with better infrastructure. In general, many effects of public goods investment on factor and output prices will be indirect. Note that the reduced form model estimated below measures these net effects, accounting for all direct and indirect effects at the regional level.

3. Capital subsidies will reduce labour demand if labour and capital are technological substitutes, but will induce it if they are complements.

4. Agri-environmental payments are linked to certain types of output which generate positive environmental externalities (for example, protection of biodiversity or a certain landscape, or reduced soil erosion). They hence make the production of these outputs economically more attractive. If these outputs are produced by using a more labour-intensive technology than conventional outputs, they will increase labour demand.

It is hence not unfounded to expect that agricultural policies may have positive effects on agricultural employment, although effects of different policy packages may be of opposite direction. What the effects are in reality is an empirical question that is addressed next.

4 Empirical strategy and data

We now derive an estimable equation of dynamic labour demand and discuss a number of challenges that arise in its empirical implementation.

4.1 Deriving an estimating equation

A standard approach in the labour economics literature has been to replace the unobserved $L*$ in (4) by a function $G(X)$ in order to obtain an estimable model, where X is a vector of determining variables (Hamermesh 1993, chapter 7). Such reduced-form approaches typically

[2] It has been argued that they may increase factor use via wealth and insurance effects (Hennessy 1998). Sckokai and Moro (2009) have shown recently for Italy that the risk-related effect of direct payments is small.

use output, factor stocks and/or prices as exogenous variables (Bond and Van Reenen 2007, 4478). We hypothesise that the various policy measures do have an impact on long-term labour demand. Furthermore, as the impact on agricultural employment may vary substantially among policy measures and may even be of opposite sign, we argue that it is necessary to analyse their influence simultaneously. Several of the policy measures are not directly paid to agricultural firms, in particular, processing and marketing as well as rural development funds. However, annual payment streams disaggregated by measures are available from Petrick and Zier (2011) at the county level. We therefore conduct the analysis at this level and assume that the theoretical model applies to a regionally representative farm. As we linearise the model below, it can be regarded as a consistent aggregation of individual farms. We postulate that optimal employment is determined by the following set of factors:

(5) $$L_{jt}* = G\left(\theta_{jt}, p_{jt}, \widetilde{Z}_{jt}, \overline{Z}_{j}\right),$$

where $L_{jt}*$ is the projected long-term agricultural employment in region j at time t, θ_{jt} is a vector of policy expenses that vary across regions and periods, p_{jt} is a vector of regionalised input and output prices at time t, $\widetilde{Z}_{jt}$ is a vector of regional characteristics that also vary across time and space, and $\overline{Z}_{j}$ a vector of time-invariant regional characteristics, including land endowments.

can be formulated in discrete time as follows:

(6) $$L_t - L_{t-1} = \gamma\left(L_t * - L_{t-1}\right).$$

Solving (6) for L_t and inserting yields an estimable reduced-form equation of L_t. Linearising this equation gives the following expression:

(7) $$L_{jt} = \lambda L_{jt-1} + \beta_1\theta_{jt} + \beta_2 p_{jt} + \beta_3\widetilde{Z}_{jt} + \beta_4\overline{Z}_{j} + \varepsilon_{jt},$$

where λ and β_i are parameter vectors to be estimated and ε_{jt} is an identically and independently distributed error term. Note that this partial adjustment model provides an estimate of the coefficient of adjustment, as $\gamma = 1-\lambda$. Concerning the effects of policy measures on labour demand, short-run and long-run effects have to be distinguished. Policies may affect current labour demand immediately, as measured by β_1. However, there is also a long-term effect via the dynamic adjustment process. In the steady state, $L_{jt} = L_{jt-1}$. Substituting this into and solving for L_{jt} leads to the long-run effect of θ_{jt}, which is $\frac{\beta_1}{(1-\lambda)} = \frac{\beta_1}{\gamma}$. The smaller γ, the slower is the adjustment of y to a new equilibrium and the bigger the effect of θ_{jt} that can only be observed in the long-run. If $\gamma = 1$ (or $\lambda = 0$), adjustment to the steady state is immediate and there is no sluggish adjustment at all. In this case, there is no effect that only occurs in the long-run. The model is transformed into a static model.

In the following, we wish to estimate in order to identify effects of the elements of θ_{jt} on L_{jt}. This is subject to two major methodological challenges. The first is the role of unobserved time-varying variables that may have an effect on regional policy expenses, as discussed in the literature on empirical incidence analysis that exploits variations in policies (Besley and Case, 2000; Imbens and Wooldridge, 2009). The second is the endogeneity of the lagged dependent variable, as discussed in the literature on dynamic panel data models (Arellano and Bond 1991; Kiviet 1995; Blundell and Bond 1998).

4.2 Endogeneity of policy variables

Simply regressing observed employment figures on a set of regional characteristics and policy expenses will lead to biased estimates if not all relevant control variables can be observed provided the control variables not included in the regression are correlated with the variables of interest. While some of these control variables are routinely published by statistical agencies, such as land resources or climatic conditions, others are unlikely to be easily recorded, such as regional human or social capital. Researchers relying only on observable variables make the assumption of 'unconfoundedness' or 'selection on observables' (Imbens and Wooldridge, 2009). They will generate spurious policy effects if they disregard relevant unobserved variables.

However, if the effects of time-invariant characteristics can be linearly separated, regional fixed-effects will eliminate the bias originating from observed and unobserved heterogeneity, thus allowing for 'selection on unobservables'. Forming first differences of leads to:

(8) $$L_{jt} - L_{jt-1} = \lambda\left(L_{jt-1} - L_{jt-2}\right) + \beta\left(x_{jt} - x_{jt-1}\right) + \varepsilon_{jt} - \varepsilon_{jt-1},$$

where x_{jt} is the vector of time-varying right-hand variables in . This equation shows that the influence of observed and latent characteristics of regions, as far as they are time invariant, as well as any other linear separable selection bias is 'swept out' of the equation. Because L_{jt-1} is correlated with ε_{jt-1} from , $L_{jt-1} - L_{jt-2}$ will be correlated with $\varepsilon_{jt} - \varepsilon_{jt-1}$ in (Cameron and Trivedi, 2005, 765). This latter problem will be addressed in the next section.

Besley and Case (2000) argue that regional political variables may have an effect on regional policy design. Whereas the procedures for calculating and administrating direct payments under the CAP are mostly settled at the European and national level, states have freedom to allocate funds on investment support, rural development and agri-environment. However, there is practically no decision power related to the CAP at the county level, our unit of observation.

With regard to direct payments, critical variables in determining payment streams are which crops are planted and how many animals are kept in a given region, particularly cattle. Similarly, the area under environmentally friendly practices or the farms' investment activities are determining the absorption of agri-environmental measures or capital subsidies. While these are decision variables of the farm managers and thus potentially endogenous, we maintain the assumption that there is an average potential of a region to absorb these payments. This potential is assumed to be completely determined by the given environmental conditions and human resources of that region. It can thus be eliminated by fixed effects. Changes in this potential over time are neglected. Transfers that are not paid on the basis of voluntary participation of farmers, such as public good investments or measures affecting the downstream sector are exogenous to the model per se.

4.3 Endogeneity of the lagged dependent variable

Estimating dynamic panel data models has recently been an active field of research and there is no single established estimator. The traditionally employed least squares dummy variable (LSDV) approach to eliminate fixed effects is to time demean the sample, which means either differencing out group means or using group dummy variables. In dynamic models like this approach will be inconsistent if T is small, because L_{jt-1} is endogenous (Nickell 1981). Anderson and Hsiao (1981) suggested to eliminate fixed effects by first differencing and use L_{jt-2} in an equation like to instrument $L_{jt-1} - L_{jt-2}$, as L_{jt-2} is uncorrelated with $\varepsilon_{jt} - \varepsilon_{jt-1}$. This approach yields consistent estimates if $N \to \infty$.[3] Arellano and Bond (1991) improved the efficiency of the instrumental variables approach by using additional lags of the lagged dependent variable in the framework of a Generalised Method of Moments (GMM) estimator.

The ensuing discussion in the literature showed that the true persistence of the dependent variable and the number of T available in the dataset have a decisive influence on the bias and efficiency of the estimator. Blundell and Bond (1998) argued that the Arellano-Bond estimator may perform poorly if the true autoregressive parameter is high (that is, values of 0.8 and above). Their alternative estimator, hereafter called BB, displayed notable efficiency gains if lagged differences were included as instruments into a level equation of the dependent variable. Previous studies indicate that labour adjustment in agriculture tends to be slow, so that the preferred estimator should be robust to high autoregressive parameters.

Unfortunately, the instrument proliferation in BB does not come without a cost. As highlighted by Roodman (2009a), it may bias the coefficient estimates of endogenous variables due to overfitting, weaken test procedures of instrument validity, and produce downward-biased standard errors. The latter was addressed by a variance correction for the two-step BB estimator due to Windmeijer (2005), which we also use in the following. Furthermore, Roodman (2009a) recommends testing results for sensitivity to reductions in instrument numbers.

Another concern is that the instrumental variables estimators are valid for large N, but their properties in small sample sizes are generally unknown. Analysts working with macro panels containing only a limited number of cross-sectional units have therefore questioned their usefulness for empirical work (Judson and Owen, 1999). Assuming strict exogeneity of the

[3] Consistency is achieved by the use of instruments, not by first differencing. There is no a-priori reason that makes time demeaning the preferred method of eliminating fixed effects vis-à-vis first differencing, and simulation results are inconclusive (Kiviet 1995).

right-hand variables other than L_{jt-1}, Kiviet (1995) made the case that the advantages of the LSDV approach in terms of efficiency could be combined with the consistency of the GMM estimators by using the latter for a correction of the former. Monte Carlo studies of small N and moderate T (for example $N = 100, T = 20$) by Judson and Owen (1999) used the correction factor developed by Kiviet (1995) to estimate a "corrected LSDV" (LSDVC). They showed that it outperformed the GMM approaches both in terms of bias and efficiency. Bruno (2005a) extended the LSDV correction procedure for application in unbalanced panels.

Flannery and Hankins (2010) were the first to also investigate how the choice of the estimator affects the estimates for the exogenous variables. In their simulation studies based on short panels of corporate finance data, BB and LSDVC ranked highest in accurately estimating the λ-parameter. Even so, the results also made clear that reliability of the β-estimates can be a problem even for these estimators if panels are short, the true autoregressive parameters are high, and the exogenous variables are highly persistent themselves. Despite these limitations, BB and LSDVC are the preferred estimators. We present results on both estimators in the following.

4.4 Control variables

Further exogenous variables were included in to control potentially confounding factors of the farms' external environment. Our focus was on input and output price data.

Wages and the local demographic structure are likely to vary both across time and space. Labour markets are typically local because of the inherent immobility of these factors. In addition, net migration out of rural areas may have led to local shortages of labour (Uhlig 2008). It also may have wider implications in terms of public goods provision by the government. We therefore included variables on wages and regional population density in $\tilde{Z}_{jt}$. The wage variable reflects the regional gross wage level in all sectors, including social benefits.[4]

Moreover, we included four different price indices that vary over time but not across regions, namely for agricultural plant and livestock products, for variable inputs, and for agricultural investment goods. These variables control time-related shocks and thus help to justify the assumption of independent error terms across counties.

Data on land prices was not available with sufficient coverage to be included in the model. We did not include factor stock variables other than labour. These were regarded as either exogenous and constant at the regional level, such as land, or are endogenous to a dynamic model, such as capital.

[4] Unfortunately, wage data differentiated by sectors is not available at the regional level.

4.5 Data

A key question posed by our research was which data aggregation level to use. In the present application, we work with regional CAP expenditure data described in Petrick and Zier (2011), which covers 69 counties ($N = 69$). In this dataset, information on CAP payments originates from paying agencies of the state agricultural ministries. Compared to farm-individual data, for example originating from the Farm Accountancy Data Network (FADN), regional data has the obvious disadvantage of hiding farm-specific detail and structure due to aggregation. However, this particular regional dataset has a couple of advantages, too. First, it represents the complete account of CAP expenditure per region, not only the funds received by a sample of farms. Furthermore, all actual agricultural employees in a given region are recorded. Problems of biased sample selection inevitably arising from an FADN dataset are thus avoided. In addition, FADN datasets are often lacking sufficient coverage of key variables. For example, Shucksmith et al (2005, 66) could only use data for agri-environmental payments and compensations for LFA from their EU-wide analysis of FADN data. Here we have all measures compiled from one consistent source, including those that are not directly paid to farmers, such as development of rural areas and processing and marketing.

Data for employment and the control variables except the agricultural price indices originates from Destatis (2009). Agricultural price indices were taken from BMELV. The data panel is slightly unbalanced. There is generally a coverage of 7 years in the right-hand variables, but the period covered differs by one year, depending on the state (Table 1). Furthermore, the number of lags available for the dependent variable varies between states.

Table 1: Overview of data coverage

	Brandenburg (N=16)	**Saxony (N=29)**	**Saxony-Anhalt (N=24)**
Dependent variable	1994-2006 (T=13)	1996-2006 (T=11)	1994-2006 (T=13)
Right-hand variables	2000-2006 (T=7)	2000-2006 (T=7)	1999-2005 (T=7)

Source: Authors' calculations.

The dataset distinguishes eight different policy measures. From the “first pillar” these are: coupled area payments; coupled livestock payments; and decoupled direct payments, that is the SFP introduced in 2005. From the “second pillar”, the measures are: development of rural areas; processing and marketing support; investment aids to farm businesses; compensatory payments to LFA; payments within the agri-environmental programme. Their individual components are described in the appendix of Petrick and Zier (2011). In addition, a dummy variable “decoupling” was included that takes the value of one in 2005 and 2006 and zero otherwise, to capture reform effects not due to the volume of payments. Descriptive statistics of the variables are given in Table 2. All monetary variables are in real terms, using the GDP deflator for Germany and taking 2000 as a base year.

Table 2: Descriptive statistics

		Mean	Std. Dev.	Min	Max	Obser-vations
Employees in agriculture	Persons	1913.10	1089.41	109	5337	815
Coupled area payments	Million EUR	9.43	9.22	0.00	39.88	483
Coupled livestock payments	Million EUR	1.12	1.60	0.00	10.52	483
Decoupled direct payments	Million EUR	3.30	7.79	0.00	46.85	483
Development of rural areas	Million EUR	3.91	3.24	0.00	22.30	483
Processing and marketing	Million EUR	0.33	1.52	-0.74[a]	23.19	483
Investment aids	Million EUR	0.67	0.72	0.00	4.05	483
LFA	Million EUR	0.66	0.81	0.00	3.35	483
Agri-environment	Million EUR	1.63	1.88	0.00	11.74	483
Population density	Persons/km²	288.18	380.90	41.36	1912.12	483
Average annual wage all sectors	Thousand EUR	24.62	1.40	21.21	29.15	483
Output price index plant production	2000=100	104.0	5.5	97.6	114.4	483
Output price index livestock production	2000=100	98.6	4.1	92.1	105.5	483
Input price index variable inputs	2000=100	104.0	4.5	93.2	110.3	483
Input price index agricultural investment goods	2000=100	102.7	2.7	98.7	107.8	483

Note: [a] Overpayment in some regions led to negative expenses in subsequent years. All monetary values expressed in real terms, using the GDP deflator for Germany.

Source: Authors' calculations. Data sources see text.

5 Estimation results

In Table 3, we show results for four different fixed effects specifications of . We estimated the LSDV with a first order autoregressive lag as a naïve reference model along with two versions of the Blundell-Bond estimator. One BB version uses the full instrument set that is available in the data and one collapses the number of instruments to one per lag length, following the procedure outlined in Roodman (2009a). Furthermore, we present results for a corrected LSDV estimator due to Kiviet (1995) and Bruno (2005a), by using the results from fully instrumented BB for initialisation.[5]

The LSDV and BB models use cluster robust standard errors based on the county variable, which controls for both serial correlation and heteroscedasticity in the LSDV model (Cameron and Trivedi 2005, 707). The BB models report heteroscedasticity-robust standard errors due to Windmeijer (2005) and robust tests for serial correlation due to Arellano and Bond (1991).

5 Estimations were carried out by using the routines xtreg and xtdpdsys implemented in Stata 12, as well as the user-written routines xtabond2 due to Roodman (2009b) and xtlsdvc due to Bruno (2005b). The latter was modified to accomodate the xtdpdsys results for initialisation.

The tests present no evidence of second-order autocorrelation. We also apply Hansen's *J*-test for instrument validity in the BB models (Hansen 1982). The test gave no evidence of misspecification in both models. The LSDVC model uses bootstrapped standard errors. The hypothesis that estimated parameters are all zero was clearly rejected in the LSDV and BB models, as indicated by the *F*- and χ^2-statistics. We followed Im et al. (2003) in testing for panel-specific unit roots in the employment variable under the assumption of fixed *T*, which seems appropriate given our short panel. Allowing for a time trend, the reported standardised test statistic ($Z_{\tilde{t}bar} = \text{-}5.697$) allowed us to reject the hypothesis that the panels were all non-stationary at the one per cent significance level.[6]

Our interest focuses on the evidence concerning lagged adjustment and the effects of policy measures. All models show that labour adjustment is sluggish, with a highly significant coefficient of adjustment. However, the reported levels differ. The LSDV result is notably lower than the other three, which is in accordance with the known downward bias of this estimator (Nickell 1981). The BB results are close to the LSDVC result, which adds strength to the evidence that the coefficient of adjustment is at about 20 per cent. As noted above, the use of methods robust to highly persistent data is thus warranted. The estimate means that, after a shock, it takes a bit more than three years to move halfway to the new steady state.[7] Adjustment is thus similar to the rate reported in Chang and Stefanou (1988) for Pennsylvania dairy farms and a bit slower than found by Stefanou et al. (1992) for German family farms, but considerably faster than estimated by, e.g., Vasavada and Chambers (1986) for aggregate US data and Pietola and Myers (2000) for Finnish hog farmers.

With regard to policy effects, there is some evidence on positive employment effects of investment support, which is significant at five per cent in the LSDV and the full BB model. According to the latter estimate, one million euro of investment aid per region creates about 20 jobs in agriculture in the short run. For this short run effect, approximately 50 thousand euro annually are required to create one additional job. Given the logic of our model, adjustment to a new employment equilibrium takes time, so that the full effects are visible only in the long run. Using the adjustment coefficient of the full BB model, the long run effect is 83 jobs in the steady state per one additional million euro of investment aid paid now. The estimate of the LSDVC model for this parameter is of a similar magnitude as the BB estimate, but the precision is much lower so that it fails to pass the ten per cent level of significance. The collapsed BB model could not isolate any remotely significant effect from investment aids. Generally, collapsing the instruments had no far-reaching effects on results.[8]

6 The test was carried out by using the xtunitroot command.

7 The median length of the lag can be obtained by solving for t^* in $\lambda^{t^*} = 0.5$ (Hamermesh 1993, 248), which is $t^* = \log_\lambda 0.5$.

8 We also estimated an alternative variant of the full BB model in which we relaxed the strict exogeneity conditions imposed so far. In particular, we allowed that contemporaneous direct payments, investment subsidies, and agri-environmental payments were determined endogenously. We used lags of order two back to the maximum possible as GMM-type instruments for these variables, leading to a total instrument count of 304. The results supported the estimated adjustment coefficient of the full BB and LSDVC models presented here, but tended to give larger standard errors for the policy variables.

Table 3: Regression estimates: dynamic employment in agriculture

Explanatory variables	LSDV Coeff.	LSDV Std. err.	BB full Coeff.	BB full Std. err.	BB collapsed Coeff.	BB collapsed Std. err.	LSDVC using BB full Coeff.	LSDVC using BB full Std. err.
Ag employment (lagged one year)	0.65 ***	0.05	0.76 ***	0.14	0.77 ***	0.21	0.80 ***	0.04
Coupled hectare payments	0.92	6.73	31.12	23.44	20.81	13.01	6.06	6.99
Coupled livestock payments	-10.88	12.81	16.90	31.93	1.45	13.41	-3.23	12.04
Decoupled direct payments	-4.06	6.97	22.45	18.73	14.30	11.58	1.38	6.84
Development of rural areas	-0.72	2.75	-4.41	6.92	-4.40	4.12	-0.71	2.87
Processing & marketing support	-4.68	3.07	1.48	7.25	3.59	4.27	-5.04	4.36
Investment aids	17.97 **	9.57	20.33 **	9.09	5.46	10.12	19.44	14.99
LFA	17.50	54.27	-7.49	76.87	-36.52	76.65	17.99	32.89
Agri-environmental scheme	1.55	3.27	2.21	4.30	-0.70	4.60	0.43	4.74
Decoupling (2005/6=1)	-104.81 *	58.48	-114.30	276.25	-174.97 **	83.07	-132.23 *	79.34
Population density	0.06	0.46	0.51	0.52	-0.24	0.59	0.27	0.47
Average annual wage all sectors	-51.77 ***	18.70	-59.12	64.21	-45.15	31.13	-41.47 **	18.27
Output price index plant production	0.30	1.64	-0.94	3.10	-0.23	1.68	0.10	1.57
Output price index livestock production	1.29	5.06	4.84	23.67	5.71	5.32	2.39	5.55
Input price index variable inputs	-10.78	11.91	-15.73	53.43	-19.72	13.13	-13.24	12.98
Input price index ag. investment goods	21.97	21.35	37.51	87.04	46.99 *	28.35	34.57	28.53
Elimination of fixed effects	Time demeaning		First differences		First differences		Time demeaning	
Number of instruments	0		80		28		80 (from BB full)	

Notes: *** (**,*): significant at the 1% (5%, 10%) level. Total number of observations=483. LSDV and BB's also include a constant.

LSDV: Adj. R^2 (within)=0.724. F-value (16,68)=77.29. p-value<0.001. Standard errors adjusted for 69 clusters.

BB full: Lags of order two back to the maximum possible are used as GMM-type instruments for the lagged dependent variable in the differenced equation using the two-step procedure. Lagged differences used as GMM-type instruments for the lagged dependent variable in the level equation. First differences of all right-hand variables used as standard instruments. Hansen test of overidentifying restrictions: χ^2 (63)=52.23. p-value=0.832. Wald test of jointly zero coefficients χ^2 (16)=2129.9. p-value<0.001. Arellano-Bond test for zero autocorrelation: p-value of order 1=0.017, p-value of order 2=0.153. Standard errors adjusted using Windmeijer's (2005) procedure. Estimation carried out using xtdpdsys in Stata 12. Hansen test performed by xtabond2.

BB collapsed: Lags of order two back to the maximum possible are used as GMM-type instruments for the lagged dependent variable in the differenced equation using the two-step procedure. Lagged differences used as GMM-type instruments for the lagged dependent variable in the level equation. Coefficients of GMM-type instruments are assumed constant across lag lengths, implemented by the "collapse" and "h(2)" options in xtabond2 (Roodman 2009b). First differences of all right-hand variables used as standard instruments. Hansen test of overidentifying restrictions: χ^2 (11)=11.62. p-value=0.393. Wald test of jointly zero coefficients χ^2 (16)=1062.4. p-value<0.001. Arellano-Bond test for zero autocorrelation: p-value of order 1=0.040, p-value of order 2=0.166. Standard errors adjusted using Windmeijer's (2005) procedure.

LSDVC: Standard errors bootstrapped with 100 replications, using xtlsdvc. Correction procedure is based on Bruno (2005a, b).

Source: Authors' calculations.

Overall, the results on policy effects suggest that the CAP has very limited impact on job creation or maintenance in agriculture. Compared to Petrick and Zier (2011), the evidence presented here clearly suggests that the dynamic aspects of labour adjustment must not be ignored. Furthermore, the positive effects of agri-environmental programmes on labour use found by Pufahl and Weiss (2009) as well as Petrick and Zier (2011) were not supported by our findings. On the other hand, also the negative effects arising from direct payments, rural development measures as well as processing and marketing aids were not borne out here. In line with Petrick and Zier (2011), we found evidence in favour of the view that the introduction of the SFP in 2005 led to labour shedding. The decoupling dummy reported in Table 3 turned out to be significant in the LSDV, collapsed BB and LSDVC models. This is a plausible result if decoupling allowed the release of labour no longer necessary to maintain the production levels previously required to obtain crop- and livestock-related subsidies. According to the estimates from the LSDVC model, decoupling reduced average employment by 132 workers per county in the short run, or about seven per cent of the average agricultural labour force per county. The estimated long-run effect is 660 workers, or 35 per cent of the work force. Taken together, our more complete dynamic specification of CAP employment impacts supports the global picture of limited or even negative policy effects drawn by earlier analysis, while a couple of qualifications are made in the details.

A result that is supported with high precision and similar magnitude by two of the four estimators is the negative effect of the general wage level on labour use in agriculture. The short- and long-run wage elasticities at sample means implied by the two models are reported in Table 4.

Table 4: Short- and long-run labour demand elasticities with regard to the general wage level

	LSDV	LSDVC
Short-run elasticity	-0.67	-0.53
Long-run elasticity	-1.90	-2.67

Note: Elasticities computed at sample means. General wage level is annual average gross wage in all sectors per county.

Source: Authors' calculations.

The negative sign is consistent with the theory presented in section 0. The level of the estimates indicates that labour adjustment is inelastic in the short run but will be much more elastic in the longer run. Although not the main focus of this study, this result identifies the off-farm wage level as an important driver of labour use in agriculture. On the other hand, the regional population density does not have an impact on agricultural employment. Surprisingly, practically none of the price indices turned out influential. The only exception is a weakly significant, positive effect of the price of investment goods on labour use in the collapsed BB model. A conclusion to be drawn from this outcome is that developments in exogenous prices at the national or macro level are of relatively little relevance for regionally specific employment adjustments.[9]

So which factors drive regional employment in agriculture? Based on our East German sample, we can say that the influence of agricultural policy has been modest, whereas regional (but not national) developments on factor markets (labour) played an important role. In addition, the highly significant autoregressive parameter in all models indicates a strong path dependency in labour adjustments; after all last year's labour stock is the best predictor of this year's employment level. It seems likely that on-farm organisational as well as legal constraints of labour restructuring limit the managers' leeway to freely adjust employment levels according to annual fluctuations in the farms' external environment. Note that despite the relatively few significant regressors, the given specification can explain more than 70 per cent of employment variation in the data.[10] Even so, it is clear that our approach cannot sort out all of the micro determinants of employment adjustment, such as managerial plans and abilities on the farms, differences in the farm labour force due to age and education, and the local availability of sufficiently qualified potential entrants. Future research using different types of data, including qualitative approaches, may shed further light on this issue.

6 Conclusions

In this article we presented results of four specifications of a dynamic employment equation with fixed effects, estimated on East German county level data. A consistent finding across the different estimators was that agricultural employment adjusts slowly to changes in the external environment. With an annual adjustment rate of about 20 to 25 per cent, it takes a bit more than three years to move halfway to the new steady state. While earlier studies on employment in family farms found even slower adjustment, estimates compiled by Hamermesh (1993, 254) imply it is much faster in manufacturing. Typical median lags in manufacturing are about a half to one year. Labour adjustment appears to be slow in agriculture, but even slower in family farms than in farms based on wage labour, as in our case.

Direct payments, measures for the development of rural areas, transfers to LFA and agri-environmental measures had no employment effect in any of the models. Two specifications suggest that job creation in the CAP framework was possible via capital subsidies. Such subsidies were mostly used to finance buildings or machinery. Apparently, increases in capital use were sufficiently complementary to labour that they slowed down labour cuts. According to our estimates, about 50 thousand euros of subsidies were required annually to create one additional job in the short run. However, capital subsidies are more effective in the long run,

[9] To capture the influence of other macro shocks not originating from price volatility, we replaced the price indices by seven year dummies in an alternative specification. None of these dummies turned out significant, and the other parameter estimates remained practically unchanged.

[10] This is shown by the "within" adjusted R^2 of the LSDV model in Table 3. It measures the share of the variance in the dependent variable that is explained by the right hand variables after group levels have been removed by time demeaning.

as they also affect the steady state equilibrium labour demand. Our results also provide evidence that the introduction of decoupled direct payments in 2005 accelerated labour cuts. A plausible interpretation of this finding is that workers were no longer necessary to maintain production levels required for receiving payments. A further finding was that a rising general wage level in all sectors of the economy reduces labour use in agriculture. In the long run, a one per cent rise in the wage level leads to job losses in agriculture in the range of 1.9 to 2.7 per cent.

In our view, the explicitly dynamic specification of the model is a step forward compared to earlier econometric evaluations of CAP effects on labour use. By using estimators representing the current state of the methodological literature, we gave evidence that at least some of the CAP measures help to achieve the political goal of job maintenance in agriculture. However, overall, the CAP appears to be not a particularly effective tool for active job promotion in agriculture. Among the measures studied here, there is no single policy instrument which has unambiguously positive employment effects. Furthermore, adjustment takes time, so that short-term successes in job creation are unwarranted. Following the results, economic developments outside agriculture have, via the general wage level, the most pronounced effect on labour use in the farm sector.

The lion's share of the CAP budget is still represented by the direct payments. Recent reform debates have shown that these payments are increasingly difficult to justify towards the public. Repeatedly, there have been proposals to cap them for larger farms, such as present in the region studied here. According to our analysis, moderate cuts in these unconditional payments would have no negative farm employment effects. If such effects are to be achieved by the CAP, focused instruments - such as investment aids - appear more promising.

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About the authors

PD Dr. Martin Petrick is deputy head of the department External Environment for Agriculture and Policy Analysis at the Leibniz-Institute of Agricultural Development in Central and Eastern Europe (IAMO). His major research areas are agricultural policy and rural development in Europe and Asia.

Dipl.-Ing. agr. Patrick Zier is a graduate researcher in the department External Environment for Agriculture and Policy Analysis at IAMO. His research focuses on agricultural policy analysis and structural change in Eastern Germany.

Address:
Leibniz-Institute of Agricultural Development in Central and Eastern Europe (IAMO)
Theodor-Lieser-Straße 2
06120 Halle (Saale), Germany
Phone: +49-345-29 28 120
Fax: +49-345-29 28 199
E-mail: petrick@iamo.de