CAP EFFECTS ON LABOUR USE IN AGRICULTURE: EVIDENCE FROM ALTERNATIVE DYNAMIC PANEL DATA MODELS

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CAP effects on labour use in agriculture: evidence from alternative dynamic panel data models

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Abstract

Our aim is to investigate whether the direct payments and rural development measures of the EU's Common Agricultural Policy (CAP) do make jobs in agriculture safer. We work with a dynamic labour demand equation that is augmented by the full set of policy instruments of the CAP. It is estimated on a unique regional panel dataset of three East German states for the period 1999-2006. We present results for three consistent estimators which differ in how they eliminate the fixed effects and how they instrument the lagged dependent variable, including estimators due to Arellano and Bond, Blundell and Bond, and a corrected least-squares dummy variable estimator due to Kiviet and Bruno. Our results suggest that there were few desirable effects on job maintenance or job creation in agriculture. While there is some indication that investment subsidies have halted labour shedding on farms, the introduction of the fully decoupled Single Farm Payment has likely contributed to significant job losses.

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

JEL-codes: Q18; J43; C23.

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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 first 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, agro-environmental payments, and a broad range of rural development measures, are supposed to create employment via these additional functions (see, e.g., EC 2006). The aim of the present article is to investigate whether the various CAP measures actually do make jobs in agriculture safer.

In the literature, sluggish labour adjustment in agriculture has been a long-standing issue. One prominent line of research has treated labour as a quasi-fix asset and studied interrelated factor demand functions in a framework of dynamic duality based on Epstein and Denny (1983). While the focus has more commonly been on capital, labour fixity has been the focus of several studies as well (Vasavada and Chambers, 1986; Stefanou et al., 1992; Pietola and Myers, 2000). These approaches establish a strong link between theory and estimation, but they have rarely been used to directly analyse policy effects on labour adjustment. Lacking data at the farm level may be one of the reasons for this neglect, however, it is also unclear how policy measures could be included in the dynamic duality formulation employed by these approaches. Recently, in studying capital investments, Sckockai and Moro (2009) and Serra et al. (2009) have made area payments an argument of the intertemporal utility function of the farmer. While this may be regarded a pragmatic solution, it is less clear how it could be extended to other policy measures. For example, several of the rural development measures are not even paid to farmers directly, but rather to local governments or downstream processors. Furthermore, if the analysis of policy effects moves to the centre of analysis, unobserved heterogeneity of beneficiaries and the endogeneity of programme participation becomes a core methodological problem (Besley and Case, 2000; Blundell and Costa Dias, 2009; Heckman and Vytlacil, 2007). The use of panel data methods has proliferated in this area, however, the highly non-linear models in the tradition of intertemporal factor demand analysis have hardly been able to exploit the power of these methods.[2]

In the following, our goal is to estimate policy effects on labour use in agriculture. Our workhorse is a dynamic labour demand equation that is augmented by the full set of policy instruments of the CAP. We estimate this on a unique regional panel dataset of the three East German States Brandenburg, Saxony, and Saxony-Anhalt. The slightly unbalanced dataset contains seven years of observations for 69 counties (*Landkreise*). Focusing on a single linearised equation allows us to make use of recent methodological advances in the analysis of dynamic panel data. Furthermore, by way of regional aggregation, we can consider the entire portfolio of first- and second-pillar measures simultaneously.[3]

The main part of the paper considers a quantitative evaluation of policy effects that builds on current methods for estimating dynamic panel data models with fixed effects. We provide results of a least-squares dummy variable (LSDV) model with a first order autoregressive lag as a naïve reference model. This model is known to give biased results but has the general property of producing small standard errors. There is an ongoing discussion which alternative performs best in samples of a moderate number of cross-sectional units, N, and a small number of periods, T. We present results for three consistent estimators which differ in how they eliminate the fixed effects and how they instrument the lagged dependent variable. We consider estimators due to Arellano and Bond (1991), Blundell and Bond (1998), and a corrected LSDV estimator due to Kiviet (1995) and Bruno (2005a). Our results suggest that there were few desirable effects on job maintenance or job creation in agriculture. While there is some indication that investment subsidies have halted labour shedding on farms, the introduction of the fully decoupled Single Farm Payment (SFP) has likely contributed to significant job losses.

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2 One exception to this rule is Thijssen (1996), who uses dynamic panel data methods to remove unobserved heterogeneity in an analysis of agricultural investment behaviour.

3 The impact of policy measures on agricultural structures at the regional level has been studied by Ahearn et al. (2005), Goetz and Debertin (2001), and Lence and Mishra (2003). While there is hence a methodological

In the following section 2, we give some background information on labour adjustment in East German agriculture. Section 3 presents a theoretical model of intertemporal labour demand and derives hypotheses about policy effects. Section 4 explains our empirical strategy and data. Section 5 presents the estimation results, while section 6 concludes.

2 Labour adjustment and the Common Agricultural Policy in East Germany after 1989

In 1989, collectivised agriculture in the former German Democratic Republic (GDR) entered the transition process with a share of eight percent in domestic value added and ten percent in domestic employment (BMELF 1991). Eight years after the beginning of reforms, labour use had gone down by 76 percent of the 1989 level (Figure 1), the strongest decline among all European transition countries. However, adjustments in the wage level paid to agricultural workers were unambiguously upwards. 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).

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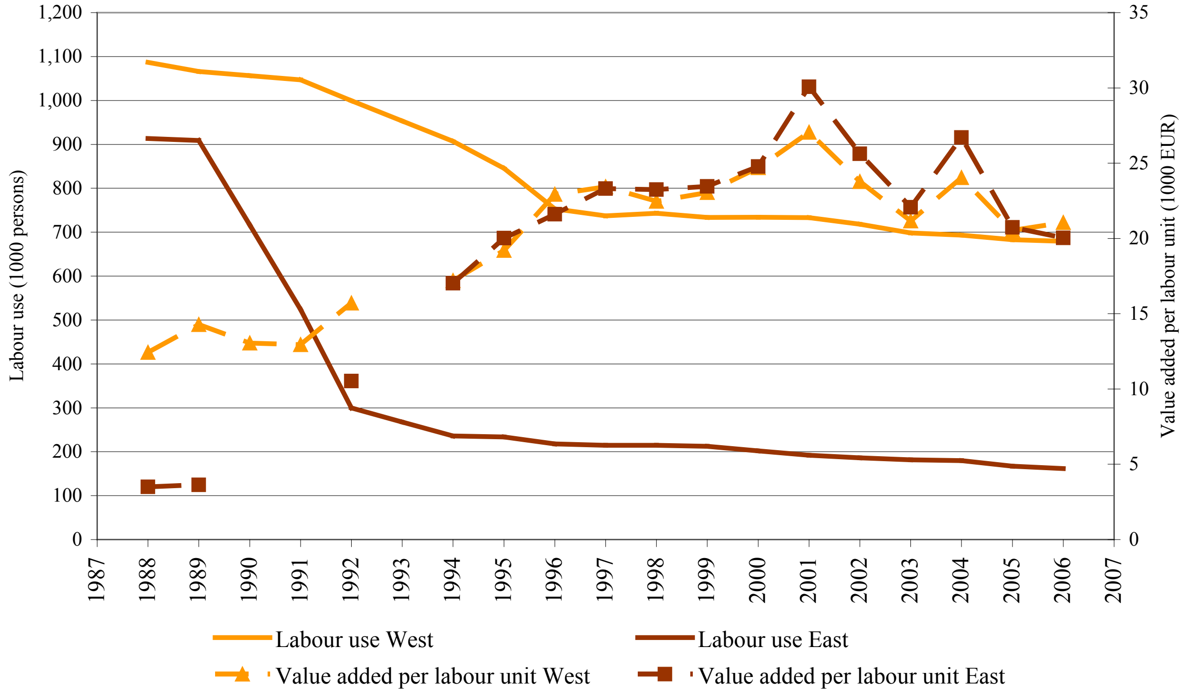
link to the present study, these authors neither look at intertemporal factor demand nor do they address the endogeneity of programme participation.

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



Notes: Labour use East Germany in 1988 and 1989 represent stock on September 30, other labour use figures are annual averages. Labour use in East and West Germany for 1990 and 1993 were estimated. Value added for East Germany in 1988 and 1989 converted into *DM* by a rate of 4.4 *DDR Mark*/*DM* (Jenkis 2005, 448). Value added is output value minus value of purchased inputs (without labour).

Sources: Labour use: 1988; 1989; 1991 Statistical Yearbooks of the GDR and FRG, var. years; 1992; 1994-2006: Destatis (2009). Value added: 1988-1990 Statistical Yearbooks of the GDR and FRG, various years; 1991-2006: Destatis (2009). Authors' calculations.

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 (Rudolph 2005). 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. In 1998, the average farm size in the East was 175 ha compared to 24 ha in the West, and more than 50 percent 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 percent 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. While farm structures are regarded as internationally competitive, politicians in the new *Länder* place much hope on the ability of agriculture to provide additional services and create new jobs, beyond their traditional role of producing food and fibre.[4] As a result of Agenda 2000 and 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 percent are co-financed by the EU (see Figure 2 for the three *Länder* studied in the following). The Single Farm Payment (SFP) was implemented in 2005. This reform decoupled payments from the specific cropping pattern of a given farm and from the number of animals kept, linking them only to a certain reference area of land in agricultural use. Single farms receiving more than 300 thousand EUR of direct payments annually have been no exception.[5] In addition, 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*. These programmes are regularly complemented by federal and state funding in the framework of the "Joint Task for the Improvement of Agricultural Structures and Coastal Protection" (Gemeinschaftsaufgabe Verbesserung der Agrarstruktur und des Küstenschutzes, GAK) (Rudolph 2005).

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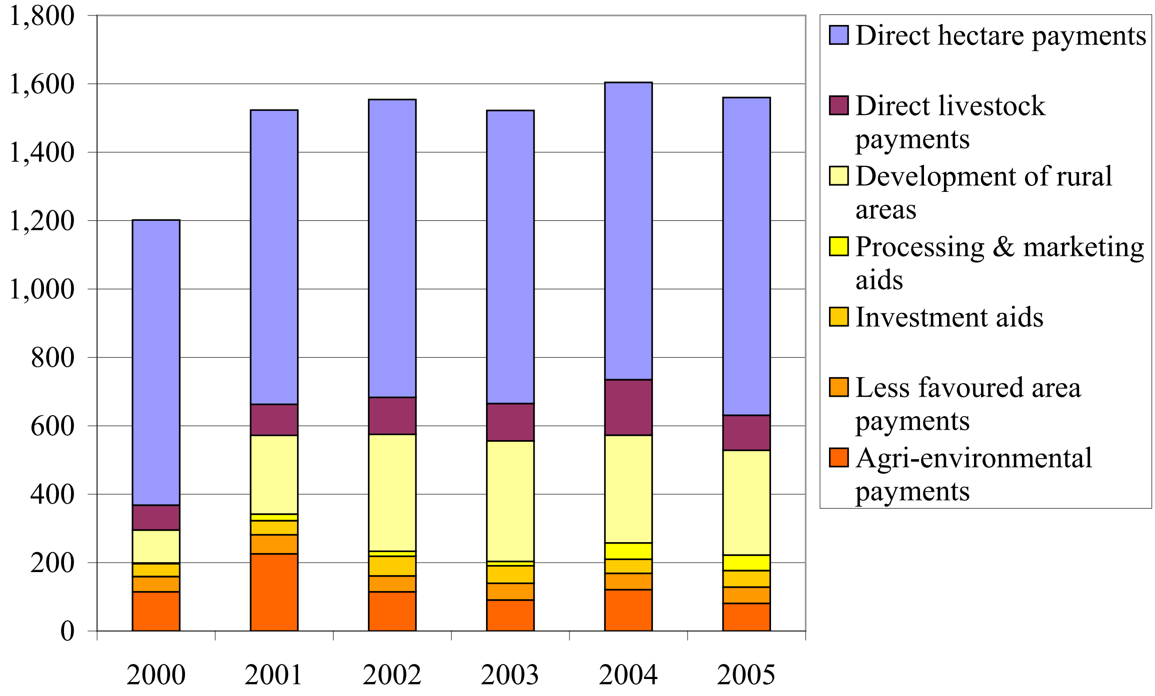
4 This is documented in various policy statements at the state level, for example in the Agricultural Report of the Land Brandenburg (Ministerium für Ländliche Entwicklung, Umwelt und Verbraucherschutz des Landes Brandenburg (MLUV) 2008).

5 This is the threshold above which farms are subject to additional modulation after the "health check" reforms of November 2008.

Figure 2: Aggregate annual CAP expenses in Brandenburg, Saxony, and Saxony-Anhalt according to main policy instruments (Million EUR)



Source: Authors' calculations based on unpublished data of state paying agencies.

East German States implemented a region-specific mix of second pillar measures. The emphasis is on instruments under the umbrella of "development of rural areas". These are mostly related to infrastructure investments, such as road construction and improvement, and are usually disbursed to local municipalities. The second largest portion of the second pillar measures goes to agro-environmental measures, which include payments for the maintenance of extensive grassland and the conversion to organic farming. In addition, some ten to twenty million euro are spent on compensatory allowances for less favoured areas (LFA), as well as on investment aids and processing and marketing support. While the former represents support for regions with below average soil conditions, the latter two are credit subsidies for a wide range of capital investments on farms and in the downstream sector.

3 Theoretical considerations and hypotheses

3.1 A theoretical model of dynamic labour adjustment

To study the effects of CAP measures on employment in agriculture, we consider a dynamic model of the price-taking agricultural firm with convex adjustment costs in labour (Nickell

1986, 481; Hamermesh 1993, 210; Chiang 1992, 106). 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.[6]

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

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

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):

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

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[6] The assumption of convex adjustment costs has been discussed controversially in the literature (Hamermesh and Pfann, 1996). In the given context of East German agriculture (compare Figure 1), the relevant question is whether, on the way to a lower steady state employment, the marginal separating costs increase with the number of workers. Many of the 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 (Welschof et al., 1993). It seems plausible that these social costs are marginally increasing.

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:

$$pf'(L^*) = w + ra\,. \tag{3}$$

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). Under quadratic adjustment costs, eq. (2) yields a general solution to the Euler 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:

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

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).

In order to analyse policy effects on long-term labour demand, it is crucial to identify how changes in exogenous conditions affect L^*. The model so far suggests that higher output prices and less productive technology tend to increase optimal labour use, while higher wages and higher one-time adjustment costs reduce it.

3.2 Hypotheses about CAP effects on labour use

We now turn to a theoretical analysis of the policy measures listed in Figure 2. The theoretical model presented in section 3.1 is clearly too simple to derive the effects of these measures on labour demand. In particular, some of the measures explicitly address production factors not included in the theoretical model so far. However, the model may be extended informally to generate hypotheses about the impact of measures:

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 generated already in the beginning of the period observed here, allocation effects will be small. Direct payments and payments for less-favoured areas will have no effect on labour use if they are fully decoupled.[7] A shift from a coupled to a decoupled policy regime, as implied by the CAP reform implemented in 2005, will therefore tend to release employment.
2. Most of the public goods investments, both for "rural development" or "processing and marketing", can be assumed to generate higher output prices (if only 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, and thus also increase equilibrium labour use.
3. Capital subsidies will reduce labour demand if labour and capital are 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. In the framework of the above model, technology parameters will be decisive for these effects. What the effects are in reality is an empirical question that is addressed next.

[7] 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.

4 Empirical strategy and data

4.1 Deriving an estimating equation

A standard approach in the labour economics literature has been to replace the unobserved $L*$ in eq. (4) by a function $G(X)$ in order to obtain an estimable model (Hamermesh 1993, chapter 7). Such reduced-form approaches typically use output, factor stocks and/or prices as exogenous variables (Bond and Van Reenen 2007, 4478). However, in the agricultural economics literature, another approach has become more popular. It has followed Epstein and Denny (1983) by using a multivariate version of eq. (4) to estimate dynamic interrelated factor demand functions in the framework of a dual optimisation model. Applications to agricultural labour in the US include Vasavada and Chambers (1986), Howard and Shumway (1988), and Luh and Stefanou (1996). Stefanou et al. (1992) studied labour adjustment in the German dairy sector and Pietola and Myers (2000) in the Finnish hog industry. The typical procedure has been to start with a specific functional form for the value function from which dynamic factor demand functions are derived by using a generalised version of Hotelling's lemma. In this way, the constraints implied by the theoretical model can either be tested or imposed and the parameters of the value function can be recovered by means of prices (see Mundlak, 2001, for a discussion of these approaches). The analysis is either based on aggregate time series or on firm-level panel data. However, these studies usually do not directly evaluate the effects of policy measures.[8] The focus has rather been on empirically identifying asset fixity and the interrelatedness of factor demand.

Our approach is different in that we emphasise the analysis of policy effects on long-term labour equilibrium by focusing on the single dynamic labour equation only. 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 (section 3.2), 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 at the regional (*Landkreis*) level. We therefore conduct the analysis at this level and assume that the theoreti-

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[8] Recent exceptions are Sckokai and Moro (2009) and Serra et al. (2009) who made direct payments an argument of the intertemporal utility function. However, they analyse investment and not labour adjustment.

cal model of section 3.1 applies to a regionally representative farm. As we linearise the model below, it can be regarded as a consistent aggregation of individual farms.

Maintaining the assumption of static expectations, farmers update their optimal labour demand each period, based on current prices and exogenous conditions. We hence postulate that optimal employment is determined by the following set of factors:

$$L_{jt}* = G\left(\theta_{jt}, p_{jt}, \tilde{Z}_{jt}, \bar{Z}_j\right), \tag{5}$$

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 prices at time *t*, $\tilde{Z}_{jt}$ is a vector of regional characteristics that also vary across time and space, and $\bar{Z}_j$ a vector of time-invariant regional characteristics, including land endowments.

Eq. (4) can be formulated in discrete time as follows:

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

Solving (6) for L_t and inserting (5) yields an estimable reduced-form equation of L_t. Linearising this equation gives the following expression:

$$L_{jt} = \lambda L_{jt-1} + \beta_1\theta_{jt} + \beta_2 p_{jt} + \beta_3\tilde{Z}_{jt} + \beta_4\bar{Z}_j + \varepsilon_{jt}, \tag{7}$$

where β_i and λ 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 eq. (7) 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 (7) by using our East German county level data set 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 regional policies (Besley and Case 2000; Smith 2004). The second is the endogeneity of the lagged dependent variable, as discussed in the literature on dynamic panel data models (Baltagi, 2008, chapter 8; Cameron and Trivedi, 2005, 763-768).

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 characteristics, which serve as control variables, can be observed. While some of these 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. This approach, called 'selection on observables' (Smith 2004, 297), will lead to spurious policy effects if such variables are omitted.

However, utilising the panel structure of the data provides a remedy for this problem. If the effects of time-invariant characteristics can indeed be linearly separated, regional fixed-effects will eliminate the bias originating from observed and unobserved heterogeneity, thus allowing for 'selection on unobservables' (Smith 2004, 304). Forming first differences of (7) leads to:

$$L_{jt} - L_{jt-1} = \lambda\left(L_{jt-1} - L_{jt-2}\right) + \beta_i\left(x_{ijt} - x_{ijt-1}\right) + \varepsilon_{jt} - \varepsilon_{jt_0}, \tag{8}$$

where x_{ijt} denotes the time-varying right-hand variables in (7). 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. However, because L_{jt-1} is correlated with ε_{jt-1} from eq. (7), $L_{jt-1} - L_{jt-2}$ will be correlated with $\varepsilon_{jt} - \varepsilon_{jt-1}$ in eq. (8) (Cameron and Trivedi, 2005, 765). This latter problem will be addressed in the next section.

As noted by Besley and Case (2000), a critical assumption in the estimation of equations like (7) is that the specification does not leave out any time-*varying*, region-specific variables that may have an influence on both θ_{jt} and L_{jt}. This assumption must be fulfilled for all of the estimators discussed before. We therefore briefly discuss the potential for policy endogeneity in our regressions. Clearly exogenous determinants of payment streams include price and subsidy levels as well as land endowments. All of these are likely to be either constant across regions (output prices and most factor prices and subsidy levels) or state-specific but time invariant (land resources). The former will be captured by a set of year dummies which are added to the estimating equation, the latter by regional fixed effects. Furthermore, 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. The question is thus which variables remain in $\tilde{Z}_{jt}$ that cannot be controlled for by using fixed effects.

Besley and Case (2000) emphasise the importance of regional political variables that may have a bearing on regional policy design. This determinant can be largely ruled out here, as the underlying political decisions are mostly made on a European level, with only some leverage left at the *Länder*, but not at the *Landkreis* level. Whereas the procedures for calculating and administrating direct payments are mostly settled at the European and national level, *Länder* have freedom to allocate funds within their Rural Development Plans (to be cofinanced from the EAGGF Guarantee section) and their Operational Programmes (in East Germany to be cofinanced from the EAGGF Guidance section) (Schubert 2002). Regional programmes are thus focusing on agri-environment and farm structures, and state governments can decide how to use funds from the modulation of direct payments. However, there is practically no decision power related to the CAP at the *Landkreis* 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. Similarly, the area under environmental-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.

Two variables that plausibly do vary across regions are the prices of labour as well as the local demographic structure. Labour markets are typically local because of the inherent immobility of these factors. In addition, net migration out of rural areas has been particularly strong in the age class between 18 and 29 years and 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 capture these trends by including variables on wages and regional population density.[9]

4.3 Endogeneity of the lagged dependent variable

Estimation of dynamic factor demand equations has been an active field of methodological research recently. The challenge has been to derive consistent estimators which are capable to eliminate fixed effects but nevertheless make efficient use of the data and perform well in samples of moderate size. As shown by Nickell (1981), the traditionally employed least squares dummy variable (LSDV) approach to eliminate fixed effects in (7) will be inconsistent if *T* is small, because L_{jt-1} is endogenous. Anderson and Hsiao (1981) have suggested to eliminate fixed effects by first differencing and use L_{jt-2} in an equation like (8) 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 $N \to \infty$. Arellano and Bond (1991) improved the efficiency of the instrumental variables approach by using further lags of the lagged dependent variable in the framework of a Generalised Method of Moments (GMM) estimator. Blundell and Bond (1998) showed that further efficiency gains are possible by also including lagged differences as instruments into a level equation of the dependent variable.

The former estimators are valid for large *N* and their properties in small sample sizes are generally not known. Analysts working with macro panels containing only a limited number of cross-sectional units have therefore argued that their usefulness for empirical work may be doubtful (Judson and Owen, 1999). Kiviet (1995) has argued 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". They show that it outper-

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[9] Data on land prices was not available with sufficient coverage to be included in the model.

formed the GMM approaches both in terms of bias and efficiency. Bruno (2005a) extended the correction procedure for application in unbalanced panels.

4.4 Data

In the present application, we work with $N = 69$, whereby the 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.

Given this data set-up, there is no unambiguous preference for one of the estimation approaches outlined before, except that the uncorrected LSDV is theoretically inconsistent. However, it is clear that our N is much smaller than in the typical applications of panel data methods to firm or household data covering several thousands of observations. As such, the corrected LSDV results may be regarded as most reliable among the four.

Data on CAP payments was collected from paying agencies of the state agricultural ministries for the periods given in Table 1. All other data was taken from official statistics (Destatis 2009). Descriptive statistics by state are given in Table 2.

Table 2: Descriptive statistics

		Mean	*Std. Dev.*	*Min*	*Max*	*N*
Brandenburg						
Employees 1st sector	n	2736	1267	227	5337	208
Direct hectare payments	Mln EUR	18.330	10.673	0.914	45.410	112
Direct livestock payments	Mln EUR	3.534	2.258	0.074	11.025	112
Development of rural areas	Mln EUR	4.149	3.134	0	17.412	112
Processing and marketing	Mln EUR	0.303	0.678	0	3.597	112
Investment aids	Mln EUR	0.879	0.648	0.001	3.286	112
Less favoured areas	Mln EUR	1.581	0.852	0.082	3.351	112
Agri-environment	Mln EUR	2.686	1.332	0.257	5.585	112
Population density	n/km²	146	171	41	764	112
Average annual wage all sectors	EUR	26.49	1.29	24.12	29.18	112
Saxony						
Employees 1st sector	n	1734	854	109	3825	319
Direct hectare payments	Mln EUR	8.593	6.194	0.022	22.790	203
Direct livestock payments	Mln EUR	1.013	0.757	0.016	3.930	203
Development of rural areas	Mln EUR	3.400	2.728	0	10.694	203
Processing and marketing	Mln EUR	0.375	2.202	0	24.298	203
Investment aids	Mln EUR	0.871	0.881	0	4.096	203
Less favoured areas	Mln EUR	0.559	0.694	0	3.088	203
Agri-environment	Mln EUR	1.653	2.449	0	11.886	203
Population density	n/km²	395	420	71	1696	203
Average annual wage all sectors	EUR	25.11	1.71	21.21	30.93	203
Saxony-Anhalt						
Employees 1st sector	n	1493	838	271	3924	312
Direct hectare payments	Mln EUR	14.071	9.109	0.500	41.109	168
Direct livestock payments	Mln EUR	0.846	1.087	0.021	7.003	168
Development of rural areas	Mln EUR	4.759	4.016	0	23.530	168
Processing and marketing [a]	Mln EUR	0.324	1.103	-0.779	12.914	168
Investment aids	Mln EUR	0.344	0.449	0	2.644	168
Less favoured areas	Mln EUR	0.239	0.454	0	2.068	168
Agri-environment	Mln EUR	1.019	1.032	0.049	4.860	168
Population density	n/km²	254	396	42	1912	168
Average annual wage all sectors	EUR	25.04	1.10	22.69	28.71	168

Note: [a] There was occasional overpayment in some regions, which led to negative expenses in subsequent years.

Source: Authors' calculations.

5 Estimation results

In the following, we show results for four different fixed effects specifications of eq. (7). We estimated the LSDV with a first order autoregressive lag as a naïve reference model (Model

A) along with two asymptotically consistent estimators which differ in their instrument set, the Arellano-Bond (model B) and the Blundell-Bond (model C) estimator. Furthermore, we present results for a corrected LSDV estimator due to Kiviet (1995) and Bruno (2005a) (model D), by using the Arellano-Bond results for initialisation.[10]

In addition to the lagged employment variable, the equation contains the seven policy aggregates listed in Figure 2. To analyse the effect of the Single Farm Payment (SFP), we add a dummy variable "Decoupling" which takes the value of one in 2005 and 2006, and zero before.[11] Furthermore, an average annual wage for all sectors, the regional population density, and year dummies for the period 2000-2004 are included.

In model (B), one cross section of observations is lost due to first differencing. Models (A) to (C) use cluster robust standard errors based on the county variable, which controls for both serial correlation and heteroscedasticity in model (A) (Cameron and Trivedi 2005, 707). Models (B) and (C) report heteroscedasticity-robust standard errors and robust tests for serial correlation due to Arellano and Bond (1991). The tests present no evidence of second-order autocorrelation. Model (D) uses bootstrapped standard errors. The hypothesis that estimated parameters were all zero was clearly rejected in all models, as indicated by the F- and χ^2-statistics.

Comparing the p-values of models (B) to (D) in Table 3, there is no unambiguous ranking of estimators in terms of efficiency possible. Neither is the Blundell-Bond model clearly superior to the Arellano-Bond, nor does the corrected LSDV estimator outperform the other two.

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[10] Estimations were carried out by using the routines xtreg, xtabond, and xtdpdsys implemented in Stata 11, as well as the user-written routine xtlsdvc due to Bruno (2005b).

[11] For 2005 and 2006, direct payments under the SFP were split into area and livestock payments according to the average distribution of the latter between 2000 and 2002, following official calculation rules (BMVEL, 2005).

Table 3: Regression estimates: policy impacts on employment in agriculture

	LSDV (A)			*Arellano-Bond (B)*			*Blundell-Bond (C)*			*Corrected LSDV using (B) (D)*		
Explanatory variables	*Coefficient*		*p-value*	*Coefficient*		*p-value*	*Coefficient*		*p-value*	*Coefficient*		*p-value*
Ag employment (lagged one year)	0.64	***	< 0.001	0.45	***	< 0.001	0.81	***	< 0.001	0.76	***	< 0.001
Direct hectare payments	-9.02		0.391	-15.61		0.176	12.88		0.136	-5.06		0.475
Direct livestock payments	4.01		0.788	-8.06		0.645	25.64		0.160	7.46		0.517
Development of rural areas	0.09		0.972	1.48		0.689	-4.71		0.416	0.05		0.988
Processing & marketing support	-6.38	**	0.029	-0.51		0.901	-1.56		0.696	-6.68		0.112
Investment aids	17.90	**	0.047	8.02		0.367	23.72	**	0.039	18.82		0.203
Less favoured areas	-6.24		0.906	-63.96		0.330	-29.49		0.498	-0.81		0.980
Agri-environmental scheme	1.73		0.614	5.72		0.222	-0.20		0.968	0.78		0.873
Decoupling (1999-2004=0; 2005/6=1)	-97.90		0.110	-151.77	**	0.025	-120.88	**	0.047	-71.50	*	0.082
Population density	-0.37		0.393	-1.40	**	0.036	0.13		0.637	-0.13		0.786
Average annual wage all sectors	-35.37	**	0.035	-47.81	*	0.071	4.53		0.506	-28.11		0.108
Number of instruments	--			71			79			71		
Number of observations	483			414			483			483		

Notes: All models include year dummies for 2000-2004 (parameters not reported). *** (**,*): significant at the 1% (5%, 10%) level.
Model (A): Includes dummy variables for 69 clusters. Adj. R^2 (overall)=0.934. F-value (16,68)=75.53. p-value<0.001. Standard errors adjusted for 69 clusters. Model
Model (B): Variables transformed into first differences. Lags of order two back to the maximum possible are used as GMM-type instruments for the lagged dependent variable using the one-step procedure. First differences of all right-hand variables used as standard instruments. Wald test of jointly zero coefficients χ^2 (16)=459.1. p-value<0.001. Arellano-Bond test for zero autocorrelation: p-value of order 1=0.021, p-value of order 2=0.219. Standard errors adjusted for 69 clusters.
Model (C): Variables transformed into first differences. 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 one-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. Wald test of jointly zero coefficients χ^2 (16)=15503.7. p-value<0.001. Arellano-Bond test for zero autocorrelation: p-value of order 1<0.001, p-value of order 2=0.125. Standard errors adjusted for 69 clusters.
Model (D): Standard errors bootstrapped with 100 replications. Correction procedure is based on Bruno (2005a, b).

Source: Authors' calculations.

Our interest focuses on the evidence concerning lagged adjustment and the effects of policy measures. All models consistently show that labour adjustment is sluggish, with a highly significant coefficient of adjustment. However, the reported levels differ considerably. While the LSDV result must be assessed with caution due to the inconsistency of the estimator, model (C) reports a considerably higher value than (B). This is in line with the finding of Blundell and Bond (1998) that the Arellano-Bond estimator may be downward biased. The Blundell-Bond results are close to the corrected LSDV results, which implies that the coefficient of adjustment is at 24 percent. This means that, after a shock, it takes about two and a half years to move halfway to the new steady state.[12] This level of the adjustment coefficient is in the range of values found in other studies on dynamic labour adjustment, such as in Stefanou et al. (1992) for German family farms and Luh and Stefanou (1996) for US farms.

With regard to policy effects, a first conclusion to be drawn is that most measures had no significant impact on agricultural employment at all. There is weak evidence that processing and marketing aid reduces employment. However, the GMM models do not support this finding and also in the corrected LSDV the parameter does not pass the ten percent level of significance. A bit stronger is the evidence on positive employment effects of investment support, which is significant at five percent in the LSDV and the Blundell-Bond models. According to the latter, one million euro of investment aid per region creates almost 24 jobs in agriculture in the short run. For this short run effect, 42 thousand euro annually are required to create one additional job. Given the logic of our model, full 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 model (C), the long run effect is 125 jobs in the steady state per one additional million euro of investment aid paid now. However, this result is not borne out in the corrected LSDV model.

A result consistently supported by models (B) to (D) is that there were significant employment losses in the years 2005 and 2006. Our interpretation is that this is due to the introduction of the decoupled SFP. It is unlikely that a severe macro effect, such as price drop, caused this fall in employment, as the revenue and profit situation of farms in Germany had notably improved in 2005 compared to previous years (BMELV 2007, 17).[13] According to our view,

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

13 The drop in value added in 2005/6 given in Figure 1 is misleading in this respect, as it is due to the fact that direct payments were no longer counted as farm revenue after decoupling in 2005.

decoupling broke the link between payments and labour allocations necessary for the maintenance of certain farm activities. The estimates range from 72 to 152 agricultural jobs per region that were lost due to decoupling in the short run (Table 4). Mean employment in the first sector per region was 1893 persons in the observed period, so that the SFP on average implied losses of 4 to 8 percent of jobs in agriculture. In the long run, these effects are even more dramatic, ranging from 276 (model B) to 637 (model C) job losses due to decoupling, or up to one third of all jobs in agriculture.

Table 4: Average short- and long-run job losses per region due to the 2005/6 effect

	Arellano-Bond (B)	*Blundell-Bond (C)*	*Corrected LSDV (D)*
Short-run losses (persons)	-152	-121	-72
Long-run losses (persons)	-276	-637	-298

Source: Authors' calculations.

Models (A) and (B) also produced significant parameters on population density and the overall wage level. The negative signs imply that fewer people work on farm in regions which are more densely populated and where wages are higher, which is a plausible result.

6 Conclusions

Our regression analysis of CAP payments in three German States reveals that there were few desirable effects on job maintenance or job creation in agriculture. The results are based on four specifications of a dynamic employment equation with fixed effects estimated on county level data. The specifications differ in how they eliminate the fixed effects and how they deal with the endogeneity of the lagged dependent variable. We found that agricultural employment adjusts slowly to shocks. On average, it takes about two and a half years to move halfway to the new steady state. Direct payments for crops and livestock, measures for the development of rural areas, transfers to less favoured areas and agro-environmental measures had no employment effect at all. Processing and marketing support to the downstream sector implied job losses in one of the four specifications. There was a significant loss of jobs in agriculture in 2005 and 2006. We suspect that this may have been a consequence of decoupling. Due to this effect, in the long run, up to one third of agricultural jobs per region are at stake.

The evidence presented here suggests that the only way to actively promote job creation in the CAP framework is via capital subsidies. These subsidies are mostly used to finance buildings or machinery. Apparently these increases in capital use were sufficiently complementary to labour that they induced relatively higher employment, that is they slowed down labour cuts. According to our estimates, 42 thousand euro of subsidies are required annually to create one additional job in the short run. However, capital subsidies are more effective in the long run, as they also affect the steady state equilibrium labour demand. Furthermore, it should be stressed that this finding is supported by only two of the four econometric specifications.[14]

We therefore conclude that, in the three East German States, the CAP mostly misses its target of safeguarding jobs. Potentially positive effects due to capital subsidies were counteracted by the recent decoupling of direct payments. Given the policy perspective to "modulate" further funds away from direct payments, this analysis calls into question whether an expansion of second pillar measures is a reasonable way to use the modulated funds.

The analysis here has focused on the goal of job creation in agriculture. With regard to other goals that may have been achieved by the CAP, such as environmental stewardship or the social goal of income redistribution, we can only conclude that their potential achievement at least has not made jobs in agriculture safer.

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[14] The finding is consistent with evaluation results of CAP investment support mandated by the European Commission (Agra CEAS, 2005, vi).

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