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Agricultural policy regime change assessment: Austrian accession to the European union

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Abstract Policy regime change evaluation involves assessing what would have been had the policy not changed. In this paper, we empirically assess the impact on Austrian pig producers of the 1995 decision of Austria to join the European Union. Applying a recently developed Hausman–Wu-Type cointegration test, we confirm the existence of a cointegrating relationship between Austrian and EU pig prices series. This relationship is used to forecast the counterfactual time path of prices. Within an expected utility framework, we compute the Austrian producer's willingness to pay to remain under the pre-accession policy. Accounting for the dual income and insurance effects, we found producers to have been under-compensated. Conventional welfare measures which do not include the insurance component would significantly underestimate the total welfare impact.

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1 Introduction

Economic integration can dramatically change the policy environment of accessional countries. In the case of a customs union, decisions to enter the European Union (EU) can have important economic consequences for the agricultural sectors of the accession country. Recognizing this, as a country transitions from its own policy regime to that of the EUs Common Agricultural Policy (CAP), special programs are often put in place to ease adjustment costs for farmers. The size of these costs is typically the subject of active public debate. Thus, the need for an accurate a priori cost assessment is clear; however, generating credible evaluations of the likely outcome can present a formidable challenge to the analyst. In this paper, we propose an approach to generate reliable estimates of the anticipated effects of policy regime change on farmers. We illustrate the approach by an analysis of the economic impact of Austria's decision to enter the EU on the Austrian slaughter pig sector in 1995.

Agricultural price and income policies are generally aimed at supporting producer income. However, public policies may yield an additional insurance component. We argue that the latter is an important component of a policy regime change assessment which heretofore has not been taken into account in the existing literature. The extent to which the income and insurance components can be expected to be altered by policy change is fundamental to the evaluation of regime-induced consequences. Recognizing this, Hennessy (1998) examined the wealth and insurance effects of policy on optimal production decisions. He argued that studies of trade and domestic policy reform in stochastic environments should consider both wealth and insurance effects. Thompson et al. (2004) extended this conceptualization to measure trade protection rates. Following Hennessy's call, in this paper, we empirically assess the impact of a one-time deterministic policy regime change on the level and variability of producer income.

Prior to accession, an important Austrian policy instrument was price supports for the slaughter pig sector. Prices were higher and relatively more stable than EU prices. However, with accession prices fell to EU levels. A major concern of Austrian pig producers about EU membership was that prices would fall to unprofitable levels. To ease the transition to this new situation, farmers were granted "digressive payments" distributed over the 4 years following entrance.[1] Despite compensation, under the umbrella of the CAP, there was widespread concern that the well being of the Austrian slaughter pig industry was at stake. As a consequence, a number of ex ante predictions were made to show the likely effects of the 1995 decision. For instance, Schneider (1994) expected pig prices to decrease an average of 23%. This decrease was expected to be compensated under the EUs CAP with digressive adjustment premia leading to an expected 17% decrease in revenue.

[1] Total payments were 5.82, 3.79, 2.33, and 0.87 (million €) from 1995 to 1999. While large in magnitude, even in 1995, payments were just 7.3% of market price.

While price declines were anticipated, the effect on farm income was the bottom-line concern. Income levels are also affected by changes in input costs, other payments, and transfers. However, as we argue here, there is an additional cost to the risk averse farmer: stability of income is preferred to instability. Therefore, if incomes were to become more volatile after accession, an added cost would be borne by the Austrian farmer due to the policy change. In other words, this may be viewed as an unanticipated cost (or benefit with more stable incomes) to the policy regime change. As a result of moving from the pre-1995 Austrian agricultural policy to the post-1995 CAP umbrella, are Austrian farmers better or worse off under the CAP? We present an analytical framework which enables the quantification of this willingness to pay. We begin with an assessment of the degree to which the Austrian and EU pig markets are economically integrated; we test for the existence of unit roots and a cointegrated relationship between Austrian and EU product prices. We introduce and make application of a new statistical test of cointegration for dynamic ordinary least squares (DOLS). Secondly, within an expected-utility-of-income framework, we use our estimated price relationships to compute what Austrian producers would have been willing to pay to remain under the pre-accession policy regime. This exercise provides the dual income-insurance components of the Austrian accession. We end with some concluding remarks.

2 Market integration

Market integration is assessed relative to the economic concept of spatial market equilibrium. Based on the law of one price (LOP), we test for the existence of a long-run stable relationship among Austrian and EU prices.[2] If we cannot reject LOP, the relationship can be used to generate reliable price forecasts. To begin, we individually examine the time series properties of Austrian and EU market price series during the period before the known point of policy change. The inability to reject nonstationarity in both series enables us to test for a potential cointegration relationship. We use the Stock and Watson (1993) Dynamic OLS estimator of the cointegrating regression coefficient and apply a Hausman–Wu-Type cointegration test. Developed by Choi et al. (2008), the test uses a GLS correction procedure to assess the robustness of the DOLS parameters. Equipped with a reliable cointegrating relationship, we proceed to forecast the time path of prices “with and without” the regime change; the difference between the forecasts is attributed to policy.

Monthly Austrian and European Union slaughter pig price data from January 1984 to October 2002 are displayed in Fig. 1. There are 226 observations; 132 before accession and 94 afterward. These data suggest that Austrian prices were clearly higher and less volatile than EU prices before January 1995 than afterward.[3] Further, it appears

[2] Prior to the MacSharry Reforms of 1992, the CAP was viewed as a unified system of administered prices. However, the reduction in price supports post-MacSharry followed by a decoupling of payments led to greater market integration. Especially for pigs, the formation of price is very much integrated with the emergence of Northern Germany as the single key market.

[3] Food prices, in general, were about 30% higher in Austria than in the EU.

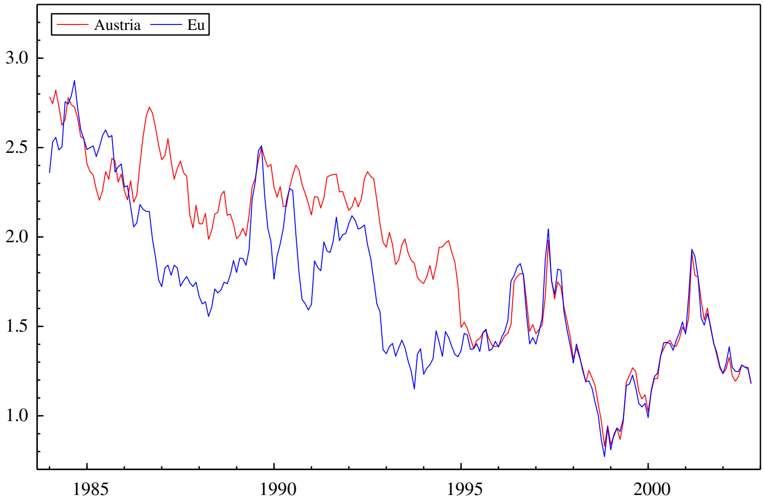

Fig. 1 Austrian and EU slaughter pig prices (€/kg)

Table 1 Austrian and EU slaughter pig prices (€/kg): mean and coefficient of variation

		1984:01–1994:12	1995:01–2002:10	Change (%)
Austrian price				
	Mean	2.243	1.379	−38.527
	CV	0.113	0.177	56.217
EU price				
	Mean	1.927	1.377	−28.565
	CV	0.214	0.193	−9.816

$\mathrm{CV} = \frac{\mathrm{SD}}{\mathrm{mean}}$

that the co-movement of the paired prices was not highly correlated before 1995 but became near perfect after Austria entered the EU. These observations are supported by the descriptive statistics for the two sub-periods shown in Table 1. Relative to pre-accession, during the post-1995 period average Austrian price levels fell by 38.5% while EU price levels fell 28.5%. A comparison of the coefficients of variation (CV)[4] between the two series reveals stark differences. While the post-1995 CVs are nearly identical, the CV changes from the pre-1995 period differ greatly; Austria's CV increased by 56% while that of the EU fell by nearly 10%. Both Fig. 1 and the descriptive statistics in Table 1 suggest a clear structural break the Austrian pig price series at the time of Austria's accession to the EU (Fig. 2).

[4] Compared to the standard deviation (SD), the coefficient of variation (CV) is attractive when the data are in different units or when data are in same units but the means are apart. Since both for EU and Austria, there are large changes in the mean prices before and after EU-accession, SD would be a misleading indicator of price uncertainty.

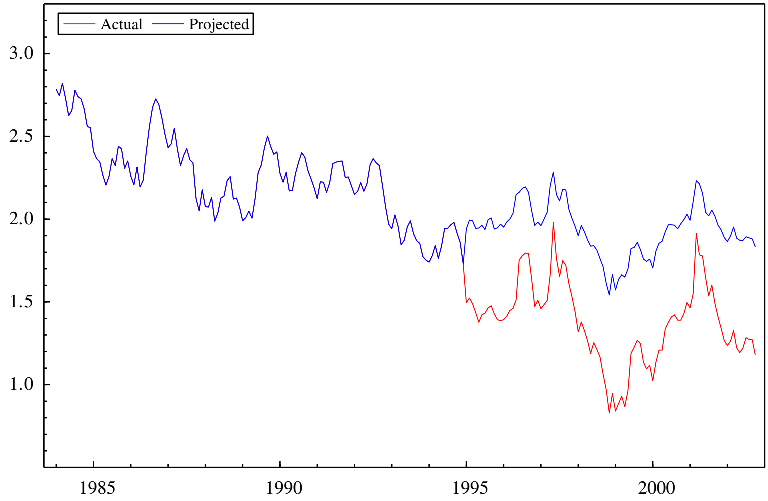

Fig. 2 Austrian pig prices: actual versus projected (€/kg)

3 Cointegration and VECM estimation

Using pre-accession data, we estimate the cointegrating regression,

$$p_t^a = \alpha + \beta p_t^e + \varepsilon_t, \tag{1}$$

where p_t^a is a logarithm of Austrian pig price and p_t^e is a logarithm of EU pig price, both in Euros. Note that this cointegration relationship might be interpreted as "weak" form of the LOP.[5] First, we test the hypothesis of unit root for both Austrian and EU pig prices only for the pre-Accession period of 1984:01–1994:12 by using ADF, PP, and DF-GLS (Elliott et al. 1996) tests. Results are presented in Table 2. In general, all three tests cannot reject the null hypothesis of unit root nonstationarity.

In order to test whether a cointegrating relationship exists, we first use the residual-based cointegration test. As presented in Table 3, the residual-based cointegration test does not reject the null of no cointegration at 10% level of significance implying Eq. (1) might not be a cointegrating regression. However, since it is well known that the ADF test suffers from low power problem, it would be too hasty to make a judgement based on the residual-based cointegration test result only. Therefore, we employ the Hausman–Wu-Type cointegration test recently proposed by Choi et al. (2008). Unlike the residual-based cointegration test where the null hypothesis is no cointegration, the null hypothesis of the Hausman–Wu-Type cointegration test is cointegration. As argued in Ogaki and Park (1998), it is desirable to test the null hypothesis of cointegration rather than that of no cointegration in many applications where economic models imply cointegration. In our application, the cointegration relationship between Austrian pig price and EU pig price is implied by the law of one price. If we are unable

[5] LOP holds only when $\beta = 1$. However, if there are some frictions, such as a tariff and a subsidy in one country, β might not be exactly equal to one. We can even interpret the difference of β from one as policy effects.

Table 2 Unit root tests for pig prices

Lag	Austria			EU		
	ADF	PP	DF-GLS	ADF	PP	DF-GLS
0	−2.667	−2.667	−2.504	−2.193	−2.193	−2.151
1	−3.218	−2.921	−3.022	−2.610	−2.394	−2.610
2	−3.025	−2.958	−2.741	−2.439	−2.447	−2.439
3	−4.012	−3.110	−3.730	−3.026	−2.553	−3.014
4	−2.669	−3.108	−2.504	−2.236	−2.549	−2.224
5	−2.897	−3.045	−2.681	−2.007	−2.484	−2.066
6	−2.684	−3.000	−2.287	−2.092	−2.449	−2.149
7	−2.121	−2.878	−1.837	−1.810	−2.386	−1.880
8	−2.066	−2.752	−1.763	−1.976	−2.335	−2.058
9	−1.990	−2.676	−1.774	−2.614	−2.359	−2.677
10	−1.670	−2.585	−1.546	−2.627	−2.394	−2.663

Testing regression includes the constant and the linear time trend. Critical values are −3.52, −4.07 and −4.36 for 10, 5 and 1% level of significance

Table 3 Dynamic OLS and Hausman–Wu-Type Cointegration Test

	$\hat{\alpha}$		$\hat{\beta}$
Dynamic OLS	0.537		0.404
	(0.036)		(0.054)
Dynamic GLS	–		0.369
	(–)		(0.167)
Hausman–Wu-Type cointegration test[a]		0.044	
Residual-based cointegration test[b]		−2.511	

Lead and lag length is set to 4. Numbers in parenthesis are HAC standard error estimates with Bartlett kernel of bandwidth parameter 4

[a] Critical values of $\chi(1)$ are 2.71, 3.84 and 6.63 for 10, 5 and 1% level of significance

[b] Lag length of 11 is selected by BIC method. Critical values of residual-based cointegration test are −3.13, −3.41 and −3.96 for 10, 5 and 1% level of significance

to reject the null of cointegration, we can proceed to estimate Eq. (1) using a cointegration estimation method, such as Stock and Watson (1993) Dynamic OLS (DOLS).

The Hausman–Wu-Type cointegration test exploits the different properties of DOLS and DGLS estimators. DOLS and DGLS estimators of β in Eq. (1) are the OLS estimators of the following regression equations, respectively:

$$p_t^a = \alpha + \beta p_t^e + \sum_{j=-q}^{q} \varphi_j \Delta p_{t-j}^e + \varepsilon_t \tag{DOLS}$$

$$\Delta p_t^a = \beta \Delta p_t^e + \sum_{j=-q}^{q} \varphi_j \Delta^2 p_{t-j}^e + \Delta\varepsilon_t \tag{DGLS}$$

Under the null hypothesis of cointegration, the Hausman–Wu-Type cointegration test statistic follows the chi-square distribution:

$$\frac{\sqrt{n}(\widehat{\beta}_{\text{GLS}} - \widehat{\beta}_{\text{DOLS}})^2}{\text{Var}(\widehat{\beta}_{\text{GLS}})} \xrightarrow{d} \chi(1).$$

The DGLS estimator is consistent regardless of if the true data generating process (DGP) is a cointegrating regression or a spurious regression. However, the DOLS estimator is super-consistent (more efficient) when true DGP is a cointegrating regression, but it is not consistent when the true DGP is a spurious regression (Choi et al. 2008). DOLS estimation and Hausman–Wu-Type cointegration test results are provided in Table 3. Based on Hausman–Wu-Type cointegration test statistic, the null hypothesis of cointegration is not rejected at 10% significance level. Thus, we have confirmation of the existence of cointegrating relationship between Austrian and EU pig prices.

In our quest to use Eq. (1) to construct the Austrian pig price series after Austria's EU-accession on January 1995 under the hypothetical condition that Austria's EU-accession had not occurred, it is necessary to additionally confirm weak exogeneity of the EU pig price; that is, it should be Austrian pig price (not EU pig price) that adjusts to deviations from the long-run equilibrium. Following Thompson et al. (2002), we estimate the Vector Error Correction Model (VECM):

$$\Delta p_t^a = \mu_1 + \sum_{i=0}^{k} \theta_{1,i}^a \Delta p_{t-i}^a + \sum_{i=0}^{k} \theta_{1,i}^e \Delta p_t^e + \varphi_1 (p_{t-1}^a - \beta p_{t-1}^e) + \varepsilon_{1,t}$$

$$\Delta p_t^e = \mu_2 + \sum_{i=0}^{k} \theta_{2,i}^a \Delta p_{t-i}^a + \sum_{i=0}^{k} \theta_{2,i}^e \Delta p_t^e + \varphi_2 (p_{t-1}^a - \beta p_{t-1}^e) + \varepsilon_{2,t}.$$

The estimation results with $k = 3$ are presented in Table 4. The weak exogeneity of EU pig price implies that $\varphi_2 = 0$. Since the t-statistic of $\varphi_2 = 0$ is $-0.041/0.081 = -0.51$, the hypothesis of the weak exogeneity cannot be rejected. On the contrary, the

Table 4 Vector error correction model estimation

	Austria		EU	
μ	−0.002	(0.003)	−0.005	(0.004)
θ_1^a	0.196	(0.093)	−0.324	(0.148)
θ_2^a	−0.111	(0.097)	0.029	(0.154)
θ_3^a	0.184	(0.091)	−0.155	(0.144)
θ_1^e	0.132	(0.063)	0.260	(0.100)
θ_2^e	−0.092	(0.065)	−0.085	(0.102)
θ_3^e	0.172	(0.063)	0.198	(0.100)
φ	−0.159	(0.051)	−0.041	(0.081)

Numbers in parenthesis are standard error estimates coefficient for Austrian pig price, φ_1, has the correct negative sign and statistically significant (t-statistic of $\varphi_1 = 0$ is $-0.159/0.051 = -3.12$).[6]

With confirmation of both a cointegrated Austrian–EU price relationship and a weak exogeneity of EU prices, we can use the estimated cointegrating regression to construct a post-1995 Austrian price series. This series is constructed under the assumption that the Austrian accession had not occurred. More specifically, the projected counterfactual Austrian pig price from 1995:01 to 2002:10 is calculated by the following formula:

$$\overline{p}^{a}_{1995:01+j} = 0.537 + 0.404 \times p^{e}_{1995:01+j}$$

where 0.537 and 0.404 are DOLS estimates for the cointegrating regression and $p^{e}_{1995:01+j}$ is actual EU pig price.

Under the CAP, a variety of income supports become available to the Austrian farmer. As a result, the relevant metric of concern is income levels and stability, not product prices. Accordingly, we compute farmer incomes (gross margin) as: revenue less variable production costs and any additional payments or other transfers. Included in our definition of revenue are three revenue-enhancing payments: digressive (premiums to facilitate adjustment), hectarage (historical entitlements), and environmental (for reduced animal density). Variable productions costs include: purchased soymeal protein, farm grown CCM (corn cob mix), piglet cost, veterinary expenses, insurance, and other costs. We express all costs and revenues in real Euros.

We compute incomes for the specific case of the farmer producing pigs mainly on farm feed. A key reason for choosing the mainly on farm feed scenario is so we would not have to consider the complexities involved by adding the potential policy-induced change in the price of purchased feed. Budget cost data show that the differences between the two scenarios were roughly the same; thus, little income differences occurred between the pre- and post accession periods. Further, during our sample period some 40–50% of Austrian pig farms operated mainly on farm feed. This category includes organic producing farms which mainly produce all their own feed. Based on this scenario, we use our statistical model to compute forecasts of producer incomes. The pre-and post-accession revenues and incomes are shown in Fig. 3.

4 Producer welfare

We utilize an expected utility framework to approximate the policy-induced income and insurance effects. Following Newbery and Stiglitz (1981), the combined policy impacts of the level (transfer) and stability (risk) of income are assessed in producer welfare terms. The total welfare effect is the sum of the transfer and risk effects. The transfer effect is the percentage change in mean income due to the policy change and the risk effect is the welfare change attributable to the policy-induced income risk.

[6] These results are robust to lag length selection from $k = 2$–5. The weak exogeneity of EU pig price is also confirmed with monthly seasonal dummies.

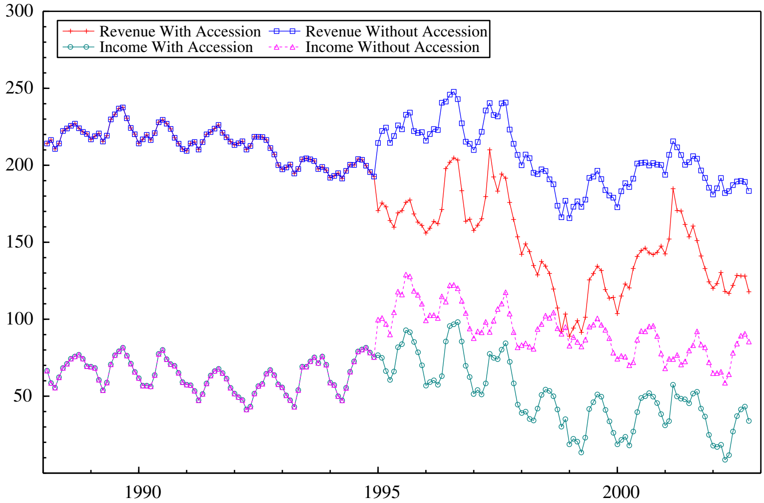

Fig. 3 Revenue and gross margin per slaughter pig: Austria, 1988–2002 (real €)

The objective is to determine the value of the policy benefit (B) needed to make the expected utility of income with accession equal to what it would have been without accession. The value of B is the willingness to pay to remain under the Austrian policy regime. For given any farmer's utility function $U(Y)$, this can be written as:

$$\mathrm{EU}(Y_0) = \mathrm{EU}(Y_1 - B) \tag{2}$$

where Y_0 is the income *with* accession, Y_1 is the income *without* accession, and B is the *absolute welfare benefit*.

To compute the *absolute welfare benefit* B, we need to know the specific functional form of the farmer's utility function. In this paper, we assume that the utility function is constant relative risk aversion (CRRA) as follows:

$$U(Y) = \frac{Y^{1-\gamma}}{1-\gamma}$$

where γ is the relative risk aversion coefficient; the greater the value of γ the more risk averse the farmer. For the reasonable values of γ, we draw upon the literature. Antle (1987) reports estimates of relative risk aversion coefficients for US agricultural producers in the range of 0.19–1.77. Since we do not have specific estimates for Austrian pig farmers, we provide result scenarios R values ranging from 0.0 to 2.0.

With CRRA utility function, we numerically solve for B to obtain our welfare measure. More specifically, we solve for B in:

$$\sum_{t=t_b}^{t_e} \frac{Y_{0,t}^{1-\gamma}}{1-\gamma} = \sum_{t=t_b}^{t_e} \frac{(Y_{1,t} - B)^{1-\gamma}}{1-\gamma}$$

Table 5 Welfare effects of Austrian EU accession for pig flattening

R	0.0	0.5	1.0	1.5	2.0
1995:01–1998:12					
B(€)	37.7	38.5	39.5	40.0	40.8
$B/\overline{Y}_0$ (%)	59.2	60.5	61.9	62.8	64.1
1995:01–1999:12					
B(€)	41.5	43.0	44.7	46.8	49.0
$B/\overline{Y}_0$ (%)	65.1	67.5	70.2	73.5	76.9
1995:01–2000:12					
B(€)	42.4	43.9	45.5	47.4	49.2
$B/\overline{Y}_0$ (%)	66.6	68.9	71.5	74.4	77.2
1995:01–2001:12					
B(€)	41.2	42.3	43.5	44.9	46.3
$B/\overline{Y}_0$ (%)	64.6	66.5	68.3	70.5	72.6
1995:01–2002:10					
B(€)	42.0	43.3	44.6	46.1	47.4
$B/\overline{Y}_0$ (%)	66.0	68.0	70.1	72.4	74.5

$\overline{Y}_0 = 63.70$€

where t_b and t_e are the beginning and ending period of welfare evaluation, respectively. t_b is always set to January 1995, however for the robustness check, we consider various t_e; December 1998, December 1999, December 2000, December 2001, and October 2002. Since our welfare evaluations are largely robust, we discuss below those results obtained from the longest time horizon, from January 1995 to October 2002.

The income, insurance and total welfare effects of the 1995 accession are shown in Table 5. We illustrate these effects for the farmer using "mainly on farm feed" and receives a hectarage premium (historical entitlement), digressive payments (premiums to facilitate adjustment), and base environmental payments (for reduced animal density). Positive sign of B indicates the cost to the farmer attributed to EU membership.

Given the pre-accession average income of 63.7€ per animal, the income (level) effect for the risk neutral farmer ($R = 0.0$) is 42.0€. This means that the average risk neutral farmer received 42.0€ less that would have been received if accession had not occurred. This implies that farmers would have been willing to pay 42.0€ per animal to have remained under the pre-accession policy regime. The number 66.0 means that farmers received on average 66% less during the post-accession period than would have been expected without accession. An insurance effect is observed as farmers become increasingly risk averse. If, for example, $R = 1.0$, the total welfare cost is 44.6€ per animal; 42.0€ is the income effect and 2.6€ is the insurance effect. The income effect is invariant to R but the cost due to increased income risk increases the more risk averse the farmer. For moderately risk averse farmers ($R = 2.0$) the insurance effect can exceed 10% of the total; a nontrivial portion. It is important to note that without accounting for the insurance effect, welfare losses can be significantly underestimated. Our policy regime change impact assessment tends to be larger than ex-ante predictions for at least two major reasons: first, we use a different statistical forecasting model, and second, we include the insurance component.

A commonly held view is that Austrian farmers received generous temporary compensation payments, especially during the first year following accession. In subsequent years, with the digression of payments this over-compensation soon evaporated. However, this view of generous compensation for overall agriculture should not be generalized to individual commodities. For instance, contrary to our findings for pig farmers, it is widely held that grain farmers were over-compensated with the newly introduced CAP direct payments. Over- or under-compensation must be evaluated on a commodity-by-commodity basis.

5 Concluding remarks

The impact of policy regime change necessitates an assessment of "what would have been the case" if policy had not changed. Accordingly, we investigate the effect on Austrian pig producer incomes of Austria's decision to entry to the European Union in 1995. First, this assessment requires an understanding of how policy influences the true data (price) generating process. Positing a weak form of LOP, we investigate the statistical relationship between Austrian and EU prices. Since tests revealed the two nonstationary series to be cointegrated, we proceeded to use the cointegrated relationship to forecast the time path of prices without regime change. We used a recently developed Hausman–Wu-Type test of cointegration for Dynamic Ordinary Least Squares. Our estimated relationship was used to forecast the time path of prices under the scenario of no regime change. These price forecasts drive producer incomes.

Second, within an expected utility maximization framework, we compute the producer's willingness-to-pay to remain under the pre-accession Austrian policy regime. Equipped with reliable income forecasts, we compute the income and insurance effects of regime change.

Our results suggest that farmers were under-compensated per unit of production by roughly 60–70%; 90% due to income level and 10% due to added income risk. Conventional welfare measures which do not include the insurance component would significantly underestimate the total welfare impact.

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