

The Political Economy of Policy Instrument Choice: Theory and Evidence from Agricultural and Food Policies

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Abstract

We study the political economy of instrument choice with an application to agricultural and food policies. We present stylized facts on the choice of policy instruments and develop a political economy theory of instrument choice. The key predictions of the model suggest a rational explanation of instrument choice patterns, based on the trade-off between transaction costs and distortions of the policies, and internal and external political constraints. Our empirical analysis supports the main predictions of the theoretical model. The shift from distortionary to less distortionary instruments is positively influenced by institutional development, a country's net trade status, and the GATT/WTO framework.

Keywords

Public Policy, Political Economy, Instrument Choice, Agricultural and Food

1. Introduction

An extensive literature on the political economy of agricultural policies has developed over the past 20 years. Papers in this literature have attempted to provide an explanation for the stylized facts on agricultural protection, such as the widely observed increase in agricultural protection when an economy grows—see de Gorter and Swinnen (2002) [1], Swinnen (2010) [2] and Anderson *et al.* (2013) [3] for reviews. Studies have attempted to provide an explanation for the stylized facts on agricultural protection, such as the widely observed increase in

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agricultural protection when an economy grows (Anderson and Hayami, 1986) [4]. Theoretical studies attempting to explain these and other facts have stressed the implications of organization costs on the political decision-making process (Gardner, 1987; Olson, 1990) [5] [6], structural factors affecting the distributional effects of agricultural protection (Anderson and Tyers, 1988; Honma and Hayami, 1986; Swinnen, 1994) [7]-[9], the relative income position of agriculture (Bullock, 1992; de Gorter and Tsur, 1991; Swinnen and de Gorter, 1993) [10]-[12] and political institutions like differences in electoral rules and the degree of democracy (Beghin and Kerallah, 1994; Swinnen *et al.*, 2001; Olper, 2001; 2007; Olper and Raimondi, 2010) [13]-[17].

These political economy studies of agricultural policy have focused primarily on explaining the *level* of policy intervention and less attention is paid to the explanation of the *instruments* used for intervention (de Gorter and Swinnen, 2002) [2]. This bias in focus is an important shortcoming of the literature. From a welfare perspective the key question should be why governments have introduced so many market distortions through agricultural policies. The distortionary effects of government interventions are equally dependent on the choice of the instrument as on the level of the intervention. Therefore the choice of instrument should be at least of equal concern as the intervention level. In fact, both the trade and agricultural policy literature and the policy debates have reflected this importance (e.g. Gardner 1983) [18]. In the policy world, the debate on the choice of instruments has been a very important element of policy discussions. The differences in distortionary effects is recognized by the WTO in the classification of agricultural policy instruments in green, blue and amber boxes—with the green box for non-trade distorting policies instruments. This distinction between the level of support and the extent of market and trade distortions is at the heart of some important policy reforms, such as those of the EU's Common Agricultural Policy (CAP) over the past two decades. In fact, one could argue that the issue of instrument choice was the key element of the CAP reform, more so than the level of support (Swinnen, 2008) [19].

Surprisingly, this attention to instrument choice in the literature and the policy debate has not translated in similar attention in studies on the political economy of agricultural and food policies where most of the focus has been on explaining the level of intervention rather than its form. A possible explanation for this bias in focus is differences in the availability of good empirical data, resulting in some well-known and puzzling stylized facts on policy level but not on instruments.¹

This paper contributed to filling this gap, first of all presenting such stylized facts, drawing on OECD data. The paper then proceeds to provide an explanation of these stylized facts.

The paper is organized as follows. In the next section we present stylized facts on instrument choice in agricultural and food policies. Then we develop a theoretical model and derive some key hypotheses. Afterwards we empirically test these hypotheses using an econometric study. The final section concludes.

2. Instrument Choice in Agricultural Policy

We first present some stylized facts on agricultural and food policy instrument choice in OECD countries over the past 25 years.²

Since 1986 the OECD calculates policy support given to agriculture. The total amount of support to agriculture is referred to as Producer Support Estimate (*PSE*).³ The *PSE* data cover 28 countries, 12 of which are not OECD members, over the period 1986-2009 (Table A1). The OECD's calculation of policy support distinguishes between several instruments. For the purpose of our analysis it is convenient to combine the instruments into "market price support" (*mps*), "input subsidies" (*is*) and "direct payments" (*dp*). Their share in total support (*PSE*) is represented by *mpsh*, *ish*, and *dpsh*, respectively.

The first instrument, *mps*, includes all transfers through tariffs, price support and subsidies directly linked to agricultural production. These instruments are typically considered as being the most distortive. The second instrument, *is*, are input subsidies and cover a very heterogeneous set of measures, spanning from investment aids

¹This argument may apply more widely, while there have been some studies in the general literature on explaining instrument choice, in particular why governments chose inefficient policies to redistribute income or protect certain groups (e.g. Cassing and Hillman, 1985; Rodrik, 1986; Coate and Morris, 1995; Acemoglu and Robinson, 2001) [20]-[23], these studies are almost exclusively theoretical. Only recently have there been a few empirical studies on the determinants of instrument choice, including Kono (2006) [24] and Ederington and Minier (2006) [25].

²For stylized facts on the historical evolution of agricultural policy instruments, see e.g. Tracy (1989) [24], Josling (2007) [25], Swinnen (2009) [26]; on the historical evolution of trade policy, see e.g. Irwin (2008) [27]; and Williamson (2003) [28].

³Initially the *PSE* calculations were only for OECD member states but more recently also some other countries, such as China and Brazil, are covered. For countries not belonging to the OECD, the time coverage is not complete: the first year observation is around 1990-1992 and the last is 2007.

and labor subsidies to land protection programs. Finally, the third instrument, *dp*, includes fully decoupled and partially decoupled agricultural payments. These instruments are generally considered the least distortive.

In the 1980s, the most important instrument was *mps*. The share of market price support in total support was 82%, whereas the share of direct payments made up only 9%, and of inputs subsidies 9%. In the next two decades the share of market price support has declined and that of direct payments increased substantially (**Table 1**). By the late 2000s the former had decreased to 49% and the later increased to 38%. In contrast, the share of input subsidies remained about the same.

The choice of policy instrument is correlated with 1) the level of development; 2) the trade status, and 3) the GATT Uruguay Round Agreement on Agriculture (URAA). **Figure 1** (left panel) illustrates a positive correlation between economic development and the use of direct payments and **Figure 1** (right panel) shows a negative correlation with the export share. In addition, **Figure 2** indicates that the shift from market price support to direct payments started in the early 1990s, which was the time of the conclusion of the URAA and has continued during the Doha WTO negotiations.

GATT/WTO rules distinguish between instruments according to their distortionary impact and limit the use of distorting measures while non-distorting measures are not regulated. More specifically, the WTO classifies agricultural policy instruments in green, blue and amber boxes—with the green box for non-trade distorting policies instruments (see Josling and Tangermann, 1999; Tangermann, 1999; Josling, 2000 for more details) [29]-[31].

In summary, these empirical indications suggest that the choice of instruments is non-random. As stylized facts, we find that the choice of instruments is correlated with three factors: 1) a country’s level of development; 2) the URAA and the Doha WTO negotiations; and 3) a country’s trade status. We now develop a theoretical model to explain these stylized observations.

3. Theory

We use the same static framework as most models in the literature and consider the choice of governments

Table 1. Support by policy instrument based on OECD PSE database (Million US \$).

	Average 1986-1988		Average 2007-2009	
	Value	Share	Value	Share
Market price support	195,839	0.82	125,215	0.49
Input subsidies	20,400	0.09	33,403	0.13
Direct payments	22,425	0.09	98,146	0.38
Total PSE	238,665	1.00	256,764	1.00
Percentage PSE	37		22	

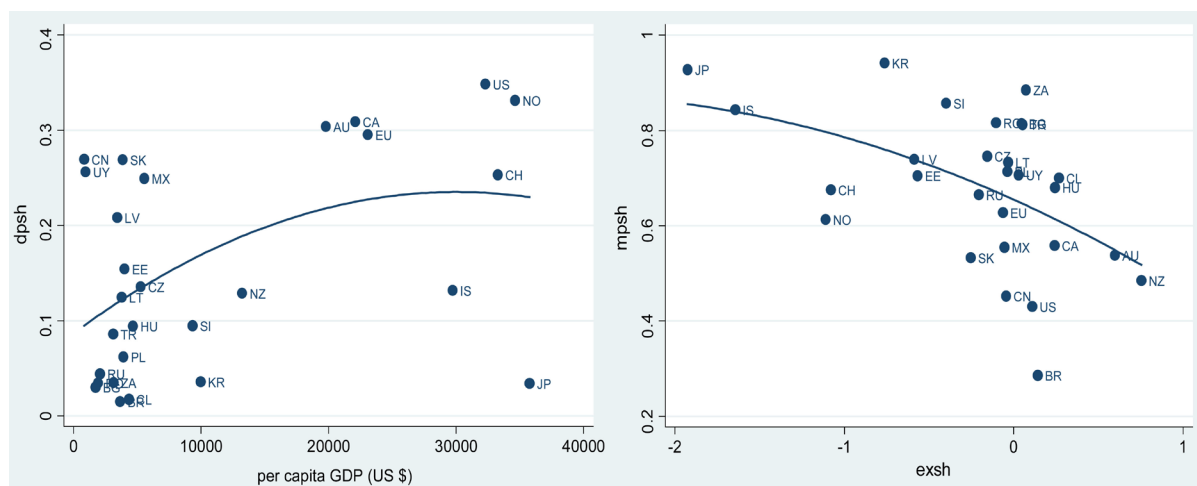


Figure 1. Relation between the share of decoupled (*dpsh*) and coupled (*mpsh*) transfers with the economic development (left panel) and the trade export share (right panel) in OECD.

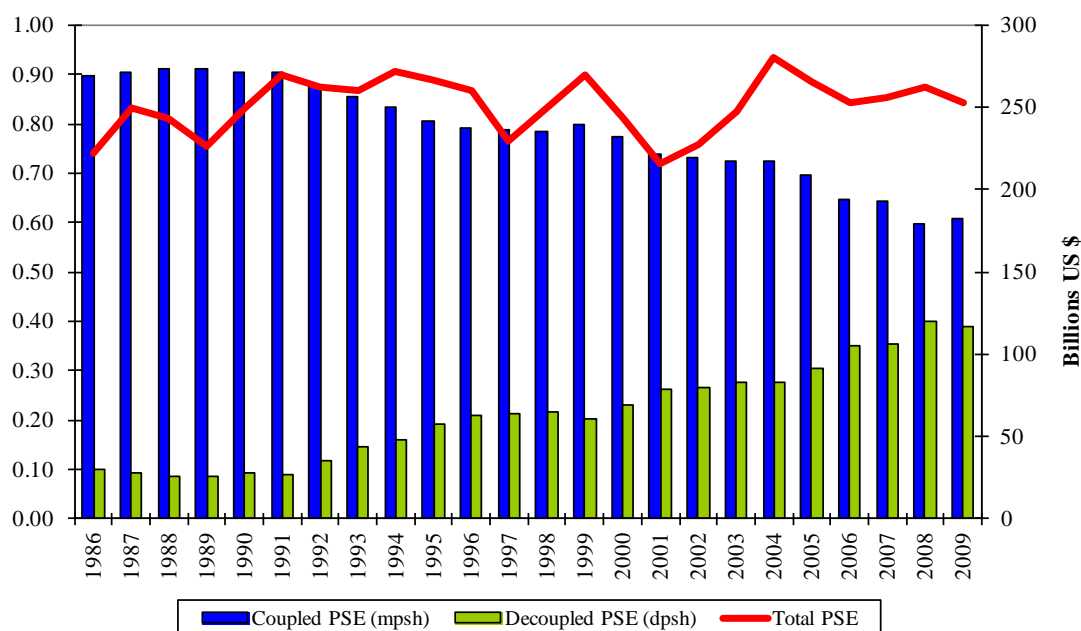


Figure 2. Evolution of total PSE (billions of US \$) and the shares of coupled and decoupled transfers in OECD. (Source: own computation on PSE/CSE database).

between instruments in the absence of existing policies (see, e.g., Hillman and Ursprung, 1988; Foster and Rausser, 1993; Kono, 2006) [32]-[34]. We assume that governments have perfect information on the impact of the various policy instruments, so there is no room for policy obfuscation. Consider that for some reason, e.g. a dramatic decline in world market prices for agricultural products, the government introduces policies to support producers' incomes.

We assume that the government has two different policy instruments at its disposal (see Hillman and Ursprung, 1988; Rodrik, 1986; Coate and Morris, 1995) [21] [22] [32] to transfer income to producers: instruments t and s , which are assumed to have the following characteristics:

	<i>Distortions</i>	<i>Transaction costs</i>	<i>Impact on government revenue</i>
t	High	Low	Positive if net importing; Negative if net exporting
s	Low	High	Strongly negative

Policy t has low transaction costs but high costs of market distortions, and has a positive (negative) impact on government revenue if the country is a net importer (exporter). Policy s causes fewer distortions but is characterized by high transaction costs, and has a strongly negative impact on government revenue, independent of the country's trade status. Even if the country is a net exporter, the impact of instrument s on government revenue is more negative than instrument t 's impact. One could think of tariffs vis-à-vis lump-sum transfers, or market price support vis-à-vis direct income support as examples of policies t and s , respectively.

As in Kono (2006) [34], we assume that governments need both voter support and money to stay in power. Money can be raised both through interest-group contributions and through revenues from the implementation of policy instruments. Our assumptions imply a modified Grossman and Helpman (1994) [35] model of government decision-making where, in line with Maggi and Rodríguez-Clare (2000) [36], the government maximizes a weighted sum of interest group contributions, policy revenues, and total voter support:

$$G(t, s) = C(t, s) + \omega^R(\alpha)R(t, s; \beta) + \omega^V V(t, s), \quad (1)$$

where G is government utility, C are the interest-group contributions, R measures the budgetary costs or revenues of the policy instruments, and V is total voter support. t and s are the income transfers of the two policy instruments, and ω^R and ω^V are the weights that the government gives to respectively revenue considerations

and total voter support. β represents the trade balance of that country, and α is an inverse measure for a country's institutional development.

In developing countries—with relatively underdeveloped institutions—raising revenue through foreign trade taxes constitutes the single largest source of public revenue (Burgess and Stern, 1993; Rodrik, 1995; Bates and Block, 2010) [37]-[39]. The revenue motive is therefore more imperative in countries with less developed institutions. Since α inversely measures the country's institutional development, we assume that the weight attached by the government to the revenue function increases with less developed institutions $\left(\frac{\partial \omega^R}{\partial \alpha} > 0\right)$.

As in Grossman and Helpman (1994) [35], we assume that the interest group consists of active lobbyists that solicit income transfers from the government. For this purpose the interest group offers the government a schedule that lists the interest group's contributions as a function of the income transfers. The interest-group contributions $C(t, s)$ rise with the level of the income transfers ($C_t > 0; C_s > 0$), but at a decreasing rate ($C_{tt} < 0; C_{ss} < 0$).⁴

The policy revenue function $R(t, s; \beta)$, is assumed to be decreasing in policy instrument s ($R_s < 0$), whereas the revenue impact of instrument t can be either positive ($R_t > 0$) or negative ($R_t < 0$), depending on the trade status of the country (respectively net importing or net exporting). We assume that ($R_s < R_t$) to represent that instrument s has a highly negative impact on government revenue, even more negative than instrument t in the case of a net-exporting country. $b^t \alpha t$ is concave in the income transfers ($R_{tt} < 0; R_{ss} < 0; R_{ts} < 0$). In line with the marginal impact of instrument t being respectively positive and negative for a net-importing and net-exporting country, the impact of an increase in the trade balance β , on instrument t 's marginal revenue impact is negative, *i.e.* $R_{t\beta} < 0$. The negative revenue impact of instrument t is independent of the trade balance: $R_{s\beta} = 0$.

The function for total voter support, $V(t, s)$, is given by

$$V(t, s) = W(t, s) - b^t \alpha t - b^s \alpha s, \quad (2)$$

where the first term, $W(t, s)$, represents total voter welfare, and the second and third terms, $b^t \alpha t$ and $b^s \alpha s$, measure the total transaction costs related to each instrument. As before, α is an inverse measure for the country's institutional development. In an unfavourable institutional environment where institutions are underdeveloped and the administrative capacity is low, transaction costs are higher for the same amount of income transfer (Burgess and Stern, 1993) [37]. Since policy t involves lower transaction costs than policy s for the same amount of income transfer, we assume that $b^s > b^t \geq 0$. Hence $b^t \alpha t$ and $b^s \alpha s$ are the average transaction costs per unit of the income transfers t and s .

As both policy instruments are distortionary measures, $W(t, s)$ is decreasing in the income transfers ($W_t < 0; W_s < 0$) at an increasing rate ($W_{tt} < 0; W_{ss} < 0; W_{ts} < 0$). Instrument t is more distorting than instrument s , so $W_t < W_s$.

The equilibrium pair of income transfers (t^*, s^*) is determined by the first order conditions (FOCs):⁵

$$\begin{cases} G_t = C_t(t, s) + \omega^R(\alpha)R_t(t, s; \beta) + \omega^V W_t(t, s) - \omega^V b^t \alpha = 0 \\ G_s = C_s(t, s) + \omega^R(\alpha)R_s(t, s; \beta) + \omega^V W_s(t, s) - \omega^V b^s \alpha = 0 \end{cases} \quad (3)$$

Define $h = \frac{t}{t+s}$ as the share of policy t in the total income transfer. Comparative statics on the equilibrium

$h^* = \frac{t^*}{t^* + s^*}$ for changes in the institutional development of a country, α , and the country's trade balance, β , yield the following two results.

⁴Subscripts t and s denote partial derivatives.

⁵The Hessian matrix of the government's objective function is $H(G) = \begin{pmatrix} G_{tt} & G_{ts} \\ G_{ts} & G_{ss} \end{pmatrix}$. In order to obtain a global maximum and to perform comparative statics, this matrix must be negative definite. Since all the Hessian's elements are negative, *i.e.* $G_{tt} < 0$, $G_{ss} < 0$, and $G_{ts} < 0$, the matrix is negative definite if $\det(H) = G_{tt}G_{ss} - G_{ts}^2 > 0$ (Winston, 2004). To secure uniqueness of the equilibrium and reaction function stability, in line with Brander and Spencer (1983) [40] and Dixit (1984) [41], we assume that the own effects of the income transfers on marginal contributions, revenue, and total voter welfare exceed cross effects such that $G_{tt} < G_{ss}$ and $G_{ss} < G_{tt}$.

Result 1: If $\frac{b^s}{b^t} > \frac{G_{ss}}{G_{ts}} > 1$ and $\frac{R_t}{R_s} > \frac{G_{ts}}{G_{ss}} < 1$: $dh^*/d\alpha > 0$.

Proof: See Appendix 1.

Result 1 implies that in countries with lower institutional capacity, where policy transaction costs are higher and the revenue motive is more important, *ceteris paribus*, the relative share of income transfer t^* is higher in equilibrium. Hence countries with less developed institutions (α higher) will apply relatively more distorting policies (h^* larger), provided that the transaction costs of the more distorting policy are sufficiently lower than that of the other policy $\frac{b^s}{b^t} > \frac{G_{ss}}{G_{ts}} > 1$, and that the less distorting policy has a sufficiently more negative impact on government revenue $\frac{R_t}{R_s} > \frac{G_{ts}}{G_{ss}} < 1$. The latter condition is always fulfilled if the more distorting policy has a positive impact on government revenue ($R_t > 0$).

To illustrate this result, take the specific case of a net-importing country. In that case, $R_t > 0$, and the second condition is fulfilled. If in addition the more distorting instrument (t) involves no transaction costs, *i.e.* $b^t = 0$, the first condition holds as well. It is clear from **Result 1** and the proof in Appendix 1 that in this specific case, an increase of the institutional capacity of a country (α lower) will always result in a higher relative share of the less distorting policy instrument s^* in equilibrium (h^* smaller). The result also holds under less strict conditions, which are discussed in Appendix 1.

Result 2: $dh^*/d\beta < 0$.

Proof: See Appendix 1.

Result 2 implies that if the trade balance of a country increases (β increase), the relative share of the more distorting policy in the total income transfer decreases (h^* decreases). For example, if for some exogenous reason a country's imports decrease, *ceteris paribus*, the country will shift to using the less distorting policy relatively more, although it involves relatively higher transaction costs.

4. Empirical Evidence

To formally test whether our theoretical hypotheses are consistent with the observed evidence on instrument choices we use the share of market price support in total support (*mpsh*) as a proxy for the instrument t and the share of direct payments in total support (*dpsh*) as a proxy for the instrument s . As we explained in Section 2, the OECD data on instrument choice cover 28 countries over the 1986-2009 period.⁶

We proxy the institutional development and administrative capacity of a country by the real GDP per capita (*gdppc*), taken from the World Development Indicators. As an indicator of the trade status we use the net export share in total production (*exsh*), based on FAO data.⁷ To capture the effect of international agreement we include a dummy variable, *d_gatt*. This dummy takes the value of 1 since 1995 (0 otherwise). 1995 was the first year of the GATT URAA implementation, which has introduced more constraints on the use of highly distortionary policy instruments like *mps*, than on lower distortionary instruments, like *dp*. In fact, fully decoupled policies which are not trade distorting are allowed under WTO principles.

As control variable and to account for path dependency and the persistence of policies, we include the level of the dependent variable in the previous period.

Finally, one may argue that from a conceptual point of view, the empirical model should also include the level of support (*PSE*). By including *PSE* as explanatory variable, one can analyze the relation between the policy level and instrument choice. However, there are two econometric reasons that render the inclusion of the level of support in our instrument choice equations problematic. First, *PSE* is endogenous, as the level of support is likely to depend itself on the policy instrument. Second, our explanatory variables, *exsh*, *gdppc*, and *d_gatt*, are also important determinants of the overall protection level. While the first problem could be solved potentially by using a simultaneous equation model, the second problem precludes finding good instruments for *PSE* in the

⁶In some cases, the PSE and the instruments share are negative: 13.3% of the observations have negative values for *mpsh* and 7.2% for *dpsh*. We dealt with this problem in three different ways. First, we recalculated *mpsh* and *dpsh* variables using absolute values of each instrument. Second, we ran the regressions excluding the negative values for *mpsh* or *dpsh*. Third, we ran the regressions with the subsample of the OECD member countries only, where there are no negative values. The model results are robust across these different samples.

⁷More specifically, $exsh = (\text{export value} - \text{import value})/\text{production value}$.

mpsh and *dpsh* equations. We therefore do not include *PSE* in the regressions.

Summarizing, in what follows we will run the following empirical specifications:

$$mpsh_{it} = \alpha_0 + \alpha_1 mpsh_{it-1} + \alpha_2 gdppc_{it-1} + \alpha_3 exsh_{it-1} + \alpha_4 d_gatt_{it-1} + v_{it}$$

$$dpsh_{it} = \beta_0 + \beta_1 mpsh_{it-1} + \beta_2 gdppc_{it-1} + \beta_3 exsh_{it-1} + \beta_4 d_gatt_{it-1} + \eta_{it}$$

where α_1 and β_1 are expected to be positive; α_1 , α_2 and α_3 are expected to be negative; and β_1 , β_2 and β_3 are expected to be positive. We first run the regressions using a simple OLS estimator, and later will do robustness tests with alternative estimation techniques.

OLS regressions are reported in **Table 2**, with Columns (1) and (5) for *mpsh* and *dpsh* regressions, respectively. All the relevant variables have their expected signs and are statistically significant at the 95% level or more. Moreover, the adjusted R^2 of the models, ranging from 0.58 to 0.80, indicates a good explanatory power of the selected variables.

gdppc has a significant negative effect on *mpsh*, the share of market price support, a relation that turn into positive when the share of direct income support, *dpsh*, is considered. These results are consistent with our theoretical argument that countries with lower administrative capacity and lower institutional development have a preference for price support. Also in line with our hypothesis, the net export share has a significant negative effect on *mpsh*, and a positive and significant effect on *dpsh*.

The 1994 GATT Agreement as captured by the dummy *d_gatt* is significantly negatively correlated with *mpsh*, and positively with *dpsh*. These results are consistent with the argument that the GATT constraints exert an effect on instrument choices: *mpsh* declined on average after the implementation of the 1994 URAA, and *dpsh* increased. In all regressions the coefficients of the lagged value of the dependent variable are positive and strongly significant. The magnitude of the lagged coefficients, ranging from 0.66 to 0.86, confirm a strong level of persistency in instrument choice.

Table 2. Regression results.

Dependent variable	<i>mpsh</i>				<i>dpsh</i>				
	Estimations method	OLS	LSDV	System GMM		OLS	LSDV	System GMM	
				One step	Two step			One step	Two step
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	
<i>gdppc</i>	-0.0018	-0.0107	-0.0035	-0.0036	0.0015	0.0085	0.0023	0.0027	
	2.98***	3.50***	3.21***	2.94***	4.14***	3.98***	2.70**	2.93***	
<i>exsh</i>	-0.0650	-0.0571	-0.1147	-0.1190	0.0132	0.0444	0.0195	0.0162	
	5.51***	1.28	5.06***	4.78***	1.45	1.77*	1.19	1.01	
<i>d_GATT</i>	-0.0511	-0.0472	-0.0882	-0.0871	0.0197	0.0273	0.0309	0.0274	
	4.03***	2.46**	4.15***	4.00***	3.16***	2.42**	4.14***	3.48***	
<i>Lagged_mps(dps)</i>	0.6608	0.4582	0.3950	0.4050	0.8038	0.5671	0.6930	0.6860	
	13.99***	8.29***	4.83***	4.72***	19.01***	8.82***	7.58***	7.62***	
Fixed effects	NO	YES	-	-	NO	YES	-	-	
Observations	517	517	517	517	517	517	517	517	
Countries	28	28	28	28	28	28	28	28	
Number of instruments	-	-	27	27	-	-	27	27	
F-statistics	82.4	78.7	-	-	152.6	197.8	-	-	
Adjusted R ²	0.58	0.64	-	-	0.75	0.80	-	-	
Test for AR(1): Pr > z	-	-	0.000	0.002	-	-	0.003	0.015	
Test for AR(2): Pr > z	-	-	0.245	0.334	-	-	0.246	0.241	
Hansen overid: Pr > chi ²	-	-	0.315	0.215	-	-	0.417	0.464	

Notes: *t*-statistics based on clustered standard errors under coefficients. All regressions include also a constant term. The System GMM estimator is implemented in STATA using the *xtabound2* routine, with the option *collapse* to limit the instruments proliferation. ***, ** and * p-value < 0.01, 0.05, 0.10 respectively.

Robustness Tests

We performed a series of additional robustness test. First, a potential problem in applying OLS to our specification is that the lagged dependent variable can be endogenous to the fixed effects in the error term, which gives the well know dynamic panel bias (see Roodman, 2009) [42]. A first step to deal with this is by removing the fixed effects from the error term, running the standard Least Square with Dummy Variables (LSDV) estimator. In doing so, we also control for any unobserved heterogeneity that are correlated with our explanatory variables.

The LSDV regression results are reported in Columns (2) and (6) for the *mpsh* and *dps* equation, respectively. As expected, the estimated coefficient of the lagged dependent variable is now lower in magnitude, but still strongly significant and has its expected sign in the two regressions. Generally speaking, all the OLS results are confirmed by the LSDV estimator, the only different being the effect of the trade status that now is estimated with less precision.

A potential problem with the (dynamic) LSDV estimator is that, when applied to a panel structure where $N > T$, namely the year dimension (T) is lower than the number of countries (N), it suffers of the so called Nickell bias, due to the endogeneity of the lagged dependent variable. To address this potential source of bias, the system GMM estimator proposed by Blundell and Bond (1998) [43] is used. This means estimating a system with the first-differences and the level equations, where the endogenous variables are instrumented by their level in the first-differenced equation and first-differenced instruments for the equation in level. Columns (3-4) and (7-8) report the results, considering both the one step and two step GMM option. As can be seen at the bottom of the columns, the Arellano-Bond tests for autocorrelation (AR1 and AR2), confirmed the presence of first order, but no second order, serial correlation, suggesting that the model dynamic is correctly specified. Moreover, the standard Hansen test confirms that in all cases our set of instruments is valid.⁸

The system GMM regressions strongly confirm our previous results, showing that the trade status and the level of development affected negatively the share of market price support and positively the share of direct payments, as well as the GATT dummy causes a shift from market price towards direct income support. Hence we can conclude that our results are very stable to different estimators.

5. Conclusions

In this paper we developed a theoretical political economy model to explain how various factors affect policy instruments choices. The theoretical model provides hypotheses on policy instrument use which are consistent with three stylized facts, *i.e.*, 1) their correlation with the level of a country's institutional development; 2) their correlation with a country's net trade position; and 3) the impact of GATT/WTO rules. Moreover, the model explains these key observations with a rational choice political economy model without having to rely on imperfect information of policy effects or on theories of bureaucratic inertia and obstruction.

In the last part of the paper we econometrically tested these theoretical predictions using OECD data on instrument choice in agricultural policy. Our empirical analysis confirms the hypotheses and provides strong evidence that the shift from distortionary to less distortionary instruments is positively influenced by institutional development, the net trade status, and the GATT/WTO framework. Moreover, we also find evidence of strong persistency of policy instruments. The main empirical limitation of the paper lies in the difficulty to directly isolate the effect of institutions from the one of economic development, due to the high correlation between the two dimensions. Hence, future development should pay particular attention to this issue, for example by establishing the extent to which different political institutions, such as forms of government or electoral rules, affect differently the patterns of policy instrument choice.

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⁸According to Roodman (2009), the instrument count does not exceed the number of groups and, to control for instrument proliferation that cause a weak Hansen test, we used the `xtabond2 collapse` option in STATA, instead of all available lags for instruments.

References

- [1] de Gorter, H. and Swinnen, J.F.M. (2002) Political Economy of Agricultural Policy. In: Gardner, B. and Rausser, G., Eds., *Handbook of Agricultural Economics*, **2B**, 1893-1943. [http://dx.doi.org/10.1016/S1574-0072\(02\)10023-5](http://dx.doi.org/10.1016/S1574-0072(02)10023-5)
- [2] Swinnen, J.F.M. (2010) The Political Economy of Agricultural and Food Policies: Recent Contributions, New Insights, and Areas for Further Research. *Applied Economic Perspectives and Policy*, **32**, 33-58. <http://dx.doi.org/10.1093/aep/p012>
- [3] Anderson, K. and Hayami, Y. (1986) The Political Economy of Agricultural Protection: East Asia in International Perspective. Allen and Unwin, Sydney.
- [4] Anderson, K., Rausser, G. and Swinnen, J.F.M. (2013) Political Economy of Public Policies: Insights from Distortions to Agricultural and Food Markets. *Journal of Economic Literature*, **51**, 423-477. <http://dx.doi.org/10.1257/jel.51.2.423>
- [5] Gardner, B.L. (1987) Causes of U.S. Farm Commodity Programs. *Journal of Political Economy*, **95**, 290-310. <http://dx.doi.org/10.1086/261456>
- [6] Olson, M. (1990) Agricultural Exploitation and Subsidization: There Is an Explanation. *Choices*, **4**, 8-11.
- [7] Anderson, K. and Tyers, R. (1988) The Pattern of Distortions to Agricultural Incentives. In: Maunder, A. and Valdes, A., Eds., *Agriculture and Governments in an Interdependent World*, Dartmouth, Aldershot, 175-186.
- [8] Honma, M. and Hayami, Y. (1986) Structure of Agricultural Protection in Industrial Countries. *Journal of International Economic*, **20**, 115-129. [http://dx.doi.org/10.1016/0022-1996\(86\)90064-4](http://dx.doi.org/10.1016/0022-1996(86)90064-4)
- [9] Swinnen, J.F.M. (1994) A Positive Theory of Agricultural Protection. *American Journal of Agricultural Economics*, **76**, 1-14. <http://dx.doi.org/10.2307/1243915>
- [10] Bullock, D.S. (1992) Redistributing Income Back to European Community Consumers and Taxpayers through the Common Agricultural Policy. *American Journal of Agricultural Economics*, **74**, 59-67. <http://dx.doi.org/10.2307/1242990>
- [11] de Gorter, H. and Tsur, Y. (1991) Explaining Price Bias in World Agriculture: The Calculus of Support-Maximizing Politicians. *American Journal of Agricultural Economics*, **73**, 1244-1254. <http://dx.doi.org/10.2307/1242452>
- [12] Swinnen, J.F.M. and de Gorter, H. (1993) Why Small Groups and Low Income Sectors Obtain Subsidies: The “Altruistic” Side of a “Self Interested” Government. *Economics and Politics*, **5**, 285-293. <http://dx.doi.org/10.1111/j.1468-0343.1993.tb00080.x>
- [13] Beghin, J.C. and Kherallah, M. (1994) Political Institutions and International Patterns of Agricultural Protection. *Review of Economics and Statistics*, **76**, 482-489. <http://dx.doi.org/10.2307/2109973>
- [14] Swinnen, J.F.M., Banerjee, A.N. and de Gorter, H. (2001) Economic Development, Institutional Economy of Agricultural Protection. An Econometric Study of Belgium since the 19th Century. *Agricultural Economics*, **26**, 25-43. <http://dx.doi.org/10.1111/j.1574-0862.2001.tb00052.x>
- [15] Olper, A. (2001) Determinants of Agricultural Protection: The Role of Democracy and Institutional Setting. *Journal of Agricultural Economics*, **52**, 75-92. <http://dx.doi.org/10.1111/j.1477-9552.2001.tb00926.x>
- [16] Olper, A. (2007) Land Inequality, Government Ideology and Agricultural Protection. *Food Policy*, **32**, 67-83. <http://dx.doi.org/10.1016/j.foodpol.2006.03.009>
- [17] Olper, A. and Raimondi, V. (2010) Constitutional Rules and Agricultural Policy Outcomes. In: Anderson, K., Ed., *The Political Economy of Agricultural Price Distortions*, Cambridge University Press, Cambridge and New York, 358-391. <http://dx.doi.org/10.1017/CBO9780511778964.016>
- [18] Gardner, B.L. (1983) Efficient Redistribution through Commodity Markets. *American Journal of Agricultural Economics*, **65**, 225-234. <http://dx.doi.org/10.2307/1240868>
- [19] Swinnen, J.F.M. (2008) The Perfect Storm: The Political Economy of the Fischler Reforms of the Common Agricultural Policy. CEPS Publications, Brussels.
- [20] Cassing James, H. and Hillman, A.L. (1985) Political Influence Motives and the Choice between Tariff and Quotas. *Journal of International Economics*, **19**, 279-290. [http://dx.doi.org/10.1016/0022-1996\(85\)90036-4](http://dx.doi.org/10.1016/0022-1996(85)90036-4)
- [21] Rodrik, D. (1986) Tariffs, Subsidies and Welfare with Endogenous Policy. *Journal of International Economics*, **21**, 285-299. [http://dx.doi.org/10.1016/0022-1996\(86\)90041-3](http://dx.doi.org/10.1016/0022-1996(86)90041-3)
- [22] Coate, S. and Morris, S. (1995) On the Form of Transfers to Special Interests. *Journal of Political Economy*, **103**, 1210-1235. <http://dx.doi.org/10.1086/601449>
- [23] Acemoglu, D. and Robinson, J.A. (2001) Inefficient Redistribution. *American Political Science Review*, **95**, 649-661. <http://dx.doi.org/10.1017/S0003055401003057>
- [24] Tracy, M. (1989) Government and Agriculture in Western Europe 1800-1988. Harvester Wheatsheaf, London.

- [25] Josling, T. (2007) Distortions to Agricultural Incentives in the EU. Paper Prepared for the World Bank Project “Distortions to Agricultural Incentives”.
- [26] Swinnen, J.F.M. (2009) The Growth of Agricultural Protection in Europe in the 19th and 20th Centuries. *The World Economy*, **32**, 1499-1537. <http://dx.doi.org/10.1111/j.1467-9701.2009.01247.x>
- [27] Irwin, D.A. (2008) Antebellum Tariff Politics: Coalition Formation and Shifting Regional Interests. *Journal of Law and Economics*, **51**, 715-741. <http://dx.doi.org/10.1086/590131>
- [28] Williamson, J.G. (2003) Was It Stolper-Samuelson, Infant Industry or Something Else? World Tariffs 1789-1938. NBER Working Paper 9656.
- [29] Josling, T. and Tangermann, S. (1999) Implementation of the WTO Agreement on Agriculture and Developments for the Next Round of Negotiations. *European Review of Agricultural Economics*, **26**, 371-388. <http://dx.doi.org/10.1093/erae/26.3.371>
- [30] Tangermann, S. (1999) Europe’s Agricultural Policies and the Millennium Round. *The World Economy*, **22**, 115-1178. <http://dx.doi.org/10.1111/1467-9701.00253>
- [31] Josling, T. (2000) New Farm Program in North America Their Treatment in the WTO: Discussion. *American Journal of Agricultural Economics*, **82**, 775-777. <http://dx.doi.org/10.1111/0002-9092.00078>
- [32] Hillman, A.L. and Ursprung, H.W. (1988) Domestic Politics, Foreign Interests and International Trade Policy. *American Economic Review*, **78**, 729-745.
- [33] Foster, W.E. and Raussler, G.C. (1993) Price-Distorting Compensation Serving the Consumer and Taxpayer Interest. *Public Choice*, **80**, 173-189. <http://dx.doi.org/10.1007/BF01047954>
- [34] Kono, D.Y. (2006) Optimal Obfuscation: Democracy and Trade Policy Transparency. *American Political Science Review*, **100**, 369-384. <http://dx.doi.org/10.1017/S0003055406062241>
- [35] Grossman, G.M. and Helpman, E. (1994) Protection for Sale. *American Economic Review*, **84**, 833-850.
- [36] Maggi, G. and Rodríguez-Clare, A. (2000) Import Penetration and the Politics of Trade Protection. *Journal of International Economics*, **51**, 287-304. [http://dx.doi.org/10.1016/S0022-1996\(99\)00029-X](http://dx.doi.org/10.1016/S0022-1996(99)00029-X)
- [37] Burgess, R. and Stern, N. (1993) Taxation and Development. *Journal of Economic Literature*, **31**, 762-830.
- [38] Rodrik, D. (1995) Political Economy of Trade Policy. In: Grossman, G. and Rogoff, K., Eds., *Handbook of International Economics*, Volume 3, North Holland, Amsterdam.
- [39] Bates, R.H. and Block, S. (2010) Agricultural Trade Interventions in Africa. In: Anderson, K., Ed., *The Political Economy of Agricultural Price Distortions*, Cambridge University Press, Cambridge and New York, 304-331. <http://dx.doi.org/10.1017/cbo9780511778964.014>
- [40] Brander, A.J. and Spencer, B.J. (1983) Strategic Commitment with R&D: The Symmetric Case. *The Bell Journal of Economics*, **14**, 225-235. <http://dx.doi.org/10.2307/3003549>
- [41] Dixit, A.K. (1984) International Trade Policy for Oligopolistic Industries. *Economic Journal*, **94**, 1-16. <http://dx.doi.org/10.2307/2232651>
- [42] Roodman, D. (2009) How to Do Xtabond2: An Introduction to Difference and System GMM in Stata. *Stata Journal*, **9**, 86-136.
- [43] Blundell, R.W. and Bond, S.R. (1998) Initial Conditions and Moment Restrictions in Dynamic Panel Data Model. *Journal of Econometrics*, **87**, 115-143. [http://dx.doi.org/10.1016/S0304-4076\(98\)00009-8](http://dx.doi.org/10.1016/S0304-4076(98)00009-8)

Appendix 1

Proof of Result 1:

Using Cramer's rule, it follows from (3) that

$$\frac{dt^*}{d\alpha} = \frac{1}{\det(H)} \left[\omega^V (b^t G_{ss} - b^s G_{ts}) - \frac{\partial \omega^R}{\partial \alpha} (R_t G_{ss} - R_s G_{ts}) \right], \quad (4)$$

$$\frac{ds^*}{d\alpha} = \frac{1}{\det(H)} \left[\omega^V (b^s G_{tt} - b^t G_{ts}) - \frac{\partial \omega^R}{\partial \alpha} (R_s G_{tt} - R_t G_{ts}) \right]. \quad (5)$$

Under the assumptions that $\frac{b^s}{b^t} > \frac{G_{ss}}{G_{ts}} > 1$ and $\frac{R_t}{R_s} > \frac{G_{ts}}{G_{ss}} < 1$, it follows that

$$\frac{dt^*}{d\alpha} > 0. \quad (6)$$

From the assumptions to secure uniqueness of the equilibrium and reaction function stability, $G_{tt} < G_{st}$, and $G_{ss} < G_{st}$ (see footnote 5), it follows that

$$\frac{ds^*}{d\alpha} > 0. \quad (7)$$

Performing comparative statics on h^* gives

$$\frac{dh^*}{d\alpha} = \frac{1-h^*}{(t^*+s^*)} \frac{dt^*}{d\alpha} \frac{h^*}{(t^*+s^*)} \frac{ds^*}{d\alpha}. \quad (8)$$

Using (6) and (7), it immediately follows that $dh^*/d\alpha > 0$.

The conditions imposed under Result 1 are stricter than necessary. Substituting Equations ((4) and (5)) into Equation (8) and rearranging gives the following condition for $dh^*/d\alpha > 0$:

$$\omega^V \left[s^* (b^t G_{ss} - b^s G_{ts}) - t^* (b^s G_{tt} - b^t G_{ts}) \right] > \frac{\partial \omega^R}{\partial \alpha} + \left[s^* (R_t G_{ss} - R_s G_{ts}) - t^* (R_s G_{tt} - R_t G_{ts}) \right]. \quad (9)$$

It is clear that condition (9) also holds, such that $dh^*/d\alpha > 0$, for several cases that violate the conditions imposed in Result 1. For example, condition (9) is satisfied when one of the two conditions is violated, but the other condition is sufficiently non-binding. Also, at $t^* = s^*$, condition (9) holds when $G_{tt} \cong G_{ss}$. Hence Result is more general than what the two conditions seem to suggest—the attractiveness of these two conditions is in their intuitive interpretation.

Proof of Result 2:

Using Cramer's rule, it follows that

$$\frac{dt^*}{d\beta} = \frac{-\omega^R}{\det(H)} [R_{t\beta} G_{ss} - R_{s\beta} G_{ts}], \quad (10)$$

$$\frac{ds^*}{d\beta} = \frac{-\omega^R}{\det(H)} [R_{s\beta} G_{tt} - R_{t\beta} G_{ts}]. \quad (11)$$

Since $R_{t\beta} < 0$ and $R_{s\beta} = 0$, it follows unambiguously that

$$\frac{dt^*}{d\beta} < 0 < \frac{ds^*}{d\beta}. \quad (12)$$

Performing comparative statics on h^* gives

$$\frac{dh^*}{d\beta} = \frac{1-h^*}{(t^*+s^*)} \frac{dt^*}{d\beta} \frac{h^*}{(t^*+s^*)} \frac{ds^*}{d\beta}. \quad (13)$$

Using (13), it follows that $dh^*/d\beta < 0$.

Appendix 2

Table A1. Support by policy instrument based on OECD PSE database (Million US \$).

	Period coverage	Initial year			Final year		
		<i>mpsh</i>	<i>ish</i>	<i>dpsh</i>	<i>mpsh</i>	<i>ish</i>	<i>dpsh</i>
<i>OECD countries</i>							
European Union	1986-2009	0.92	0.05	0.03	0.25	0.14	0.60
United States	1986-2009	0.44	0.18	0.37	0.18	0.30	0.51
Australia	1986-2009	0.75	0.17	0.08	0.00	0.51	0.49
Switzerland	1986-2009	0.83	0.10	0.07	0.54	0.06	0.41
Norway	1986-2009	0.72	0.10	0.18	0.55	0.05	0.40
Canada	1986-2009	0.56	0.16	0.28	0.58	0.07	0.35
Slovakia	1986-2003	0.88	0.08	0.04	0.51	0.19	0.31
Iceland	1986-2009	0.93	0.07	0.00	0.65	0.08	0.27
Czech Republic	1986-2003	0.82	0.05	0.14	0.64	0.11	0.25
Mexico	1986-2009	n.a.	n.a.	n.a.	0.35	0.43	0.22
Hungary	1986-2003	0.80	0.06	0.14	0.54	0.30	0.16
Japan	1986-2009	0.93	0.04	0.03	0.88	0.03	0.09
Turkey	1986-2009	0.77	0.23	0.00	0.92	0.03	0.05
Korea	1986-2009	0.99	0.00	0.00	0.92	0.03	0.05
New Zealand	1986-2009	0.19	0.48	0.32	0.42	0.55	0.03
Poland	1986-2003	0.75	0.17	0.08	0.60	0.38	0.01
<i>Non OECD countries</i>							
Latvia	1986-2003	0.99	0.01	0.00	0.26	0.36	0.38
China	1993-2007	n.a.	n.a.	n.a.	0.21	0.48	0.31
Ukraine	1986-2007	0.81	0.19	0.00	0.39	0.37	0.24
Slovenia	1986-2003	0.88	0.09	0.03	0.73	0.05	0.22
Estonia	1986-2003	0.88	0.10	0.01	0.51	0.32	0.17
Lithuania	1986-2003	1.00	0.00	0.00	0.63	0.23	0.14
Romania	1986-2005	0.96	0.03	0.00	0.89	0.07	0.04
Russia	1986-2007	0.65	0.24	0.11	0.54	0.44	0.03
Brazil	1995-2007	n.a.	n.a.	n.a.	0.53	0.46	0.02
Chile	1990-2007	0.96	0.04	0.00	0.13	0.86	0.01
South Africa	1994-2007	0.97	0.03	0.01	0.69	0.31	0.00
Bulgaria	1986-2005	0.99	0.01	0.00	0.31	0.69	0.00

Notes: The policy instruments considered are based on the following items of the PSE database: “market price support” refers to support based on commodity outputs (items A1 and A2, of the PSE database); “input subsidies” is the sum of payments based on input use and miscellaneous payments (items B and G); “direct payments” refer to different payments decoupled or partially decoupled from production (items from C to F). *mpsh*, *ish* and *dpsh* are the share of market price support, inputs subsidies and direct payments on total support, respectively. *Source:* own computation based on OECD PSE/CSE database (2010).