

# Relief for the environment? The importance of an increasingly unimportant industrial sector\*

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## Abstract

Deindustrialisation, stagnant real incomes of production workers and increasing inequality are latter-day features of many economies. It's common to assume that such developments pressure policy-makers to relax environmental standards. However, when heavily polluting industries become less important economically, their political importance also tends to diminish. Consequently, a regulator may increase the stringency of environmental policies. We find that declining industrial employment translates into stricter environmental standards. In contrast to previous studies, but consistent with our argument, we find that greater income inequality is associated with policies that promote a cleaner environment.

*Keywords:* Environmental regulations; deindustrialisation; income inequality; extreme bounds analysis

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# 1 Introduction

Among the more controversial views about economic growth and globalisation is that both will eventually benefit the environment (Arrow *et al.*, 1995). In part, this view is predicated on the nature of structural changes that are normally associated with trade liberalisation and economic development. More specifically, economic growth and the shift of production away from polluting sectors and “dirty” technologies help to arrest the deterioration in the environment. In addition, environmental quality is a normal good and wealthier economies will invest more heavily in environmental improvements and clean-up. According to this line of argument, another implication is that developing countries inevitably focus first on manufacturing production and basic forms of production, while tolerating some degradation in the quality of the environment. Compounding this feature is the fact that the political pressures associated with industrialisation are also likely to be influential. The factor owners employed in manufacturing industries lobby for less regulation of polluting activities. This accelerates the decay of the environment.

With the inevitable economic decline of basic manufacturing activities in more mature economies, the declining significance of basic manufacturing in industrialised countries may very well create social pressures that reduce the demand for pollution abatement. For instance, it has been argued that greater inequality of wealth and income could be bad news for the environment (see Boyce, 1994 and Torras and Boyce, 1998). Other studies show that the pattern of sectoral resource ownership matters and that greater income inequality can yield either stricter or weaker environmental policies. For example, McAusland (2003) shows that the owners of clean factors of production may be less green voters because they may bear the burden of pollution taxes through adverse terms of trade effects on the production of “clean goods”. However, in this paper we propose the argument that associated with falling industrial wages may be declining political influence exercised by the factor owners in the polluting manufacturing industries of the economy. These latter features are likely to be manifested in the political process, i.e., voting for change and a cleaner environment. In other words, structural change may not only involve less reliance being placed on the use of polluting inputs but also may have the signal virtue of altering the demand for environmental policies.

More liberalised trade and the rapid onset of skill-biased technological change have been linked with the declining real incomes received by production workers in manufacturing industries.<sup>1</sup> Free

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<sup>1</sup>There are many excellent surveys of the enormous literature on international trade and labour market outcomes, e.g., Johnson and Stafford (1999).

trade raises national income which, in aggregate terms, increases the value placed on the environment. Political economic considerations are therefore likely to lead to a cleaner environment. Trade liberalisation, which some authors continue to associate with increasing income inequality in OECD countries, may therefore be a “pro-environment” policy (see Grossman and Krueger, 1993 and Bommer and Schulze, 1999, for instance).

Associated with this relatively sanguine view has been an empirical relationship – in the form of an inverted U-shaped curve – between per capita income and various measures of environmental degradation. The relationship, or the *environmental Kuznets curve*, has been investigated for a wide variety of environmental indicators (e.g., Shafik, 1994; Selden and Song, 1994; Grossman and Krueger, 1995; Dinda, 2005). For any country, the implication is that economic growth will be associated with environmental degradation until a “critical” level of per capita income is attained; from that point, there will be an improvement in environmental conditions.

Of course, the turning points in the relationship between economic growth and environmental quality can be affected by the policies implemented by decision-makers (Shafik, 1994; Grossman and Krueger, 1995). Consequently, different political processes do not all imply that societies will grow their way out of environmental problems or that policies that promote economic growth can substitute for environmental policies. It is one objective of this paper to investigate the political pressures that may either reinforce or abrogate the environmental Kuznets curve.

This paper is also indirectly related to the political economy literature that deals with the effect of income inequality on redistributive policies and economic growth (e.g., Alesina and Rodrik, 1994 and Saint Paul and Verdier, 1996). A standard argument is that when income is more unequally distributed, the median voter is likely to be relatively less endowed with capital, either physical or human, and to thus favour a higher rate of capital taxation. A similar argument may well apply to pollution abatement policies. For instance, if the median voter is a low income worker who receives his or her livelihood from supplying labour to the basic manufacturing or pollution-intensive sectors, then greater income inequality may be associated with damage to the environment because it reduces the demand for pollution abatement. However, in this paper we will argue that precisely the opposite may occur. That is, increased income inequality may actually be associated with increases in the demand for environmental protection.

While environmental policies are shaped by the importance of potentially affected constituencies, the relative political importance of different constituencies is likely to change over time. The idea of an interaction between industry decline and endogenous policy formation is not a novel

one, of course (e.g., Cassing and Hillman, 1985). However, the perspective explored here is that the declining economic significance of polluting sectors in a developed economy is likely to be associated with greater income inequality. In turn, this is likely to reduce the “political clout” of the factor owners in the polluting sectors. In particular, as the workers in these sectors of the economy become less important economically, as reflected by their falling real incomes and falling employment levels, they also become less influential politically.<sup>2</sup> Consequently, a regulator, motivated by political considerations, will increase the stringency of environmental regulations. Of course, dynamic comparative advantages dictate that mature, developed economies shift resources away from basic manufacturing activities.

In the next section, we set out a simple model and derive some results that highlight the relationship between the sectoral decline of manufacturing and the stringency of environmental policies. In section 3, we present different types of empirical evidence to test the key findings of our model. First, we show that deindustrialisation may have a “silver lining” in terms of reducing emissions from basic manufacturing activities. Specifically, we show that organic water pollution and industrial employment levels are close complements. The latter finding is consistent with the environmental Kuznets curve hypothesis. Secondly, we investigate whether the erosion of labour market institutions that have traditionally supported blue-collar interests and lowered the inequality of earnings have led to tougher environmental regulations. We show that unionisation is strongly linked to the observed pattern of environmental taxation of industry relative to households. We conclude section 3 with a careful econometric study of panel data. In particular, we use extreme bounds analysis to examine whether countries with greater income inequality and declining manufacturing employment have more stringent environment policies. The last section concludes.

## 2 The model

Consider an economy with two types of jobs: “blue-collar” and “white-collar”, say. Further, assume that pollution creates blue-collar jobs (e.g., manufacturing) only.<sup>3</sup> All other jobs are white-collar (e.g., services, high-tech). Pollution afflicts all workers, however. A policy-maker must reconcile the conflict between blue-collar jobs and environmental quality, while also seeking the support of

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<sup>2</sup>This implies that “economic power” and political power are both unequally distributed (see Barro, 1999, p.4).

<sup>3</sup>In many developed countries, the combination of advanced pollution abatement technologies, as well as the toxic waste generated from office-situated photocopiers, suggests that it may not be entirely appropriate to classify blue-collar work as polluting and white-collar work as not.

both groups of workers.

To make matters transparent, assume that the economy has one blue-collar worker, indexed by  $b$ , and one white-collar worker, indexed by  $w$ . For expositional purposes, we assume that the white-collar worker is always employed. The blue-collar worker can be in one of two states at time  $t$ , employment ( $e$ ) or unemployment ( $u$ ). Each worker receives income  $y_t^{i,e}$  if they work,  $i = b, w$ . If unemployed, the blue-collar worker receives income  $y_t^{b,u}$ . We assume that the  $\{y_t^{i,j}\}$  are deterministic processes beyond the decision-maker's control. If worker  $i$  supplies one unit of labour inelastically each period,  $y_t^{i,e}$  can be interpreted as the wage rate in period  $t$  for worker  $i$ .

At time  $t$ , manufacturing generates a residual called pollution,  $s_t$ . Pollution and blue-collar labour are complementary inputs. If the policy-maker wants industry to create more blue-collar jobs, he must allow greater production - and pollution. The number of blue-collar jobs is given by the demand curve,  $l_t^b = f(y_t^{b,e}, s_t)$ , with

$$\frac{\partial f(y_t^{b,e}, s_t)}{\partial y_t^{b,e}} = f_{y_t} < 0, \quad \frac{\partial f(y_t^{b,e}, s_t)}{\partial s_t} = f_{s_t} > 0 \quad \text{and} \quad \frac{\partial^2 f(y_t^{b,e}, s_t)}{\partial s_t^2} = f_{s_t s_t} < 0.$$

The pollution stock,  $p_t$ , decays at rate  $\delta \in [0, 1]$ . The transition equation is

$$p_{t+1} = (1 - \delta)p_t + s_t.$$

Equating  $\delta$  to 1 gives the classic case of a pollutant that dissipates immediately;  $\delta = 0$  is the case for a pollutant that never dissipates.<sup>4</sup> All residents suffer from  $p_t$ .

Workers are assumed to have von Neumann-Morgenstern utility functions. A worker's expected utility, not yet knowing whether he will be employed or unemployed, is

$$E_t U_t^i = l_t^i U(y_t^{i,e}, p_t) + (1 - l_t^i) U(y_t^{i,u}, p_t), \text{ for } i = b, w.$$

Since the number of blue-collar workers has been normalised to one,  $(1 - l_t^b)$  can be interpreted as the unemployment rate for blue-collar workers. (Recall that white-collar workers are always employed, i.e.,  $l_t^w = 1$ .)

In traditional political economy models it is assumed that the policy-maker maximises a weighted average of the welfare of constituents over his career. The policy-maker might be a politician who considers voter welfare to win elections, or he might be a regulator, who considers

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<sup>4</sup>Iodine-129, a fission product of uranium 235, is an example of the latter, its half-life stretches into millions of years (Carter, 1987).

constituent welfare to win promotions. It is often assumed that a committee, containing conflicting interests, would make the same decision as a single leader who reconciles these interests. In a democracy with complete information, the number of decision-makers affects the time and resource costs of the decision, but does not affect the decision itself. This has been termed a “dialectic” model of interest groups (Becker, 1983).

In the current context, the political weights that a policy-maker assigns to the welfare of blue- and white-collar workers may reflect the relative political influence of the two types of workers. Different weights may be attributed to interest groups according to the degree of organisation or unionisation or may simply vary with size of membership, for instance. In the latter case, however, note that the weights may vary inversely or directly with group size depending on such issues as the nature of the political process and whether simple majority voting rules or free-riding are more important for determining the political future of decision-makers. These issues are discussed further below.

The common agency model developed by Bernheim and Whinston (1986), and applied by Grossman and Helpman (1994), provides microeconomic foundations for the political weights that are assigned to each interest group in a society. Grossman and Helpman show that if policy-makers, when choosing a policy ( $s$ , say), care about interest groups’ welfare ( $V^j(s)$ ) on one hand and about campaign contributions on the other hand, then they actually end up maximising a weighted sum of the interest groups’ objective functions. That is, the policy-maker will choose a policy  $s$  to maximise

$$V^g(s) = \sum_{j \in I} (I^j + \alpha^j) V^j(s), \quad (1)$$

where  $V^g(s)$  denotes the policy-maker’s welfare function,  $I$  is the set of all interest groups, the indicator function,  $I^j$ , equals 1 if the interest group is engaged in lobbying activities;  $I^j = 0$  otherwise.<sup>5</sup>

Grossman and Helpman concentrate on studying the distortionary effects of lobbying and assume that each group is originally given the same weight  $\alpha^j = \alpha$ ,  $j \in I$ . From equation (1), it is clear that, “despite” the presence of lobbying, the outcome will be equal to the efficient solution

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<sup>5</sup>The application of the common agency framework by Grossman and Helpman to model the political decision-making process does have its limitations. As a practical matter, political contributions by organised lobby groups are illegal in some countries. From a theoretical viewpoint there is a two-sided moral hazard problem associated with either politicians reneging on their policy promises after contributions are received or lobbies reneging on promised contributions once preferred policies are locked in place. However, note that the policy-maker’s objective described by equation (1) is quite general and could be alternatively motivated by a linear, additive version of a political-support function (Hillman, 1982).

selected by the utilitarian social planner that would assign equal weights to everybody. The political system creates inefficiencies when some groups in the economy do *not* lobby. Naturally, the policy-maker more heavily weights the policy preferences of the interest groups that do lobby (see Potters and Sloof, 1996 and Aidt, 1998). In the spirit of Grossman and Helpman, we analyse cases in which both or either of the blue- and white-collar workers may organise lobbies to help attain their preferred environmental policy settings, perhaps, via a trade union and an environmental lobby, respectively.

By construction, in the present model the welfare of the two interest groups is transparent. This contrasts to the set-up of political economy models that deal with the more complicated issue of how policy-makers coordinate national environmental policy and trade policy. In Hillman and Ursprung (1992), for instance, whether environmentalists support higher trade barriers depends on whether it is consumption or production that pollutes. In this paper, the welfare of *all* workers is adversely affected by greater pollution-intensive production. However, blue-collar workers also benefit from higher pollution. While white-collar workers would always prefer smaller production and pollution, the interests and the lobbying stance of blue-collar workers therefore depend on the elasticity of production and employment with respect to greater pollution emissions as well as the reservation utility if they were to be unemployed. The relative positions taken by the two interest groups are therefore always likely to be opposed to one another.

Returning to the problem at hand, the decision-maker's problem is

$$\max \text{ over } s_t E_0 \sum_{t=0}^{\infty} (1 + \rho)^{-t} ((I^b + \alpha)U_t^b + (I^w + \alpha)U_t^w) \quad (2)$$

subject to  $p_{t+1} = (1 - \delta)p_t + s_t$ ,  $p_0$  given,

where  $\rho > 0$  is the rate of time preference. Letting  $p_t$  be the state and  $s_t$  be the control, Bellman's equation is

$$V_t(p_t) = \max E_t \{ \theta^b U_t^b + \theta^w U_t^w + \beta V_{t+1}(p_{t+1}) \}, \quad (3)$$

where  $\beta = (1 + \rho)^{-1}$  and  $\theta^i = \{\alpha, (1 + \alpha)\}$ ,  $i = b, w$ .

Standard solution techniques yield the Euler equation (see Appendix for details),

$$\begin{aligned} \theta^b f_{s_t} \Delta_t &+ \beta \left( \theta^b l_{t+1}^b U_{p_{t+1}}^{b,e} + \theta^b (1 - l_{t+1}^b) U_{p_{t+1}}^{b,u} + \theta^w U_{p_{t+1}}^w \right) \\ &- (1 - \delta) \beta \theta^b f_{s_{t+1}} \Delta_{t+1} + \beta \varepsilon_{t+1} = 0, \end{aligned} \quad (4)$$

where  $\Delta_t = U(y_t^{b,e}, p_t) - U(y_t^{b,u}, p_t)$  and  $\varepsilon_{t+1} = E_t V'_{t+1} - V'_{t+1}$  is the one-period ahead forecast error ( $\varepsilon_{t+1}$  is serially independent with mean zero).

In general, it is difficult to find closed-form solutions for the optimal  $s_t$  (or  $p_{t+1}$ ) sequence. However, a simple perfect foresight example does illustrate some of the fundamental driving forces. For example, consider  $U(y_t^i, p_t) = y_t^i - \gamma p_t$ ,  $\gamma > 0$ ,  $i = b, w$  and  $f(y_t^{b,e}, s_t) = \ln s_t - \kappa \ln y_t^{b,e}$ ,  $\kappa > 0$ . By appropriate substitutions into equation 4 and solving the difference equation, we obtain

$$s_t = \frac{(1 - (1 - \delta)\beta)x_t}{\beta\gamma(1 + \psi)}, \quad (5)$$

where  $x_t = y_t^{b,e} - y_t^{b,u}$  and  $\psi = \frac{\theta^w}{\theta^b}$ .

To highlight the effects of changing real wages for blue-collar employees we suppose that  $x_t = \omega g^{-t}$ , where  $g > 1$  if wages are shrinking (growing) exponentially.<sup>6</sup> From equation (4), it follows that

$$s_t = \frac{(\rho + \delta)\omega}{\gamma(1 + \psi)g^t}. \quad (6)$$

The comparative dynamic effects are summarised in the following Proposition.

**Proposition 1:** Suppose that blue-collar income is given by  $x_t = \omega g^{-t}$ , then the stringency of environmental regulations falls in

- (a) blue-collar income,  $\omega$ ;
- (b) the policy-maker's discount rate,  $\rho$ ;
- (c) the pollution decay rate,  $\delta$ .

Environmental regulations are stricter

- (d) the higher is the rate of diminution of blue-collar income,  $g$ ;
- (e) the higher is the marginal disutility of pollution,  $\gamma$ ;

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<sup>6</sup>Alternatively,  $g$  may represent a direct measure of wage inequality between white- and blue-collar workers. For example, define  $y_t^{w,e}/y_t^{b,e} = g_t$  and  $y_t^{b,u} = 0$ .



(f) when blue-collar workers do *not* lobby the policy-maker.

**Proof:** Differentiating equation (6) yields parts (a) through (e). As for part (f), note that if only blue-collar workers lobby, then  $\psi^b = \frac{\alpha}{1+\alpha}$ , if only white-collar workers lobby, then  $\psi^w = \frac{1+\alpha}{\alpha}$ , and if both groups lobby, then  $\psi^{b,w} = 1$ . Clearly,  $\psi^b < \psi^{b,w} < \psi^w$ . Finally, note that environmental quality improves in  $\psi$ .

There are some transparent implications. For example, if the policy-maker discounts the future more heavily, then this is associated with deteriorating environmental quality. Congleton (1992) shows that autocratic countries are inclined to select less stringent environmental regulations. He argues that dictators tend to have shorter time horizons (i.e., higher  $\rho$ ) and are less likely to adopt pro-environment policies, since the benefits of doing so are likely to accrue only after they have left office, whereas the costs are incurred earlier.<sup>7</sup> A higher pollution decay rate,  $\delta$ , also increases pollution. This somewhat counter-intuitive result occurs because worker utility depends on  $p_t$ , and not on  $p_{t+1}$ , so that the policy-maker is likely to take a less conservative attitude with a pollutant that dissipates immediately, as opposed to the case for a pollutant that never dissipates. Consistent with this finding is that policy-makers are more likely to be “policy-active” for the types of pollutants with short-term and local impacts (see Barbier, 1997 and Emerson and Pendleton, 2004).

Higher income for workers in pollution-producing industries ( $\omega$ ) is associated with an increasing amount of economic importance attached to the polluting sector of the economy. However, if this economic “weight” falls, then because environmental quality is a normal good, the stringency of environmental regulations rises over time and consequently, so too does the quality of the environment. Strictly speaking, it is not just the continued erosion of the earnings of blue-collar workers that beneficially impacts pollution emissions. More generally, it is the falling *relative* earnings of working in the polluting sector. For example, if the income while unemployed increases more rapidly than the income while employed in the polluting sector, then the same benefit to the environment results. Thus, some authors have argued that more generous unemployment benefits and changes to cash transfer and income tax systems have arisen to ensure worker acquiescence to potentially disruptive microeconomic reforms, such as trade liberalisation (e.g., Rodrik, 1998). Hence, while the earned income distribution may have widened in many OECD countries, the same is not true for the post-tax and post-government transfer distribution of income (see, e.g.,

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<sup>7</sup>For reasons other than the expected shorter duration of dictatorships, Olson (1993) argues that dictators wish to maximise tax revenues and thus oppose any policies that would reduce revenue, e.g., those that result from increased pollution abatement expenditures.

Smeeding and Gottschalk, 1995).

Recall that when each interest group receives an equal political weight that this is equivalent to the utilitarian social planner's problem. However, if a lower political weight is attached to blue-collar worker interests ( $\psi$ ), then less importance is attached to the polluting sectors of the economy. As noted above, the increased likelihood of free-riding in larger political constituencies poses problems for a straightforward interpretation of the political weights attached to interest groups by policy-makers. In democratic countries, government officials may favour groups with more members. Larger groups are also likely to have greater electoral resources. However, groups with more members tend to be prone to free-riding problems. In addition, larger groups are likely to be costlier to organise, more difficult to develop a coherent and consistent platform for, and to involve greater difficulties in ensuring the political participation of all members. Potters and Sloof (1996) provide a fairly comprehensive survey of the empirical effects of group size on political outcomes. Overall, they conclude that free-riding is, in fact, a serious problem for larger, *unorganised* groups. On the other hand, larger, *organised* groups, such as trade unions, for example, wield greater influence. In deindustrialising economies, the reduction in blue-collar power has in part been manifested by the declining influence of trade unionism (see Freeman, 1993). Clearly, deunionisation is likely to reinforce the declining political significance attached to the blue-collar interests in relaxing environmental standards.<sup>8</sup> We discuss this further in the next section.

**Corollary 1** (*The environmental Kuznets curve*): Environmental quality has an inverted U-shape in economies in which blue-collar incomes initially rise and then fall. Hence, deindustrialisation is associated with a cleaner environment.

**Proof:** Differentiating equation (6) with respect to  $t$  yields

$$-k(g^{-t}) \ln g \begin{cases} > 0, & \text{if } g < 1 \\ = 0, & \text{if } g = 1 \\ < 0, & \text{if } g > 1 \end{cases} \quad (7)$$

where  $k = \frac{(\rho+\delta)\omega}{\gamma(1+\psi)}$ .

The environmental effects of an eventual decline of blue-collar income are evident from equa-

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<sup>8</sup>Fredriksson and Gaston (1999) show that the stance taken by the trade union movement on the environmental policies is far from unambiguous. Among other things, union "environmentalism" may depend on the risk of unemployment for their members as well as the presence of unemployed, non-unionised "outsiders".

tion (7). In deindustrialising economies, the relative income from basic manufacturing and heavily polluting sectors eventually declines (i.e.,  $g$  increases over time) and an environmental Kuznets curve is thereby generated. That is, pollution initially rises, but then eventually falls. It can be argued that the growth rate of blue-collar income reflects known patterns of structural change in industrialised economies. Forward-looking anticipation of a declining polluting industrial sector is also likely to lead to “making hay while the sun shines” behaviour. Anticipating rapid future declines in workers’ returns to polluting activities lead to opportunistic, but short-term increases in polluting activity.

As  $g$  rises, the improvement in environmental quality becomes more marked. The forces that reflect the growing economic unimportance of the polluting sector of the economy generate the shape of an environmental Kuznets curve.<sup>9</sup> The political importance of the polluting sector affects the state of environmental quality. As the political weight attached to blue-collar workers declines, then so too does the stock of pollution. Hence, falling  $g$  and  $\psi$  work together to accelerate the improvement in environment quality.

### 3 Empirical implications and evidence

In this section we present three types of evidence to examine the predictions of the model. More informally, in the next two sub-sections we present some simple tabulations and correlations. First, we show that industrial employment and polluting activity go hand-in-hand. Secondly, we show that more unionised economies, which also tend to have more equitable earnings distributions, favour the imposition of eco-taxes on consumers rather than on industry. Finally, and more importantly, we present a formal econometric analysis using panel data to investigate the determinants of the stringency of environmental policy.

#### 3.1 Deindustrialisation and the environment

A key feature of the model is that deindustrialisation is associated with a cleaner environment. This occurs for two reasons – an economic one and a political one. The economic reason involves the trade-off between a cleaner environment and the production of basic manufacturing goods – the latter requires the employment of blue-collar workers. In turn, higher employment in this sector is associated with greater pollution emissions.

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<sup>9</sup>In the context of the present model, it is important to note that “balanced growth”, which increased both blue-collar and white-collar wages equally, would not generate the same hump-shaped environmental Kuznets curve. It is the declining economic and political importance of the polluting sector that is crucial.

There are also political pressures that imply a positive correlation between manufacturing employment levels and pollution emissions. For example, if underlying economic growth and dynamic comparative advantages reduce production and lower employment in basic manufacturing activities, then the remaining workers in these sectors are likely to receive smaller consideration in the political process. Consequently, policy-makers weight more heavily the preferences of workers (and voters) involved in the production of “cleaner” goods. Doing so, of course, simply reinforces the decline in basic manufacturing industry.

It follows that industrial employment and pollution emissions are highly and positively correlated with one another. To provide a visual perspective of this feature, we present plots for organic water pollutants and industrial employment for 18 countries in figure 1. The data are from a study of industrial emissions for a limited number of countries by Hettige *et al.* (1998). Data on water pollution are more readily available than other emissions data because most industrial pollution control programmes start by regulating organic water emissions. Such data are also fairly reliable because sampling techniques for measuring water pollution are more widely understood and much less expensive than those for air pollution. The emissions estimates represent biochemical oxygen demand (BOD) in kilograms per day for each country and year.<sup>10</sup> The employment data are from the United Nations Industrial Development Organisation (UNIDO). The data series for each country are for 1975 to 1992 (or 1993). For ease of comparison, the emissions and employment data were converted to indices (with 1975 as the base year).

– Figure 1 here –

A couple of observations are immediate. First, as anticipated, changes in industrial employment and water pollution are strongly complementary. For instance, Canada suffered very steep reductions in its manufacturing employment in the late 1980s and early 1990s due to an unexpectedly severe recession and to the, somewhat more debatable, effects of the passage of the Canada-U.S. Free Trade Agreement (see Gaston and Treffer, 1997). In non-OECD countries in which manufacturing employment has grown rapidly (e.g., Singapore) there has been a corresponding increase in pollution emissions.

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<sup>10</sup>Emissions of organic pollutants from industrial activities are a major source of water quality degradation. The Hettige *et al.* (1998) data are based on measurements of plant-level water pollution in a number of countries. The focus is on organic water pollution as indicated by the presence of organic matter, metals, minerals, sediment, bacteria and toxic chemicals. The pollution is measured by biochemical oxygen demand, BOD, because it provides the most plentiful and reliable source of comparable cross-country emissions data. BOD measures the strength of an organic waste in terms of the amount of oxygen consumed in breaking it down. A sewage overload in natural waters exhausts the water’s dissolved oxygen content. Wastewater treatment, by contrast, reduces BOD. (The previous discussion is drawn from *World Development Indicators* (1999, p.143).)

One of the key findings of the Hettige *et al.* (1998) study was the remarkable stability of emissions per worker across a large range of countries. It follows that government attempts to regulate or limit water pollution is tantamount to regulating employment in the industries that generate heavy water pollution (in particular, the primary metals, paper and pulp, chemicals, food and beverages, and textiles industries). Marxsen (1997) argues that the increasing prevalence of pollution control laws has been partly responsible for the productivity growth slowdown, and *inter alia* employment, in the U.S. manufacturing sector.

In those countries that have experienced the most marked deindustrialisation (see Baldwin and Martin, 1999), e.g., the United Kingdom, pollution emissions have steadily fallen as employment in manufacturing industry has declined over the entire time period. In the relatively few developed countries that have not deindustrialised as rapidly, e.g., Japan, emissions have remained relatively unchanged. To the extent that most developed countries have been deindustrialising, it is not surprising that the balance of evidence seems to favour the environmental Kuznets curve hypothesis for most measures of water pollution (see Grossman and Krueger, 1995). Baldwin and Martin (1999) note that the “first wave” of globalisation (pre-WW1) which generated rapid economic development for many countries was characterised by rapid industrialisation. In contrast, the “second wave” of globalisation (since 1960), which generated rapid income growth for many developed countries, has been characterised by a process of deindustrialisation and the associated steady decline in industrial employment.

### 3.2 Unions and the environment

Next, we examine a subsidiary implication of the model developed in section 2. Specifically, as the political institutions that have traditionally supported blue-collar interests have declined in importance, associated environmental regulations have toughened.

Portney (1982) argues that increasing unemployment pressures policy-makers to ease environmental standards.<sup>11</sup> By implication, political pressure brought to bear on environmental policy-makers is greater when industrial employment levels fall. Fredriksson and Gaston (1999) have noted that unions lobbying on behalf of unemployed members may encourage policy-makers to respond favourably to calls for easing environmental restrictions. Yandle (1983) found that state expenditures on environmental regulation in the United States were negatively related to the num-

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<sup>11</sup>In reference to the stance taken by European labour unions on environmental regulation, Klepper (1992, p.253) notes that the primary objective of securing or increasing employment was thought to be threatened by environmental policies.

ber of workers in polluting industries and positively related to the percentage of the manufacturing industry workforce that was unionised. He interpreted the former relationship as evidence that policy-makers operate according to an environmental quality versus jobs trade-off and the latter relationship as evidence of union rent-seeking.<sup>12</sup>

Overall, one expects that the decline of unionisation in many countries has helped the passage of more stringent environmental regulations that affect industry. On the surface, the evidence on this point is somewhat mixed. According to Tobey's (1990) indices of environmental stringency, two of the three countries with the strictest environmental standards (the United States and Japan) have among the lowest rates of union membership in the world (as well as the lowest percentage of workers covered by collective bargaining agreements). However, the third, Sweden, has among the highest rates of unionisation in the world. Fredriksson and Gaston (1999) explain this phenomenon by noting that the ambiguous stance of the trade union movement on environmental policies depends on the exposure to unemployment of their *own* members. It needs to be emphasised that it is the actual level of industrial employment, rather than the rates of unionisation of a presumably smaller pool of manufacturing workers in deindustrialising economies, that may be of greater significance for policy-makers.

In this paper, special interest groups, representing blue-collar and white-collar interests, help to determine the stringency of environmental policy. In most countries, trade unions are the most visible advocates of blue-collar interests.<sup>13</sup> If blue-collar workers perceive a trade-off between environmental regulations and jobs, unions are likely to oppose policies that threaten manufacturing employment.

Many OECD countries have recently introduced, or are considering implementing, fiscal instruments or "eco-taxes" for environmental management. Consider table 1, which illustrates a specific example of an "eco-tax". The data in columns (1) and (2) contain data on tax rates for household-use and industrial-use fuel for a number of OECD countries. A number of features are apparent. For example, the household use tax rate is highest in France and the industrial use tax rate is highest in Switzerland and both tax rates are lowest for the United States. The differences in tax rates across countries reflect a number of influences, including such disparate factors as the political importance of community environmental concerns as well as fiscal considerations.

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<sup>12</sup>Endersby and Munger (1992) found that union contributions were given disproportionately to members of Congress who were members of committees with legislative and regulatory jurisdiction over activities that would affect labour.

<sup>13</sup>Of significance, for the purpose of this paper at least, is that countries with encompassing labour market institutions (i.e., large unionised sectors with centralised bargaining) are characterised by lower wage inequality (see e.g., Rowthorn, 1992, Freeman, 1993 and OECD, 1997).

– Table 1 here –

Of more interest is the difference between the rates of taxation for industrial and household use. Column (3) indicates that the tax rate on industrial use fuel is 61 percent of the tax rate on household use fuel in Denmark; in the United States, the industrial use tax rate is 15 percent higher than the household use tax rate. Once again, there are likely to be a number of determinants of these cross-national differences. However, these differences are also likely to reflect the importance of industry concerns (i.e., shareholders and workers). Environmental and community concerns are likely to be reflected in the tax rate levels. On the other hand, national differences in the *relative* tax rates for industrial and household fuels are likely to reflect the *relative* influence of industry vis-a-vis households in the political process in which tax rates are determined. Moreover, institutional features of the labour market are important determinants of industry and union lobbying incentives, and consequently, the observed pattern of environmental policy.

Consider the last column of table 1 - “Union coordination index”. Layard *et al.* (1994, pp.80-81) argue that when unions have a national focus (designated by an index of ‘3’), they take into account the common interests of the workforce in full employment “rather than bargaining as atomistic groups of insiders” (designated by an index of ‘1’). The data reveal that bargaining at the national level is negatively related to the tax rate disparity (the ‘Ratio’ column). Of course, this correlation may be purely coincidental. On the other hand, it appears that a strong coordinated union movement is associated with relatively higher tax burdens on households (i.e., which comprise blue-collar and white-collar workers) as opposed to industry (which primarily employ the blue-collar workers). Overall, unionisation does appear to be strongly linked to the observed pattern of environmental taxation of industry relative to households.

### **3.3 Inequality, industrial employment and environmental regulation**

To conclude the empirical section, we present a formal econometric analysis of the determinants of environmental policy. As much as possible, we follow the empirical specifications previously suggested in the literature. The major innovation, of course, is the introduction of variables suggested by our own model. Another major step forward is our use of extreme bounds analysis to isolate the most important determinants of environmental regulation.

### 3.3.1 Data and variables

To proxy environmental stringency we use the lead content of gasoline. This measure has been used in previous research (e.g., Damania *et al.*, 2003); its major advantage is that the data are available as a panel for the period from 1982 to 1992 and for up to 48 countries. We transform the series by taking the logarithm of it and multiplying it by -1. It is denoted by *LREGS*.<sup>14</sup>

Our primary focus is examining the importance of blue-collar workers for the political process that shapes the environmental policy. Damania *et al.* (2003) use the percentage of the labour force employed in industry (*INDSHEMP*) to proxy political pressure by industrial workers.<sup>15</sup> This pressure is also central to our model's predictions. Since environmental regulations may increase employment uncertainty, industrial workers use their political power to prevent stricter regulations. The other variable important in our model is wage inequality. The stringency of environmental regulations is predicted to increase as blue-collar income declines. If wages fall exogenously (e.g., due to skilled-labour biased or sector-biased technological change that favours white-collar workers), then we predict a more stringent environmental policy (i.e., a lower lead content of gasoline).

In a recent paper, McAusland (2003) argues that greener pollution policies could be associated with either greater or smaller income inequality. In earlier research, inequality has often been associated with an intensification of polluting activities (e.g., Boyce, 1994; Torras and Boyce, 1998). Our model's predictions point in precisely the *opposite* direction. That is, as an economy deindustrialises income inequality may increase as the wages paid to manufacturing workers in low-tech, pollution-intensive industries fall. As it does so, the influence of blue-collar workers in the policy-making process declines.<sup>16</sup> To measure income inequality we use the Gini coefficient data recently updated and re-calculated by Francois and Rojas-Romagosa (2005).<sup>17</sup>

Needless to say, a large number of other variables have been proposed as determinants of the level of environmental stringency. Cole *et al.* (2006) use the urban population share (*URBAN*) to test whether a greater exposure to industrial pollution by a larger number of citizens increases environmental stringency. Cole *et al.* (2006) also argue that the demand for environmental quality increases with per capita income (*LGDPPC*). On the other hand, Congleton (1992) emphasises

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<sup>14</sup>A rise in the transformed *LREGS* therefore represents a higher level of environmental stringency. Hilton and Levinson (1998) and Octel (1982-92) provide a more detailed description of these data.

<sup>15</sup>Unless stated otherwise, all data is taken from the World Bank's World Development Indicators (WDI 2003).

<sup>16</sup>Consistent evidence is provided by Taylor (1998). He uses data for State expenditures per capita in the United States for hazardous waste in the 1980's and for air pollution in the 1960's and rejects the hypothesis that there is a trade-off between future environmental quality and current manufacturing jobs.

<sup>17</sup>These authors address measurement error problems in the well-known World Bank inequality dataset of Deininger and Squire (1996) and produce a new dataset of consistent inequality series.



that the effect of per capita income is theoretically indeterminate (even though he estimates a positive relationship in his study). He argues that despite the fact that the demand for environmental quality is likely to be increasing with personal wealth, voters and tax-payers also have to bear a higher share of the costs associated with environmental regulations. These costs reduce national income. A similar ambiguity is predicted for the effect of population density (*LPOPDENS*). Congleton argues that population also serves as a proxy for a country's human capital resources.

Damania *et al.* (2003) contend that more open economies will have higher environmental standards. McAusland (2003) shows that trade openness and the pattern of factor ownership are important determinants of the preference for pollution standards. If an economy is small and open then environmental policies have no effects on the terms of trade. Hence, if the poor have a larger relative stake in the production of dirty goods, then they may vote for weaker policies when the economy is open because there are no beneficial terms of trade effects associated with environmental regulations. We therefore use trade intensity (*TRADE*), measured by the ratio of trade flows to GDP. Another commonly used openness measure is foreign direct investment, which we measure as the ratio of the net inflows of FDI to GDP (*FDIGDP*).<sup>18</sup> As a final proxy for openness we employ the KOF-globalisation index (*GLOBAL*) (see Dreher, 2006).<sup>19</sup>

Congleton (1992) argues that autocratic countries have lower environmental standards because their rulers have shorter time horizons. Consequently, the incentives to invest in environmental protection are lower. Following Congleton, we also include a dictatorship dummy (*DICT*) which takes the value one if the executive index of electoral competitiveness is smaller than three (see Beck *et al.*, 1999). In addition, we employ *POLFREE* which we measure as the average of the Freedom House (2005) indices for civil liberties and political rights. Another included variable is *LEFT*, which measures whether the chief executive has a left-wing orientation or not.<sup>20</sup> Neumayer (2003) argues that a left-wing executive is traditionally more likely to care for the interests of blue-collar workers. As they work mostly in dirty sectors this may reduce environmental stringency (see also Fredriksson and Gaston, 1999). However, Neumayer notes that left-wing governments might also be more amenable to policies that protect the environment.<sup>21</sup>

Damania *et al.* (2003), Fredriksson *et al.* (2003) and Fredriksson and Svensson. (2003) emphasise the role that corruption might play in affecting the political agenda. Accordingly, we include

<sup>18</sup>Fredriksson and Gaston (2000) argue that greater capital mobility weakens the incentive for capital owners to lobby, and *ceteris paribus*, may lead to stricter environmental policies.

<sup>19</sup>This index incorporates economic as well as the political and social dimensions of globalisation.

<sup>20</sup>This variable is taken from Beck *et al.* (1999).

<sup>21</sup>Neumayer argues that blue-collar workers are likely to be among the first exposed to the effects of environmental degradation.

*CORRUPT* to measure the level of government corruption. This variable is the “Government Honesty” variable reported by the International Country Risk Guide (ICRG).<sup>22</sup> For a summary of all variables, their sources, their descriptions as well as the study that originally proposed them see table 2; table 3 gives the descriptive statistics and correlations of the variables.

– Tables 2 and 3 here –

### 3.3.2 Extreme Bounds Analysis

Since there are several studies that investigate the effects on environmental stringency, there is a long list of potential explanatory variables. Studies often restrict their analysis to certain subsets of these variables and often ignore the effects of any omitted variable bias when other variables are not included. In addition to any model uncertainty, the limited number of observations often restricts the power of statistical tests that rule out irrelevant explanatory variables.

In order to address these issues we use extreme bounds analysis (EBA), as proposed by Leamer (1983) and Levine and Renelt (1992). EBA enables us to examine which explanatory variables are robustly related to our stringency measure and is a relatively neutral way of coping with the problem of selecting variables for an empirical model in situations where there are conflicting or inconclusive suggestions in the literature.

To conduct an EBA, equations of the following general form are estimated

$$Y = \beta_M M + \beta_F F + \beta_Z Z + v \quad (8)$$

where  $Y$  is the dependent variable,  $M$  is a vector of “commonly accepted” explanatory variables and  $F$  is a vector containing the variables of interest. The vector  $Z$  contains up to three possible additional explanatory variables (as in Levine and Renelt, 1992) which, according to the broader literature, are related to the dependent variable. The error term is  $v$ . The EBA test for a variable in  $F$  states that if the lower extreme bound for  $\beta_F$  – i.e., the lowest value for  $\beta_F$  minus two standard deviations – is negative, while the upper extreme bound for  $\beta_F$  – i.e., the highest value for  $\beta_F$  plus two standard deviations – is positive, the variable  $F$  is *not* robustly related to  $Y$ .

Sala-i-Martin (1997) argues that this testing criterion is far too strong for any variable to ever pass it. If the distribution of the parameter of interest has both positive and negative support, then a researcher is bound to find at least one regression model for which the estimated coefficient changes sign if enough regressions are run. Consequently, in what follows we not only report

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<sup>22</sup>For details see Knack and Keefer (1995).

the extreme bounds, but also the percentage of the regressions in which the coefficient of the variable  $F$  is statistically different from zero. Moreover, instead of only analysing the extreme bounds of the estimates of the coefficient of a particular variable, we follow Sala-i-Martin's (1997) recommended procedure and analyse the entire distribution. Accordingly, we also report the unweighted parameter estimate of  $\beta_F$  and its standard error, as well as the unweighted cumulative distribution function,  $CDF(0)$ . The latter represents the proportion of the cumulative distribution function lying on each side of zero.  $CDF(0)$  indicates the larger of the areas under the density function either above or below zero, i.e., whether this happens to be  $CDF(0)$  or  $1 - CDF(0)$ . So  $CDF(0)$  always lies between 0.5 and 1.0. However, in contrast to Sala-i-Martin, we use the unweighted, instead of the weighted,  $CDF(0)$ .<sup>23</sup>

Another objection to EBA is that the initial partition of variables in the  $M$  and in the  $Z$  vector is likely to be arbitrary. However, as pointed out by Temple (2000), there is no reason why standard model selection procedures (such as testing down from a general specification) cannot be used in advance to identify variables that are particularly relevant. Furthermore, some variables are included in the large majority of studies and are by now common in this branch of the literature.

In our view, the inclusion of  $LGDP$  in the  $M$  vector is the only non-contentious inclusion as a regressor. In the literature on the environmental Kuznets curve the relationship between GDP per capita and environmental quality has been widely discussed. Therefore, this variable may also play an important role in determining the stringency of environmental policy. While it is tempting to include our central variables ( $INDSHEMP$  and  $INEQUAL$ ) in the  $M$  matrix, we are conscious of not prejudging the importance of our model and the outcome of the EBA.

### 3.3.3 The results

As a preliminary to the EBA, we ran a first regression using  $LGDP$  as well as our central variables and conducted specification tests to test whether we have to correct for country- and/or time-specific effects in our panel setup. As a result of these tests we include random country-effects in all equations.<sup>24</sup>

Table 4 depicts the results of the EBA.<sup>25</sup> The criterion for considering a variable to be robustly

<sup>23</sup>Sala-i-Martin (1997) proposes using the integrated likelihood to construct a weighted  $CDF(0)$ . However, missing observations for some of the variables poses a problem. Sturm and de Haan (2002) show that the goodness-of-fit measure may not be a good indicator of the probability that a model is the true model and that the weights constructed in this way are not invariant to linear transformations of the dependent variable. Hence, changing scales could result in different outcomes and conclusions. We therefore employ the unweighted version.

<sup>24</sup>For readability, the results of the specification test of the EBA are not shown. (However, they are available from the authors upon request).

<sup>25</sup>Since there are substantial differences in the number of observations for each variable, which potentially could influence our results, we opt to restrict our sample based on our inequality measure and hence ensure a more

related to stringency is the  $CDF(0)$  value. Sala-i-Martin (1997) suggested considering a variable to be robust if the  $CDF(0)$  criterion is greater than 0.90. Instead we follow Sturm and de Haan’s (2005) proposal to use a stricter threshold value of 0.95, due to the two-sided nature of the test.

– Table 4 here –

Turning to the results of the EBA we see that real GDP per capita ( $LGDP$ ) is robustly and positively linked to the level of stringency.<sup>26</sup> This result is also found in the existing empirical literature. Therefore, to some extent it resolves the potential theoretical ambiguity which Congleton (1992) highlighted.<sup>27</sup>

We now turn to the extended model. Here each of the variables takes the role of the  $F$  vector once with the other 11 variables used in 175 combinations to test the robustness of this particular variable. The variable  $LEFT$ , representing a left-wing chief executive, is negatively related to stringency. A left-wing executive traditionally cares for the interest of industrial workers and might therefore be reluctant to increase environmental stringency. The share of the urbanised population ( $URBAN$ ) has a negative relationship with stringency. Hence, citizens living in urban areas tolerate lower levels of environmental stringency.

From the viewpoint of our simple model, the most important findings of our analysis are the results for  $INDSHEMP$  and  $INEQUAL$ . According to the EBA the former variable is robustly negatively related to the measure of stringency. Therefore, just as our theory suggests, a declining blue-collar labour force is associated with diminished blue-collar political power and leads to more stringent environmental regulations. Our result for inequality is the more novel finding, however. Greater dispersion in incomes is associated with a more, and *not* less, stringent environmental policy. While this finding stands in stark contrast to previous research, it is consistent with our model. All other variables that are proposed in the literature, as being an influencing factor for the stringency of environmental agenda setting, clearly fail to meet the robustness criterion of a  $CDF(0)$  value above 0.95. This is another major finding of the EBA, although obviously a rather negative one with respect to the extant literature.

In order to evaluate the relevance of the variables we also estimate the magnitude of the impact that all variables have on the policy stringency measure. We do this by calculating the effect that a shock of one standard deviation of each variable has on  $LREGS$ . We therefore multiply the

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homogeneous sample.

<sup>26</sup>This result is based on 231 regressions.

<sup>27</sup>We also tested for a potential non-linear relationship by including the squared term into the model. However, this is not supported by the data.

average EBA coefficient with the standard deviation of the respective variable and rank them in descending order according to absolute value.<sup>28</sup> The resulting ranking is included in table 4 in the column “Impact Rank”. The five variables that exhibit a robust relation with the measure of stringency are among the six variables which have the biggest impact on the dependent variable. In addition we report the histograms of the coefficients of our two central variables. Figure 2 reveals that the estimated coefficients of the key variables are distributed close to their respective means and that there are no major outliers.

– Figure 2 here –

Concerning the robustness of our results we use the five variables that the EBA suggests are robustly linked to environmental stringency and estimate our final model. The results are contained in table 5.

– Table 5 here –

Based on purely statistical criteria, it is the preferred model. Again, the specification tests lead us to include country-specific random effects.<sup>29</sup> However, we also present the results when adding time-specific effects as a further robustness check.<sup>30</sup> Potentially worrisome is the relatively small number of observations (due to the list-wise deletion of missing observations on key variables). In order to test whether our results are driven by the small sample we linearly interpolate our inequality measure to create more observations. The results of these estimations are summarised in table 6. Except for *URBAN*, in the case of country- and time-specific fixed effects all variables are robust to the estimation technique, the inclusion of time effects and the sample size, i.e., they are statistically significant at conventional levels. Overall, the results of the EBA are reinforced.

– Table 6 here –

To summarise, there is obviously strong support for the argument forwarded in this paper. Namely, that declining economic significance is associated with a decline in political significance. Both of these factors reinforce one another and lead to more stringent environmental regulation.

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<sup>28</sup>Since the estimation results include country-specific random effects we de-mean all variables. Failing to do so could seriously bias the results since the country-specific effects that were already taken into account would again contribute to the result.

<sup>29</sup>For comparison, table 5 also contains the results for country-specific fixed effects.

<sup>30</sup>We also estimated model specifications that included a time trend. There were no changes to the results.

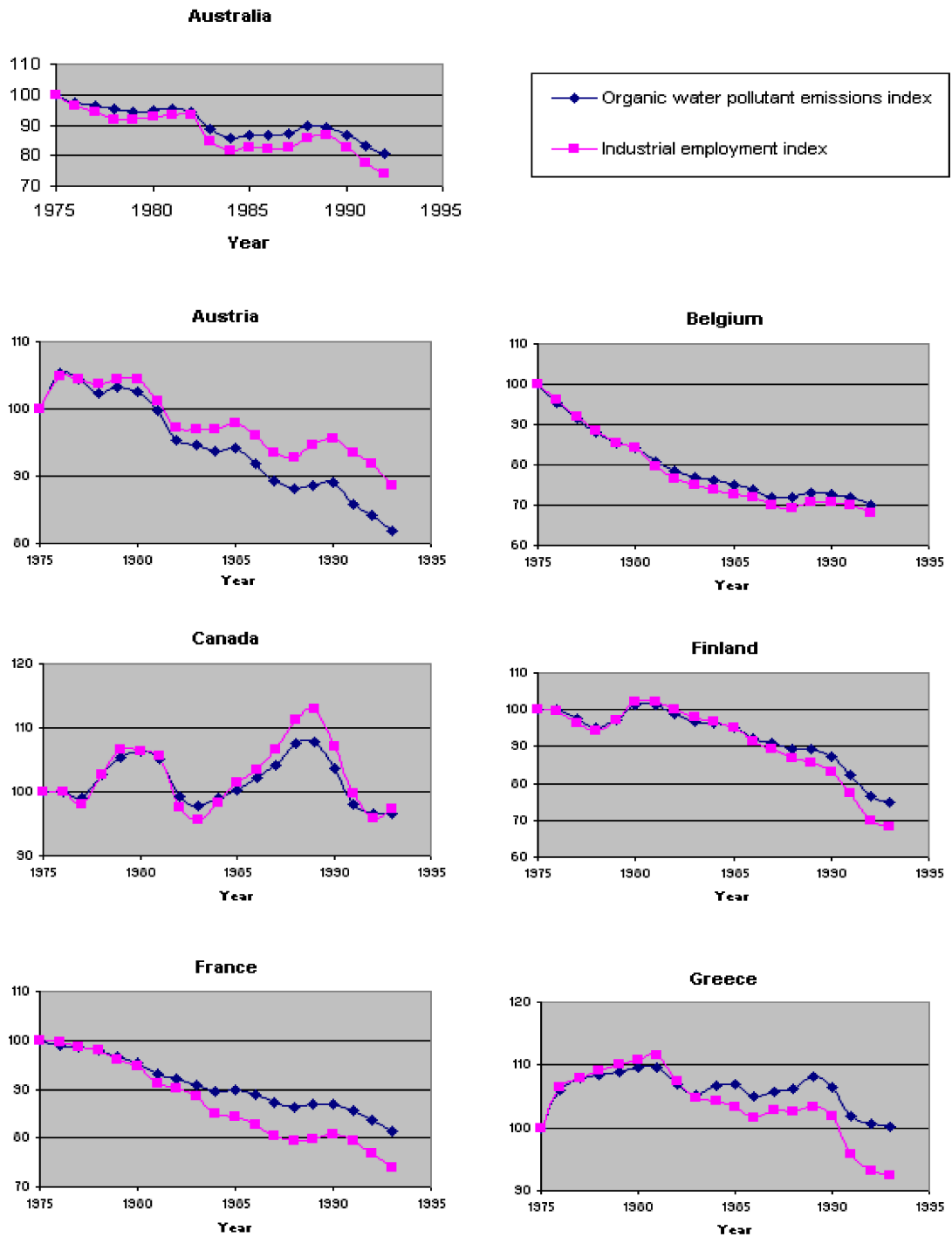
## 4 Conclusion

Our paper has emphasised that political and economic considerations interact to help explain the observed relationship between measures of economic development and environmental quality. Deindustrialisation, falling real incomes of production workers and a greater dispersion of income are increasingly prominent features in many industrialised countries. From a political economy perspective, such features can also explain observed environmental policies.

When the social and economic consequences of either high unemployment or falling incomes in manufacturing industries are high, policy-makers may be tempted to ease environmental regulations. Symmetrically, as deindustrialisation proceeds, as reflected by declining industrial employment and the falling wages and incomes for workers in basic manufacturing and pollution-intensive industries, environmental stringency increases. That is, as these sectors of the economy become less important economically, they are also likely to carry less weight politically. Consequently, a regulator optimally increases the stringency of environmental regulations. The argument is simple and straightforward. Dynamic comparative advantages dictate that mature, developed economies shift resources away from basic manufacturing activities. Environmental policy simply reinforces this movement. Political action and behaviour such as this provides an explanation for the inverted U-shape of the environmental Kuznets curve.

To some readers, the argument developed in this paper may seem overly optimistic from the point of the view of the environment and overly cynical from a social equity perspective. The risk of being over-cynical is particularly acute for those who believe that a sense of social justice should prevail during times of rapid deindustrialisation and falling blue-collar worker incomes. In turn, the social and political pressures may be thought to help override the demand for increased regulatory stringency. If this were in fact the case, it would be expected that environmental policies are least stringent in those industrialised and democratic countries in which income inequality is greatest. The evidence presented in the paper is consistent with the exact opposite view. That is, countries with the strictest environmental standards tend to be those with the greatest dispersion in their incomes.

Figure 1: Water Pollution and Industrial Employment



Source: Authors' calculations based on data from Hettige *et al.* (1998).

Figure 1: Water Pollution and Industrial Employment-continued

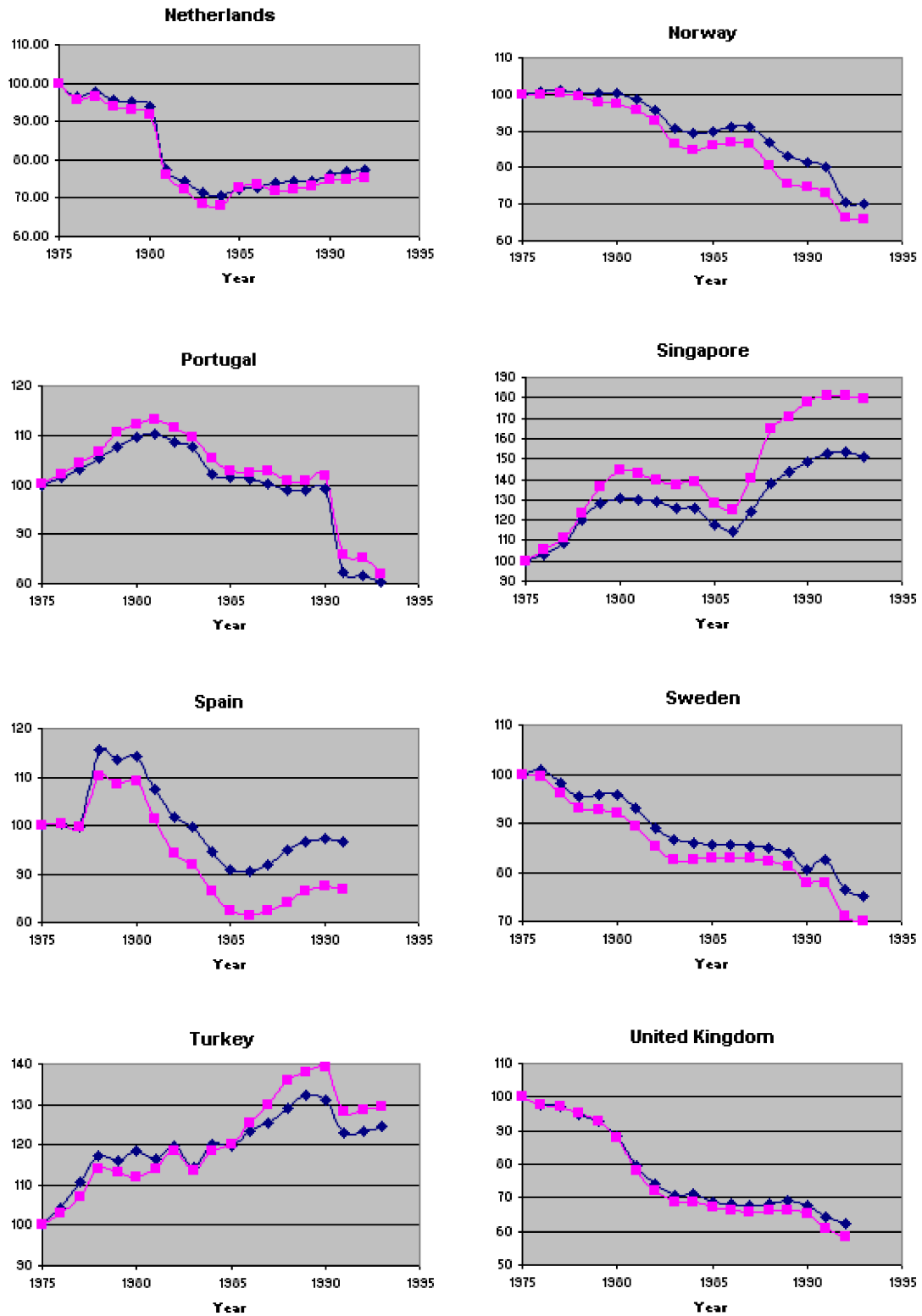




Figure 1: Water Pollution and Industrial Employment-continued

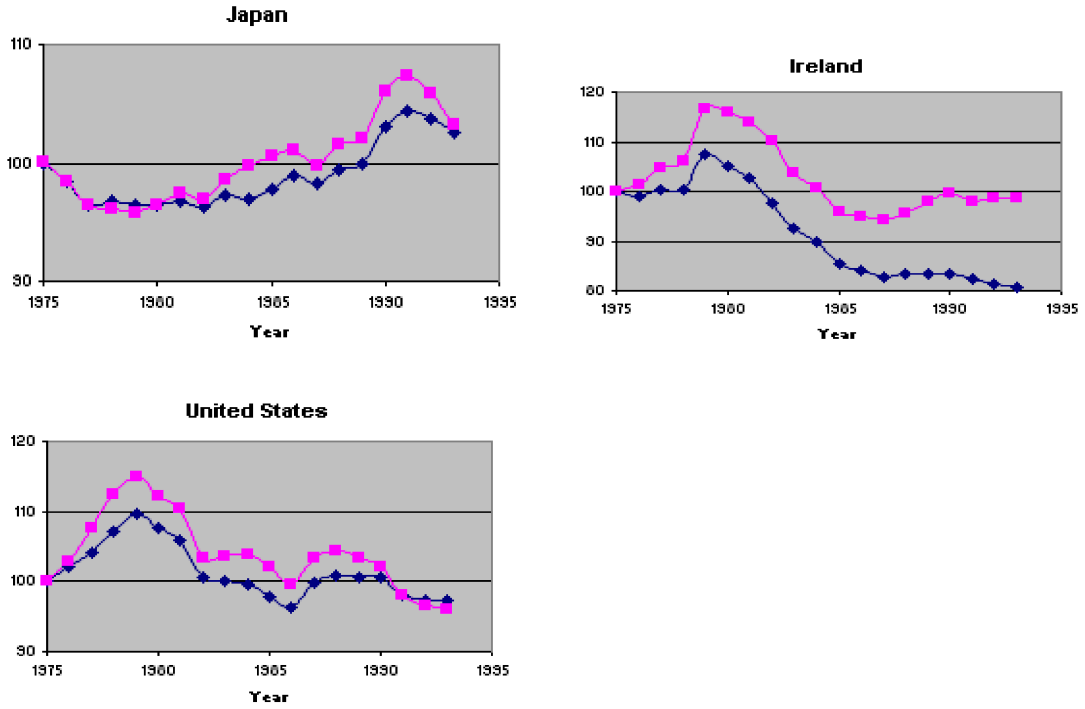
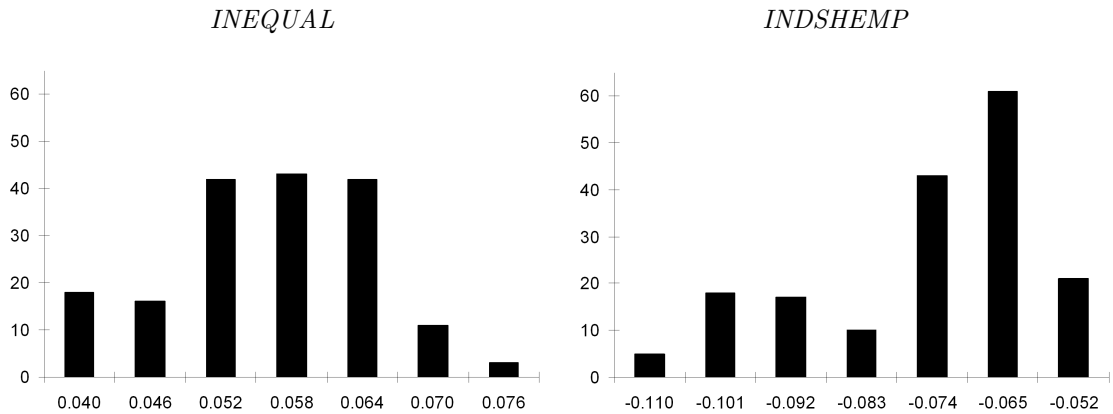


Figure 2: Histograms of the EBA coefficients



Note: The frequency distributions summarise the coefficients of the 175 regressions of the EBA for the respective variable. The number beneath each bar indicates the upper bound of the bin.

Table 1: Total taxes as per cent of end-user price for automotive fuels, 1994

Country	Household Use (1)	Industrial Use (2)	Ratio (2)÷(1)	Union co- ordination index
Denmark	68.0	41.5	0.61	3
Sweden	76.5	48.3	0.63	3
Norway	67.3	46.0	0.68	3
Austria	63.9	49.1	0.77	3
Belgium	74.2	57.3	0.77	2
Netherlands	75.9	59.7	0.79	2
France	80.8	65.1	0.81	2
Portugal	73.5	59.4	0.81	2
Germany	76.9	62.5	0.81	2
Spain	68.6	56.9	0.83	2
Canada	50.0	41.6	0.83	1
Italy	76.1	65.1	0.86	2
U.K.	73.5	63.6	0.87	1
Switzerland	71.3	68.9	0.97	1
U.S.A.	34.4	39.6	1.15	1

Sources: Columns (1) and (2): OECD (1995), table 2, p.48. Last column: Layard *et al.* (1994), table 6, p.78. values represent: 3 = High (National coordination); 2 = Intermediate; 1 = Low (firm-level or uncoordinated).

Table 2: Variables – definitions, sources and hypotheses

Variable	Source	Description	Sign	Proposed by
LREGS	Octel (1982-1992)	Log of lead content of gasoline, multiplied by (-1)		Damania <i>et al.</i> (2003)
CORRUPT	ICRG	“Government Honesty”, higher values indicate less corruption	+	Damania <i>et al.</i> (2003)
DICT	Beck <i>et al.</i> (1999)	Dummy variable for dictatorship (executive index of electoral competitiveness < 3)	-	Congleton (1992)
FDIGDP	WDI (2003)	Net inflows of foreign direct investment (% of GDP)	+	Cole <i>et al.</i> (2006)
GLOBAL	Dreher (2006)	KOF Index of Globalisation	+	This paper
INDSHEMP	WDI (2003)	Employment in industry sector (% of total employment)	-	Damania <i>et al.</i> (2003)
INEQUAL	Francois and Rojas-Romagosa (2005)	Gini coefficient – household income	+	This paper
LEFT	Beck <i>et al.</i> (1999)	Dummy variable for the chief executive’s party being left-wing	?	Neumayer (2003)
LGDPPC	WDI (2003)	Log of GDP per capita (in constant 1995 \$US)	?	Congleton (1992)
LPOPDENS	WDI (2003)	Log of population per hectare	?	Congleton (1992)
POLFREE	FHI (2005)	Average of “Civil Liberties” and “Political Rights”	+	Congleton (1992)
TRADE	WDI (2003)	Trade intensity ((imports + exports)/GDP)	+	Damania <i>et al.</i> (2003)
URBAN	WDI (2003)	Urban population (% of total)	+	Damania <i>et al.</i> (2003)

Note: ‘Sign’ refers to the expected sign of the variable according to the literature ‘+/-’ indicates a positive/negative sign while ‘?’ represents an *a priori* indeterminate effect.

Table 3: Variables – descriptive statistics and correlation matrix

	<b>Mean</b>	<b>S.D.</b>	<b>(1)</b>	<b>(2)</b>	<b>(3)</b>	<b>(4)</b>	<b>(5)</b>	<b>(6)</b>	<b>(7)</b>	<b>(8)</b>	<b>(9)</b>	<b>(10)</b>	<b>(11)</b>	<b>(12)</b>	<b>(13)</b>
<b>(1)</b> LREGS	-0.37	0.95	341	0.418	-0.225	-0.047	0.528	0.349	-0.439	-0.110	0.490	0.285	-0.396	0.101	0.383
<b>(2)</b> CORRUPT	3.49	1.55	328	328	-0.360	0.190	0.769	0.533	-0.381	0.130	0.707	-0.208	-0.540	0.340	0.608
<b>(3)</b> DICT	0.22	0.42	341	328	341	-0.138	-0.357	-0.264	0.471	-0.167	-0.537	-0.062	0.635	-0.172	-0.372
<b>(4)</b> FDIGDP	1.17	1.71	324	311	324	324	0.250	0.127	0.087	0.068	0.131	-0.240	-0.182	0.504	0.060
<b>(5)</b> GLOBAL	1.91	0.74	310	310	310	293	310	0.502	-0.582	0.053	0.865	-0.013	-0.706	0.438	0.771
<b>(6)</b> INDSHEMP	26.18	8.03	194	194	194	192	193	194	-0.399	0.180	0.724	0.037	-0.527	0.096	0.565
<b>(7)</b> INEQUAL	43.37	10.49	57	57	57	57	57	52	57	-0.159	-0.428	-0.365	0.516	-0.429	-0.037
<b>(8)</b> LEFT	0.36	0.48	332	319	332	315	302	188	55	332	-0.104	-0.193	-0.040	-0.143	-0.212
<b>(9)</b> LGDPPC	7.78	1.76	335	322	335	324	304	194	57	326	335	0.017	-0.797	0.327	0.869
<b>(10)</b> LPOPDENS	-0.45	1.36	330	319	330	313	310	194	57	322	324	330	-0.110	0.187	-0.058
<b>(11)</b> POLFREE	3.32	1.92	341	328	341	324	310	194	57	332	335	330	341	-0.248	-0.670
<b>(12)</b> TRADE	49.28	27.30	333	320	333	322	302	194	57	324	333	322	333	333	0.235
<b>(13)</b> URBAN	52.34	25.74	341	328	341	324	310	194	57	332	335	330	341	333	341

Note: The first two columns report the mean and the standard deviation (S.D.) of each series; the upper-right part of the remaining table reports correlation coefficients, the main diagonal gives the number of observations for each variable, while the lower left shows the number of observations used to calculate the correlation coefficients.

Table 4: EBA results (dependent variable: LREGS)

Variable	Lower Bound	Upper Bound	%Sign.	Unwght. CDF(0)	Unwght. $\beta$	Std. Error	Impact Rank
<i>Base Model</i>							
LGDPPC	-0.482	3.861	83.55	0.9878	1.513	0.445	3
<i>Extended Model</i>							
INDSHEMP	-0.189	0.023	95.43	0.9813	-0.079	0.034	4
LEFT	-1.398	0.150	93.14	0.9779	-0.626	0.285	1
URBAN	-0.193	0.035	74.29	0.9651	-0.069	0.035	2
INEQUAL	-0.029	0.130	70.86	0.9575	0.053	0.029	6
LPOPDENS	-1.115	3.579	34.29	0.9194	0.753	0.508	11
POLFREE	-0.292	0.583	33.14	0.8756	0.121	0.096	7
FDIGDP	-0.112	0.275	9.71	0.8625	0.077	0.065	8
GLOBAL	-1.685	2.893	40.57	0.8501	0.866	0.665	5
TRADE	-0.044	0.026	10.86	0.8066	-0.010	0.010	9
DICT	-1.272	1.157	14.29	0.7807	0.218	0.277	10
CORRUPT	-0.648	0.438	2.86	0.5707	-0.046	0.145	12

Note: Results based on 231 (base model) and 175 (extended model) regressions, respectively, using country-specific random effects. ‘%Sign’ refers to the percentage of the regressions in which the respective variable is significant at the 10% significance level. ‘Impact Rank’ lists the variables in descending order according to the impact of a one standard deviation shock. The standard deviation is calculated after de-meaning each variable to correct for country-specific effects. Variables are sorted according to the CDF(0) criterion.

Table 5: Final model (dependent variable: LREGS)

	Random effects	Fixed effects
LGDPPC	3.652*** (0.519)	4.363*** (0.629)
INDSHEMP	-0.098*** (0.028)	-0.132*** (0.033)
LEFT	-0.713*** (0.232)	-0.756*** (0.253)
URBAN	-0.107*** (0.035)	-0.080* (0.049)
INEQUAL	0.062*** (0.024)	0.063** (0.027)
Constant	-24.221*** (3.472)	-
Observations	50	50
R-squared	0.944	0.925
Hausman test ( $H_0$ : random effects, $H_1$ : fixed effects)	2.863	-
F-test (significance of country-specific random effects)	4.573***	-
LR-test ( $H_0$ : pooled OLS, $H_1$ : country-specific fixed effects)	-	107.7***

Note: \*/\*\*/\*\* indicates significance at the 10%/5%/1% level. Both regressions contain country-specific effects. Standard errors are displayed in parentheses.

Table 6: Robustness checks (dependent variable: LREGS)

	(a)	(b)	(c)	(d)
LGDPPC	3.724*** (0.533)	4.821** (2.141)	3.405*** (0.479)	1.883*** (0.466)
INDSHEMP	-0.100*** (0.030)	-0.128*** (0.046)	-0.093*** (0.030)	-0.065** (0.028)
LEFT	-0.702*** (0.239)	-0.970*** (0.344)	-0.378* (0.232)	-0.478** (0.234)
URBAN	-0.106*** (0.036)	-0.116 (0.076)	-0.100*** (0.037)	-0.064** (0.029)
INEQUAL	0.063** (0.025)	0.057* (0.031)	-	-
INEQUAL-interpolated	-	-	0.060** (0.029)	0.041* (0.026)
Constant	-24.872*** (3.595)	-	-0.378*** (0.232)	0.041*** (0.026)
Observations	50	50	102	102
R-squared	0.941	0.913	0.921	0.934
Hausman test (H <sub>0</sub> : random effects, H <sub>1</sub> : fixed effects)	0.841	-	2.965	0.891
F-test (sign. of country-specific random effects)	4.812***	-	12.960***	18.815***
F-test (sign. of time-specific random effects)	1.329	-	-	4.840***
LR-test (H <sub>0</sub> : pooled OLS, H <sub>1</sub> : country and time-specific fixed effects)	-	120.0***	-	-

Note: \*/\*\*/\*\* indicates significance at the 10%/5%/1% level. Columns (a) and (d) contain country- and time-specific random effects, (b) includes country- and time-specific fixed effects, and (c) incorporates country-specific random effects. Standard errors are displayed in parentheses.

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## A Derivation of the Euler Equation

The first-order condition for the maximisation of Bellman's Equation is

$$\theta^b \frac{\partial E_t U_t^b}{\partial s_t} + \theta^w \frac{\partial E_t U_t^w}{\partial s_t} + \beta E_t V'_{t+1}(p_{t+1}) = 0.$$

Rearranging and simplifying we have

$$\theta^b f_{s_t} \Delta_t + \beta E_t V'_{t+1}(p_{t+1}) = 0, \quad (\text{A-1})$$

where  $\Delta_t = U(y_t^{b,e}, p_t) - U(y_t^{b,u}, p_t)$ . Differentiating the value function yields

$$V'_t(p_t) = \theta^b \frac{\partial E_t U_t^b}{\partial p_t} + \theta^w \frac{\partial E_t U_t^w}{\partial p_t} + (1 - \delta) \beta E_t V'_{t+1}(p_{t+1}).$$

After simplifying we have

$$V'_t(p_t) = \theta^b l_t^b U_{p_t}^{b,e} + \theta^b (1 - l_t^b) U_{p_t}^{b,u} + \theta^w U_{p_t}^w + (1 - \delta) \beta E_t V'_{t+1}(p_{t+1}). \quad (\text{A-2})$$

Substituting (A-2) into (A-1) yields

$$\theta^b f_{s_t} \Delta_t + (1 - \delta)^{-1} (V'_t(p_t) - \theta^b l_t^b U_{p_t}^{b,e} - \theta^b (1 - l_t^b) U_{p_t}^{b,u} - \theta^w U_{p_t}^w) = 0, \quad \text{or}$$

$$V'_t(p_t) = (\theta^b l_t^b U_{p_t}^{b,e} + \theta^b (1 - l_t^b) U_{p_t}^{b,u} + \theta^w U_{p_t}^w) - (1 - \delta) \theta^b f_{s_t} \Delta_t. \quad (\text{A-3})$$

Letting  $\varepsilon_{t+1} = E_t V'_{t+1} - V'_{t+1}$  be the one-period ahead forecast error, then substituting (A-3) into (A-1) yields the Euler equation (i.e., equation (4) in the text).