# SUMMARY

*The relationship between a firm's size and its productivity level varies considerably across OECD countries, suggesting that some countries are more successful at channelling resources to high productivity firms than others. In this paper, we examine the extent to which these differences depend on regulations affecting product, labour and credit markets, and assess their relevance for aggregate productivity. To this purpose, we exploit a decomposition of industry productivity into a moment of the firm productivity distribution (the unweighted mean), and a moment of the joint distribution with firm size (the covariance between productivity and market shares –* allocative efficiency*). We apply such decomposition to a cross section of more than 800 country-industry cells and estimate the relevance of regulation policies for each of the two terms exploiting cross-industry differences in exposure to the policy. Our results suggest that there is an economically and statistically robust negative relationship between policy-induced frictions and productivity, though the specific channel depends on the policy considered. In the case of employment protection legislation, product market regulations (including barriers to entry and bankruptcy legislation) and restrictions on foreign direct investment, this is largely traceable to the worsening of allocative efficiency (i.e. a lower correspondence between a firm's size and its productivity level). By contrast, the adverse impact of financial market under-development on aggregate productivity tends to arise through shifts in the firm productivity distribution (i.e. a lower unweighted mean). Furthermore, stringent regulations are more disruptive to resource allocation in more innovative sectors.*

# Public policy and resource allocation: Evidence from firms in OECD countries

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## 1. INTRODUCTION

Differences in income per capita across OECD countries mainly reflect large and persistent differences in labour productivity. At the same time, boosting labour productivity growth is an urgent policy priority, especially in those countries where declining working age populations pose a major headwind to future improvements in living standards. Given existing constraints to fiscal expansion, however, the policy options available to governments appear to be narrowing. In this context, policy makers are increasingly looking to structural (supply-side) reforms to improve productivity performance and evaluating the contribution of such reforms to economic performance thus represents a fruitful area for economic policy research.

Cross-country differences in aggregate-level productivity performance are increasingly being linked to the widespread asymmetry and heterogeneity in firm performance within sectors. Indeed, the distribution of firm productivity and size is typically not clustered around the mean (as would be the case with a normal distribution) but is instead characterised by many below-average performers and a smaller number of star performers (Haltiwanger, 2011). Moreover, the degree of heterogeneity is striking: even within

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narrowly defined industries in the United States, for example, the 90th percentile of the productivity distribution makes, on average, about twice as much output with the same measured inputs as a plant at the 10th percentile (Syverson 2004), while firms in the top quartile of the within industry size distribution are on average 80 times larger than firms in the bottom quartile (Bartelsman et al., 2009). Similar, though not identical, patterns are also observable in other developed economies, while the heterogeneity in firm performance in China and India is even more pervasive (Hsieh and Klenow, 2009).

At any point in time, differences in aggregate labour productivity will reflect: (i) the productivity distribution of firms (*i.e.* the fraction of "better" relative to "worse" firms); and (ii) the extent to which, all else equal, it is the more productive firms that command a larger share of aggregate employment (*i.e.* allocative efficiency), which will be the outcome of the shift in resources across firms in previous periods. While the former component has been the subject of much research, reflecting a number of within-firm factors (*e.g.* managerial quality; intangible assets; innovation strategies; idiosyncratic shocks), economic researchers are increasingly linking the pattern of resource allocation within sectors to aggregate-level economic performance.

Indeed, estimates of allocative efficiency (AE) – based on the within-industry covariance between a firm's size and its productivity level – vary considerably across OECD countries, suggesting that some of them are more successful at channelling resources to high productivity firms than others. For example, in the United States, manufacturing sector labour productivity is 50% higher due to the actual allocation of employment across firms, compared to a hypothetical situation where labour is uniformly allocated across firms, irrespective of their productivity. While a similar pattern holds for some countries of Northern Europe, it turns out that allocative efficiency is considerably lower in other OECD economies, particularly those of Southern Europe.

The working hypothesis in this paper is that these apparent differences in the efficiency of resource allocation are closely related to regulations affecting product, labour and credit markets. To the extent they raise the costs of workforce adjustments, for example, dismissal costs are likely to induce firms not to hire workers even if their marginal product exceeds market wage, and/or to retain workers whose wage exceeds their productivity. Similarly, if product market regulations restrict the extent of competition through higher barriers to entry, there may be less pressure on incumbent firms to allocate resources efficiently. While the link between market reforms and productivity has received much attention, there is very little direct cross-country evidence on whether the productivity gains associated with reform episodes are realised through the channel of more efficient resource allocation.

Preliminary analysis suggests that the cross-country differences in allocative efficiency cited above are strongly and significantly correlated with a number of structural policies of interest. However, these correlations might be driven by a large amount of (country specific) unobserved characteristics; for example, allocative efficiency might be low if large inefficient firms (a symptom of low AE) were successful at lobbying governments for protection measures. To circumvent these problems, we exploit the idea that regulations are more binding for some industries than others, an approach in the spirit of

Rajan and Zingales (1998). This technique has been employed in many policy-oriented empirical works, as it allows one to assess the role of countrywide institutions, while controlling for time invariant country- (and industry-) specific factors.[1]

Looking across a sample of private non-farm sectors, we find evidence that more stringent product and labour market regulations adversely affect AE. More precisely, we find that higher barriers to firm entry and creditor-friendly bankruptcy legislation tend to disproportionately lower AE in industries characterised by high firm turnover relative to low turnover industries. Similarly, tighter employment protection legislation is found to disproportionately lower the efficiency of employment allocation in high layoff and high turnover industries. These results are robust to a variety of robustness tests including instrumental variables regressions to control for the possible endogeneity of policies affecting product and labour markets. Indeed, our estimates may actually understate the overall impact of policy-induced distortions on resource allocation to the extent that they do not account for the impacts of regulation on resource flows between sectors, which are likely to reinforce the within-sector effects that we identify.

Additional exercises provide a sensitivity check on some of the policy conclusions and further insight into the possible channels through which policy distortions affect AE. First, stringent product and labour market regulations and bankruptcy legislation are more disruptive to AE in more innovative sectors, which are likely to be subject to greater technological change and thus place a high option value on flexibility. Second, the impact of product market regulations on resource allocation is confirmed by analysis of industry specific measures of regulation for a sub-sample of service industries. While these results are based on a relatively small number of sectors, they also suggest that restrictions on foreign direct investment (FDI) reduce the efficiency of resource allocation.

While restrictions to competition in finance and low financial development do not seem to be related to AE, both are associated with lower unweighted average industry productivity than otherwise. In other words, in countries with low financial development (or high banking regulation) the first moment of the productivity distribution tends to be disproportionately lower in industries more dependent on external financing. These findings suggest that less effective financial markets affect aggregate productivity by shifting the distribution of active firms towards lower levels of productivity, rather than by altering the allocation of employment across existing productive units.

The paper is structured as follows: Section 2 presents evidence on cross-country differences in the efficiency of resource allocation, while Section 3 explores the potential for public policies to influence AE and presents some preliminary evidence for this hypothesis. In Section 4, we describe our empirical approach to identify the impact of policies on AE while Section 5 discusses the econometric results. In Section 6, we subject the core results to a battery of robustness tests while Section 7 offers some concluding thoughts.

[1] For the case of Product Market Regulation, see for example Klapper et al. (2006), Fisman and Sarria Allende (2010), Ciccone and Papaioannou (2007); for Employment Protection regulation, Micco and Pagés (2006), Bassanini et al. (2009) and Cingano et al. (2010).

## 2. THE EFFICIENCY OF RESOURCE ALLOCATION VARIES ACROSS OECD COUNTRIES

Cross-country differences in income per capita mainly reflect large and persistent differences in productivity, while those in labour utilisation play a more modest role (Figure 1). An emerging literature links these differences in aggregate productivity to the misallocation of resources across firms, which is defined as a departure from the optimal allocation of resources (or static allocative efficiency). In core models of heterogeneous firms (see Lucas, 1978; Melitz, 2003), the optimal allocation features a positive relationship between firms' productivity and their size: allocative efficiency requires resources to be allocated to their highest valued use, which implies that at any point in time, the most productive firms are also the largest.[2] Accordingly, recent firm-level empirical studies exploit the firm-level covariance between productivity and size as a robust, model-based measure to assess the pattern and relevance of resource misallocation across countries and over time (see Bartelsman et al., 2009, 2013; Eslava et. al. 2010; Deng et al., 2007).

These studies reveal that there is considerable variation in static allocative efficiency – i.e. the strength of the link between firm productivity and size – with important implications for aggregate productivity. Using high quality firm level data for eight countries over the 1990s, Bartelsman et al., (2013) show that the covariance between firm size and firm productivity in the United States is higher than in Western European and particularly Eastern European economies.[3] Nevertheless, significant increases in this covariance term have been observed over time in Eastern European countries, China and Colombia, and such improvements have generally coincided with episodes of market-oriented reform (Bartelsman et al., 2009; Eslava et. al. 2010; Deng et al., 2007). At the same time, these differences in resource allocation carry important consequences for aggregate performance: estimates suggest that if China and India were able to align their efficiency of resource allocation to that observed in the United States, manufacturing productivity could rise by 30-50% in China and 40-60% in India (Hsieh and Klenow, 2009).

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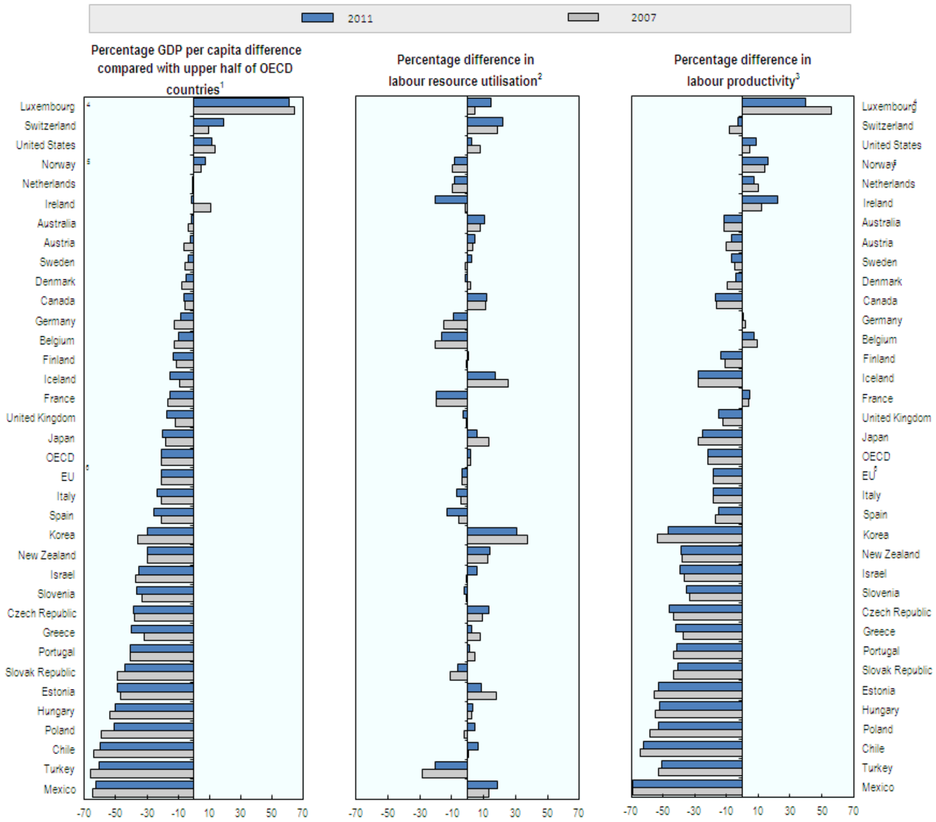
[2] This result is obtained in models featuring different assumptions regarding the available technologies and market structure. In "span-of-control" models following Lucas' (1978), markets are perfectly competitive and individual firms have decreasing returns to scale technologies (this ensures that the most productive producer does take over the market). But the same is true for models of monopolistic competition where technologies are constant or increasing return to scale (due to the presence of fixed costs of entry). See Hopenhayn (2011) for a discussion.

[3] The countries included in this analysis are: France, Germany, Hungary, the Netherlands, Romania, Slovenia, United Kingdom and the United States.




1. Compared to the simple average of the 17 OECD countries with highest GDP per capita in 2011 and 2007, based on 2011 and 2007 purchasing power parities (PPPs). The sum of the percentage difference in labour resource utilisation and labour productivity do not add up exactly to the GDP per capita difference since the decomposition is multiplicative.
2. Labour resource utilisation is measured as the total number of hours worked per capita.
3. Labour productivity is measured as GDP per hour worked.
4. In the case of Luxembourg, the population is augmented by the number of cross-border workers in order to take into account their contribution to GDP.
5. Data refer to GDP for mainland Norway which excludes petroleum production and shipping. While total GDP overestimates the sustainable income potential, mainland GDP slightly underestimates it since returns on the financial assets held by the petroleum fund abroad are not included.
6. Average of European Union countries in the OECD.
*Source*: OECD National Accounts Statistics (Database); OECD (2012), OECD Economic Outlook No. 92: Statistics and Projections (Database); OECD, Employment Outlook (Database).

**Figure 1. Large cross-country differences in income per capita are mostly accounted for by productivity gaps**

Following the relevant literature, we compute allocative efficiency exploiting the cross sectional decomposition of productivity developed by Olley and Pakes (1996). They observed that an index of productivity of industry $j$, defined as the weighted average of firm-level productivity ($P_j = \sum_{i \in j} \theta_i P_i$), can be written as:

$$\sum_{i \in j} \theta_i P_i = \bar{P}_j + \sum_{i \in j} (\theta_i - \bar{\theta}_j)(P_i - \bar{P}_j) \qquad (1)$$

where $\bar{P}_j = 1/N_j \sum_{i \in j} P_i$ is the unweighted firm productivity mean, $\theta_i$ is a measure of the relative size of each firm (e.g. the firm employment share) and $\bar{\theta}_j$ is the average share at the industry level. Hence, aggregate productivity ($P_j$) can be decomposed into a moment of the firm productivity distribution (i.e. the unweighted mean) and a joint moment with the firm size distribution reflecting the extent to which firms with higher efficiency also have a larger relative size (i.e. the "Olley-Pakes covariance" term). This second term is the allocative efficiency (AE) measure used in our analysis.

Allocative efficiency can therefore simply be obtained as the difference between the weighted and the unweighted average of the firm productivity distribution ($AE_j = P_j - \bar{P}_j$). Suppose that firm employment shares are used as weights. Then, a positive value of this difference captures the fact that the allocation of employment across firms with different productivity levels is such to assign relatively larger weights to relatively more productive firms. Note that if employment was allocated randomly across firms, then the weighted and unweighted productivity averages would coincide ($P_j = \bar{P}_j$), so that $AE = 0$. Hence, AE being positive can also be interpreted as the increase in the industry productivity index $P_j$ that can be traced to the actual allocation of employment across firms within the industry relative to a baseline scenario in which employment is randomly allocated.[4]

We also follow the relevant literature by implementing (1) with a productivity measure ($P_i$) in logs (see Foster et al 2001; Bartelsman et al 2013). We restrict our analysis to three main productivity measures and associated weights to capture relative firm size: two measures of labour productivity with employment shares as weights and one measure of total factor productivity (TFP) with value-added shares as weights. One advantage of measuring firm productivity in logs is that the AE term can be read as the *percentage* (log point) *change* in the productivity index $P_j$ due to the actual allocation of employment across firms relative to a baseline (random allocation) scenario.[5] Notice that, because AE measures productivity differences relative to a benchmark (i.e. unweighted average productivity), changes in AE can be read as changes in productivity only if the benchmark is held constant. Accordingly, in our empirical analysis we will relate regulatory policies to each of the two components.[6]

Despite its increased popularity, the Olley-Pakes decomposition has several limitations. One that is potentially relevant for our analysis is that the productivity-size covariance term might reflect industry-specific characteristics. The strength of the (positive) productivity-size relationship predicted by models of heterogeneous firms will in fact depend on technological characteristics, such as the curvature of the production function, or the amount of the fixed costs to operate (see Hopenhayn, 2011, Bartelsman et al., 2013). Production function estimates suggest that production technologies differ, in particular, across industries. If this is the case, inference based on cross-industry comparisons of the allocative efficiency term might not just reflect the underlying patterns of resource allocation. Throughout our empirical analysis, however, we will account for

[4] Note also that while comparisons of the levels of either weighted or unweighted productivity across sectors or countries is problematic due to measurement issues (e.g. index number) focusing on the difference between the two implies that any confounding factor affecting both productivity measures is differenced out (Bartelsman et al., 2009).

[5] Measuring firm productivity in logs also implies that, as in recent studies of the consequences of misallocation based on models of monopolistic competition, the aggregate index $P_j$ is a geometric average of the firm productivity distribution.

[6] Cross sectional differences in the unweighted productivity mean are often thought of as capturing differences in the underlying firm productivity distributions (i.e. in the average "quality" of active firms).

industry-specific fixed-effects, meaning that our analysis will not exploit variation in allocative efficiency that reflects additive common industry characteristics. A second limitation is that, being static, such decomposition does not allow capturing the contribution of dynamic reallocation to aggregate productivity growth. The key mechanisms through which this process occurs are firm turnover (i.e. entry and exit) and shifts in resources across incumbent firms with different productivity levels. Within-country studies generally underline the productivity-enhancing role of such reallocation activity, highlighting in particular the positive contribution of net entry and of the rapid expansion of initially smaller, young firms.[7] Understanding whether and to what extent regulatory policies affect each of these mechanisms would therefore be extremely relevant. However, systematic cross-country evidence on the contribution of dynamic allocative efficiency to productivity growth is fraught with interpretational and measurement difficulties mainly related to the comparability across countries of the entry and exit of firms (see Box 1: the ORBIS database). For this reason, we focus on the relationship between firm productivity and size at a single point in time, and interpret this measure as representative of the extent to which resources are reallocated away from less productive to more productive uses over preceding time periods.

To explore cross-country differences in the efficiency of resource allocation, we use a harmonized firm level data set covering a cross-section of non-farm business industries – that is, industries 15-74 according to NACE Rev 1.1, excluding mining and financial intermediation – in 21 OECD countries for the year 2005 (data for the United States are also provided but this country is excluded from much of the analysis; see Section 4). After implementing some common procedures to address a number of measurement issues and enhance cross-country comparability (See Box 1 for details), data on labour productivity and employment are available from some 1.34 million firm-level observations, corresponding to 64,000 observations on average per country. From these firm-level data, indicators of AE and productivity at the 2-digit industry level are constructed, resulting in 834 country-industry observations or around 40 industry observations per country. See Table A1 in the Supplementary Appendix for a summary of the firm level data that underpins the analysis in the paper.

**Box 1: the ORBIS database**

Our analysis exploits cross-country firm-level data from ORBIS, a commercial database provided to the OECD by Bureau Van Dijk, which contains administrative data on tens of millions of firms worldwide. The financial and balance sheet information in ORBIS is initially collected by local Chambers of Commerce and in turn, is relayed to Bureau Van

[7] For example, Disney et al., (2003) show that for the United Kingdom this reallocation accounts for more than 80% of aggregate total factor productivity growth in the manufacturing sector. Baldwin and Gu (2006) for Canada find that this reallocation accounts for about 70% of aggregate labour productivity growth. Foster et al. (2006) find that entry and exit explain almost all labour productivity growth of the US retail sector. See Foster et al (2001) for a methodological discussion of the main approaches to measure the contribution of reallocation to productivity growth.

Dijk through some 40 different information providers (see Pinto Ribeiro *et al.*, 2010).

While representing a potentially useful tool to analyse the cross-country patterns in productivity, ORBIS has a number of drawbacks. The main issue relates to representativeness, with firms in certain industries and many smaller and younger firms typically under-represented. Accordingly, the ORBIS sample of firms was aligned with the distribution of the firm population as reflected in the Structural Demographic Business Statistics (SDBS) collected by the OECD and Eurostat, which is based on confidential national business registers.[8] This post-stratification procedure is of course based on the assumption that within each specific cell ORBIS firms are representative of the true population – an assumption that may be problematic if the nature of selection varies across countries.[9]

In order to maximise data coverage and to alleviate some measurement concerns, we mainly focus on an operating revenue turnover-based measure of labour productivity since not all firms report value added and capital and this problem tends to vary across countries. However, the use of a turnover-based measure of labour productivity does create some interpretation issues (see Section 6.2). Accordingly, we re-estimate our dependent variable using value-added based measures of labour productivity and total factor productivity (TFP), and show that our baseline estimates are largely robust to these alternative AE estimates, which of course are based on a smaller sample of countries and firms. The TFP estimates are obtained using a Solow residual technique using (2-digit NACE Rev 1.1) sectoral (country-specific) labour cost shares from the OECD STAN, and data on value added, employment and book capital from ORBIS. The capital stock is calculated using the Perpetual Inventory Method where real investment is calculated as the difference between the current and lagged book value of fixed tangible assets plus depreciation, deflated by country and industry specific investment deflators. Since we use industry-level deflators in absence of firm level price deflators, the measures of TFP (as well as those of labour productivity) are revenue based measures. It is also possible to obtain similar results using TFP estimates based on an OLS production function estimation approach, but we do not show these results for sake of brevity.

We exclude firms with one employee as well as firms in the top and bottom 1% of the labour productivity distribution from the sample – a relatively common data cleaning technique in the literature. Finally, in ORBIS it is not possible to accurately distinguish entry into the market from entry into the sample and exit from the market from exit from the sample. This prevents us from undertaking a dynamic decomposition of industry productivity growth that accounts for the contribution of entry and exit, and thus we

[8] The post-stratification procedure applies re-sampling weights based on the number of employees in each SDBS country-industry-size class cell to “scale-up” the number of ORBIS observations in each cell so that they match those observed in the SDBS (see Gal, 2013). For example, if SDBS employment is 30% higher than ORBIS employment in a given cell, then the 30% "extra" employment is obtained by drawing firms randomly from the pool of ORBIS firms, such that the "extra" firms will make up for the missing 30%.

[9] Moreover, post-stratification weights do not address the issue of how accurately are aggregates (such as allocative efficiency) measured when the underlying number of available units is small; this issue will be tackled empirically by weighting OLS regression estimates by the number of available observations in each country-industry cell.

focus on a decomposition of the level of labour productivity at a single point in time.

By way of introduction and for purely illustrative purposes, Figure 2 shows the distribution of our AE index across countries for the manufacturing sector in 2005. The estimates suggest that some countries are more successful at channelling resources to high productivity firms than others. Indeed, more productive firms are likely to account for a much larger share of manufacturing employment in the United States and some Nordic countries than in some Continental and Southern European countries, where there is considerable scope to improve the efficiency of resource allocation. For example, we find that AE is around 0.5 in the United States; that is, in the US manufacturing industry the labour productivity index defined in (1) is around 50 per cent higher than it would be if employment shares were randomly allocated across firms. Our results also suggest that AE is higher in the United States than in many Continental European countries. Both findings are similar to those obtained in the smaller but higher quality sample used by Bartelsman et al., (2013) based on national business registers largely from the 1990s. Despite the significant difference in time period between the two studies, more detailed analysis based on a comparison of 2-digit manufacturing industries for the six countries that overlap between the two studies reveals a positive and statistically significant relationship between these two measures of AE.[10]

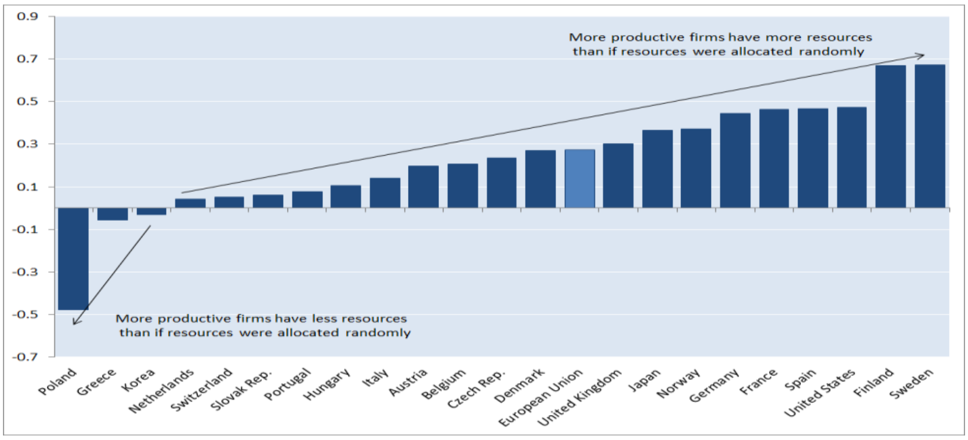



**Figure 2. OECD countries differ in their ability to allocate labour to the most productive firms: covariance across firms between firm size and labour productivity; log points; manufacturing sector in selected OECD countries in 2005**

While it is reassuring that our estimates of allocative efficiency for the manufacturing sector resemble those in leading studies based on high quality data, Table A2 also contains estimates for the services sector and the total business sector. To the best of our knowledge, these are the first estimates of allocative efficiency beyond the manufacturing

[10] For example, the correlation coefficient between our AE measure (using 2005 data) and the Bartelsman et al., (2013) AE measure (using data largely from the 1990s) is 0.35, which is statistically significant at the 1% level. This analysis is based on 48 country-industry pairs for 9 manufacturing sectors (between 20 and 33 under the NACE Rev 1.1 classification) for the following six countries: France, Germany, Hungary, the Netherlands, United Kingdom and the United States.

sector so it is difficult to verify their reliability. Nevertheless, some interesting patterns emerge. In particular, allocative efficiency tends to be much lower in the market services sector than in the manufacturing sector, which in turn yields estimates for the total business sector which are somewhat lower than those presented in Figure 2 (for example, for the EU as a whole, estimated AE is around 0.27, 0.04 and 0.14 for manufacturing, services and total business sectors respectively). The relatively lower efficiency of resource allocation in the services sector might be a symptom of lower competitive pressure, which could reflect technological characteristics or the fact that services are less trade-exposed than manufacturing and that pro-competition product market reforms have generally been more extensive in the manufacturing sector than the services sector (OECD, 2013b). Indeed, we address the latter channel more explicitly in Section 5.2, when we explore the direct impact of anticompetitive service regulation and FDI restrictions on AE in a sub-sample of service sectors.

## 3. PUBLIC POLICY AND RESOURCE ALLOCATION

The preliminary evidence presented above suggests that the efficiency of resource allocation – as measured by the within-industry correlation between a firm's size and its productivity level varies – considerably across OECD countries. What explains these cross-country differences in the efficiency of resource allocation? The working hypothesis in this paper is that regulations affecting product, labour and credit markets – which tend vary significantly across OECD countries – affect resource flows and can thus explain why some countries are more successful at channelling resources to high productivity firms than others. This section discusses the existing evidence for this hypothesis, and the channels through which specific policies influence the efficiency of resource allocation.

### 3.1. Theoretical Literature

Resource misallocation arises when an economy features distortions implying that marginal product (or the marginal value) of inputs is not equated across productive units; in this case, an appropriate reallocation of production factors from low- to high-productivity firms would raise aggregate output. Firm-specific distortions are in fact the main ingredient of a growing theoretical literature attempting to explain productivity differences between countries or industries with the misallocation of resources across firms within each country or industry. While most papers take an agnostic view as to the specific source of such distortions (Restuccia and Rogerson, 2008; Hsieh and Klenow, 2009; Alfaro et al., 2009; and Bartelsman et al, 2013)[11], some works explicitly focus on

[11] The relevance of firm-level deviations from the optimal allocation of resources for aggregate productivity is assessed through calibration exercises on the extent and dispersion of distortions. Restuccia and Rogerson (2008) calibrate to the US economy a model of firm dynamics featuring "scale distortions" (i.e. taxes and subsidies on revenues) and "factor mix distortions" (*i.e.* taxes and subsidies on capital). In turn, they quantify the potential extent of output losses due to misallocation associated with plausible alternative distributions of such distortions. Hsieh and Klenow (2009) fit a similar model to the data to directly estimate the distribution of marginal products in China and India and infer the extent of the underlying distortions by comparing the dispersion of productivity in those countries to that of the US. Similarly, Alfaro et al., (2009) argue that the distribution of firm size in most developing countries is markedly different from the presumed efficient US distribution. Finally, Bartelsman, et al. (2013) calibrate a model featuring policy-induced distortions and firm participation decisions, to match the patterns of AE (as measured in the present paper) in a relatively small sample of economies. They show that a considerable amount of the cross country variation in AE can be explained by differences in the extent and dispersion of distortions across firms, and increasing this dispersion yields a decline in AE.

the role of labour and product market regulation, or financial frictions and credit underdevelopment (Hopenhayn and Rogerson, 1993; Barseghyan and DiCecio, 2011; Moscoso-Boedo and Mukoyama, 2012; Buera et al., 2010).[12]

The potential link between policy induced distortions and allocative efficiency is probably clearest in the context of “size contingent” policies such as labour market regulations that only affect firms above a certain size threshold. When firms are heterogeneous, size-contingent labour regulation creates allocative inefficiencies because too little employment is allocated to relatively high productive firms (both those who remain undersized because of the cost of regulation, and those above the threshold who bear it, see Garicano *et al.*, 2013). More generally, it is a standard equilibrium outcome of dynamic models featuring costs of workforce adjustments (as dismissal costs) that firms will find it optimal not to hire workers even if their short-term marginal product exceeds market wage, and/or to retain workers whose wage exceeds their productivity, therefore lowering allocative efficiency (e.g. Hopenhayn and Rogerson, 1993).

The link between competition in product markets and productivity is the subject of a large literature (see Syverson (2011) for a recent survey) arguing that heightened competition impacts both the “within-firm” and “between-firm” components of aggregate productivity growth. The between effect arises when more efficient producers grow faster than less efficient ones. The within effect typically comes from individual producers (typically, larger incumbents) being induced to become more efficient by taking costly productivity-raising actions that they may otherwise not. Both channels potentially determine a negative link between anticompetitive regulation of product markets (as entry barriers) and allocative efficiency.[13] Bankruptcy regimes (i.e. exit regulation) may also have an impact on allocative efficiency. Bankruptcy codes that excessively punish business failure could affect AE by hampering entrepreneurial start-up activity, thereby implying less competitive pressure on incumbents, and by raising the likelihood that resources are trapped in inefficient firms. It should be noted, however, that anticompetitive product market regulations may also increase efficiency, through the selection of firms in the market: higher barriers in a market imply that entrants (usually smaller than average firms) must have relatively higher productivity levels than in an unconstrained economy.

The list of policies that generate idiosyncratic distortions is not limited to product and labour market regulation. It includes, notably, financial development and the degree of banking competition. Several recent works propose models in which financial frictions, formalized as a collateral constraint imposing restrictions on the ratio between productive capital and private wealth, induce differences in the returns to capital across individual producers (i.e. misallocation of capital, see Greenwood et al., 2010; Buera et al., 2011;

[12] Hopenhayn and Rogerson (1993) quantify the average TFP losses due to the wedges on employment adjustment induced by layoff costs; Barseghyan and DiCecio (2011) demonstrate the negative consequences of entry costs for aggregate outcomes as they reduce the productivity of the marginal entrant through a general equilibrium effect on factor prices. Moscoso-Boedo and Mukoyama (2012) evaluate the effects of both entry regulation and firing costs. Buera et al., (2010) highlight the role of financial frictions and credit underdevelopment in distorting the allocation of capital across heterogeneous production units and also their entry/exit decisions, thereby lowering aggregate and sector-level TFP.

[13] For example, recent works show that entry barriers reduce the market share of more productive firms relative to low productivity firms when a low number of competitors reduces the elasticity of substitution between varieties sold in the market (see Syverson, 2004; Poschke, 2010).

Midrigan and Xu 2013). More generally, non-competitive or poorly regulated banking systems may offer favourable interest rates on loans to select producers adversely (Peek and Rosengren (2005) or based on non-economic factors (Kwaja and Mian, 2005).

At the same time, more developed financial markets are often associated with more efficient selection of firms at entry, which should positively contribute to aggregate productivity through the first term in equation (1), i.e. unweighted average productivity. For example, in general equilibrium models with heterogeneous agents collateral constraints would lower entry of potentially more productive entrepreneurs compared with wealthier but less talented ones (Lloyd-Ellis and Bernhardt 2000; Evans and Jovanovic 1989).[14]

Each of these specific examples is of interest, and ultimately it is important to understand the quantitative significance of specific policies, regulations or institutions.

## 3.2. Empirical evidence

Despite the large theoretical literature, empirical evidence on strength of the link between policies and allocative efficiency is rather limited, and confined to the case of specific countries and policies. As regards to product market regulations, the seminal paper by Olley and Pakes (1996) showed that the deregulation of US telecommunications raised allocative efficiency in that industry. Arnold et al., (2008) found a negative impact of service regulation on allocative efficiency in service intensive downstream industries. Garicano et al., (2013) and Braguinsky et al., (2011) show that size contingent labour regulation induces productive firms to be undersized, with relevant aggregate consequences in the cases of France and Portugal, respectively.

Peek and Rosengren (2005) argue that substantial misallocation of capital is traceable to the existing bank regulation and supervision rules in Japan. Other papers present evidence that low financial development leads to misallocation of credit across producers in developing countries (Banerjee and Munshi, 2004; Banerjee and Duflo, 2005; Kwaja and Mian, 2005; Larrain and Stumpner, 2013). However, recent model-based simulations by Midrigan and Xu (2013) show that the bulk of the productivity losses traceable to financial frictions arise due to worse selection of firms at entry rather than the misallocation of capital across active firms. In term of our decomposition (1), this implies that financial development should be more relevant for unweighted average productivity than for allocative efficiency.

Against this background, we assess the impact of a set of country-wide policy measures on the efficiency of resource allocation in a sample of 21 OECD countries. Throughout the paper, we mainly focus on policy indicators provided by the OECD, which are summarised in Table A3. For example, we use OECD product market regulation indices to measure the extent of "anti-competitive" regulations; that is, regulations "that inhibit

[14] In addition, Aghion et al (2007) develop a simple theoretical model of credit constraints and entry where only observed short-term production capacity can be used as a firm's collateral for borrowing, as opposed to long-term productivity, and arrive at similar conclusions. Of course, it is also possible for selection to work in the opposite direction if financial development increases access to credit but the distribution of quality of those demanding the credit deteriorates at the same time (Cetorelli, 2009).

competition in markets where competition is viable".[15] To measure regulatory distortion affecting financial markets, an index of banking regulation constructed by de Serres et al., (2006) is employed, which measures entry barriers, line of business restrictions and the impact of state control on the level playing field. Crucially, however, this index abstracts from norms aimed at achieving financial stability objectives, which are likely to have more limited implications for competition and are clearly desirable from a macroeconomic perspective. Following standard practice in the literature, financial development is proxied by the ratio of private credit to GDP, while the OECD Employment Protection Legislation Index is employed to proxy the extent of labour market rigidities (see Bassanini et al., 2009). Finally, an indicator measuring as the cost to close a business – sourced from the World Bank – is included to explore the impact of bankruptcy legislation on allocative efficiency.[16]

For illustrative purposes, in Figure 3 we relate the index of AE to various framework policy indicators, observed at the country level.[17] Countries with high barriers to entry and more stringent bankruptcy arrangements are characterised by a lower covariance between firm size and their productivity (i.e. lower AE). The same is true with respect to the stringency of Banking Regulation, while Employment Protection Legislation appears unrelated to AE.

[15] As outlined in Nicoletti et al. (2000), restrictions to competition were defined either as barriers to access in markets that are inherently competitive or as government interferences with market mechanisms in areas in which there are no obvious reasons why such mechanisms should not be operating freely (e.g. price controls imposed in competitive industries as road freight or retail distribution). Moreover, the construction of the indicators assumes that regulatory patterns do not reflect cross-country differences in the level of public concern for the market failures that motivate regulations, but rather reflect regulatory failure or policies adverse to competition.

[16] Unless otherwise noted, we used policy data referred to 2003; allocative efficiency is measured in 2005 as this is the year in which the underlying firm level data guarantee the largest coverage of countries and industries, and the lowest number of missing values for the relevant variables.

[17] Country level AE is obtained applying decomposition (1) to the entire sample of firms in a country, and might therefore be affected by cross-country differences in the underlying industry composition. However, constructing average indexes for each country aggregating industry-level indicators using a common set of (US) industry weights (as in Bartelsmann et al, 2013) does not alter the picture.

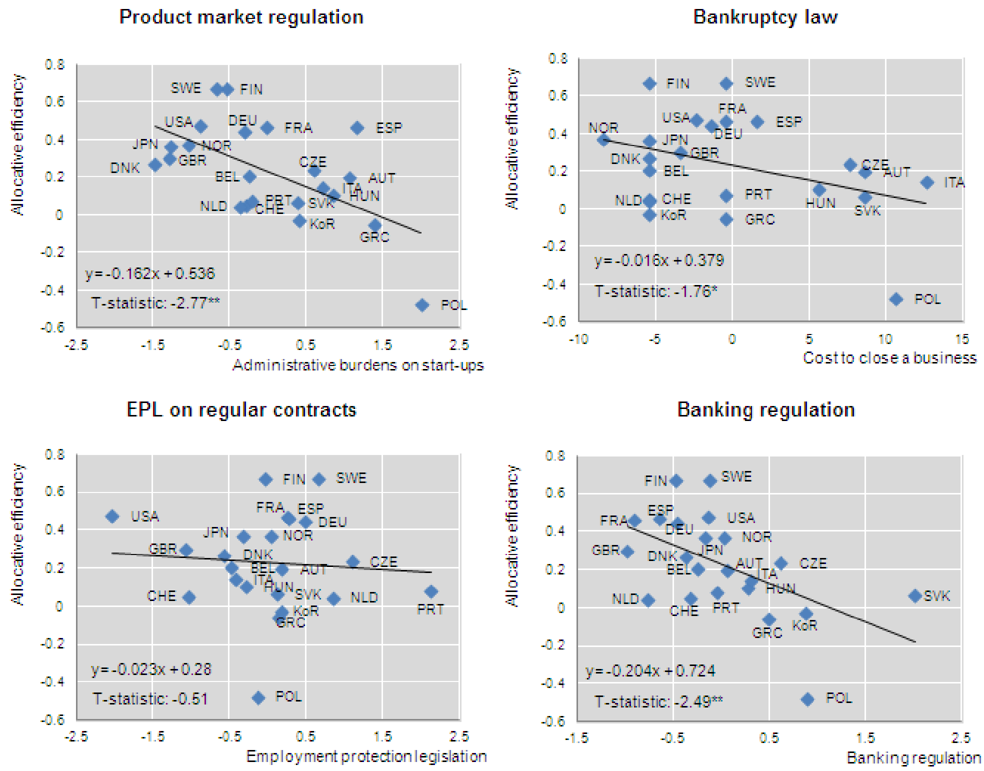



**Figure 3. Allocative efficiency in the manufacturing sector and framework policies (mid 2000s)**

Cross country correlations are clearly only suggestive of the possible link between policy induced distortions and the patterns of AE, due to the likely biases induced by observed and unobserved country-level confounding factors and reverse causality. The next section describes how these issues are tackled in the empirical analysis.

## 4. EMPIRICAL FRAMEWORK

To account for the potential sources of spurious correlation between country-level policies and economic outcomes, all our estimates are conditioned on country specific fixed-effects. Whenever plausible instruments are available, we also report 2SLS estimates. To gain within country variability in the policy variables of interest, we will exploit cross-industry heterogeneity in exposure to a given policy (e.g. the relevance of technology-induced entry costs in the case of entry barriers), and achieve identification from comparing the differential AE between highly and marginally exposed industries in countries with different levels of regulation. In a complementary exercise, we more directly examine the link between AE and regulation exploiting the availability of industry specific measures of regulation for a sub-sample of service industries. In this case, identification will be driven by within-country variation in service-specific policies and AE.

### 4.1. Identification of the impact of country-level policies

In the absence of industry-specific policy indicators covering all industries in our sample, we exploit cross-industry cross-country data and a differences-in-differences specification accounting for country-level time invariant unobserved characteristics. This approach, popularised by Rajan and Zingales (1998) is based on the assumption that there exist industries that have 'naturally' high exposure to a given policy (i.e. the treatment group), and such industries – to the extent that the policy is relevant to the outcome of interest – should be disproportionally more affected than other industries (i.e. the control group). To see this more clearly, consider the case of entry regulation. In this case, the baseline assumption is that there exist industries that have 'naturally' high entry barriers (possibly due to capital intensiveness of production or technological complexity), and industries where these barriers are almost negligible, and this pattern does not vary across countries. In this case, the marginal impact of an increase in the administrative cost of entry (the treatment) on AE could be expected to be smaller in industries where 'natural' entry barriers are very high, than in industries where entry barriers are low. Under the additional identifying assumption that AE is only affected by entry regulations via the reallocation channel, this constitutes evidence that entry regulations affect resource allocation.

If the presence of technological characteristics affects industry exposure to regulation, we would expect to see this effect empirically in the interaction of industry exposure and regulation. Our regressions will thus take the form:

$$AE_{jc} = \beta\left(Exp_j * Reg_c\right) + \mu_j + \mu_c + \varepsilon_{jc} \qquad (2)$$

where $AE_{j,c}$ measures allocative efficiency at the country-industry level, $Reg_c$ measures the stringency of product, labour or credit market policies in country $c$ and $Exp_j$ is an industry-level index aiming at capturing differences in the relevance of regulation for firms operating in different sectors. Interacting country-level policy variables with industry variables makes it possible to condition our estimates on country and industry fixed-effects, respectively $\mu_c$ and $\mu_j$. In this context, estimating $\beta < 0$ would indicate that easing regulation boosts AE disproportionately more in highly exposed industries than in less exposed industries.

Industry-level indexes of exposure are taken from the large literature exploiting the same framework to infer the relevance of country-level policies on a number of economic outcomes. Empirical studies on the relevance of Product Market Regulation (PMR) use firm turnover in a benchmark country (e.g. the US) as index of industry exposure to entry barriers, since industries with high natural entry barriers will likely exhibit relatively low turnover rates (firm turnover is also used to identify the impact of bankruptcy codes since such legislation will be more relevant to more dynamic sectors. See e.g. Fisman and Sarria Allende, 2010; Klapper et al., 2006). Similarly, studies examining the role of Employment protection legislation (EPL) use measures of worker reallocation (job turnover or layoff rates, see Micco and Pagés, 2006; Bassanini et al., 2009) to identify industries that are more likely to be affected by relatively stringent employment protection. Finally, to explore the impact of banking regulation and financial underdevelopment on AE, we use external finance dependency as an index of sectoral exposure (see Rajan and Zingales, 1998).

Following the above mentioned literature, all industry-level indexes of exposure used in the analysis are computed from US data. This common practice is generally motivated on two grounds. First, to the extent that United States is generally perceived to be a low regulation (i.e. “frictionless”) country, using US data mitigates concerns regarding the possible endogeneity of exposure to the level of regulation. Second, most of the individual-level data required to compute the indexes are not available in the case of other countries. Accordingly, the United States is excluded from the analysis.

Table A3 in the Supplementary Appendix reports details on the country level policy variables of interest and the corresponding industry-level exposure variables used in the difference in differences estimator; descriptive statistics of these variables are contained in Table A4. We also experimented with a number of other policy variables but these produced inconclusive results (see Section 6.3).

## 4.2. Identification of the impact of sectoral-level policies

To test the robustness of some of our policy conclusions based on the differences-in-difference estimates, we utilise industry-level variation in Product Market Regulation and Foreign Direct Investment (FDI) restrictions in services. One advantage of this estimation approach is that it allows us to more directly infer the average effect of policies on AE, but as explained in Section 4.3, the resulting estimates may be more prone to bias arising from endogeneity and reverse causality than the differences-in-differences estimates. We estimate the following cross-sectional regression:

$$AE_{s,c} = \theta Reg_{s,c} + \eta_s + \eta_c + \epsilon_{sc} \qquad (3)$$

where $AE_{s,c}$ represents allocative efficiency in each service industry $s$ and country $c$ according to decomposition (1); $\eta_s$ and $\eta_c$ capture industry- and country fixed-effects, respectively, and $Reg_{s,c}$ measures the level of anti-competitive restrictions in Energy, Retail trade, Air and Land Transportation, Post and telecommunication, and Professional services. Unless otherwise noted, $Reg_{s,c}$ is a simple average of the OECD sectoral regulation sub-indexes measuring barriers to entry, the regulation of market conduct (such as restrictions on the legal form of businesses, bans to advertising etc.) and price control, and excludes public ownership. Hence, our estimates of the impact of service regulation on AE in services ($\theta$) exploit cross-industry variation within countries (accounting for common industry-specific factors).

For the same industries, the OECD also collects specific measures of statutory restrictions on foreign direct investment (FDI Regulatory Restrictiveness Index). For details, see Table A3. Accordingly, we extend the above analysis to test the influence of barriers to foreign competition through FDI in domestic service markets on AE.

### 4.3. Identification concerns

While the empirical approaches illustrated above absorb country-specific time invariant observable and unobservable characteristics that might induce spurious correlation, endogeneity remains a concern. These concerns may be stronger in the case of equation (3), since service specific regulation might be a consequence, not a cause of the efficiency of resource allocation (e.g. reverse causality). This would be the case, for example if, within a country, service industries with low efficiency (that is, sectors where large firms are inefficient, on average) were characterised by high policy-induced entry barriers for political economy reasons (i.e. inefficient firms lobby for industry protection measures). Unfortunately, instruments for industry-level regulation across countries are not available.

Endogeneity and reverse causality are less of a concern in the context of specification (2), due to the differences-in-differences specification and the fact that the policy variables of interest do not vary at the country-industry level.[18] One could still argue, however, that inefficient incumbent firms in sectors with higher natural turnover could be more vocal in lobbying for protection in the form of higher policy-induced entry costs than firms in sectors with lower natural turnover. As in the previous case, this would imply that our estimates would overstate the extent to which the negative impact of barriers to entry on AE in high turnover sectors exceeds the effect on low turnover ones. To test the robustness of our core estimates to the endogeneity critique, we re-estimate equation (2) using an instrumental variables approach. Based on the existing literature, we were able to identify a few potential – albeit crude – instruments for product and labour market regulation, largely based on the characteristics of the legal system.

Legal origin is shown to have a high explanatory power of governments' proclivity to intervene in the economy (see La Porta et al., 1999). The broad idea is that while the common law tradition started in the 17th century reflects the English Parliament and aristocracy's intent to limit the power of the sovereign (and therefore put emphasis on restraining the government and on protecting the individual against the government), a civil legal tradition reflects the intent to build institutions to further the power of the State. Following Barseghyan (2008), we use a classification of countries based on the origin of their commercial laws to instrument barriers to entry.[19] To generate exogenous variation in EPL, we follow Bassanini et al., (2009) who used both an indicator (obtained from the above mentioned variable) of whether a country has a common law as opposed to a civil law system, and a refinement of this dichotomy based on information on civil codes. The idea is again that while countries with common law systems tend to have relatively few regulatory provisions concerning labour contracts, most civil law systems, and especially those with a single codified civil code, tend to regulate more.

[18] While the differences-in-differences approach popularized by Rajan and Zingales (1998) has been applied to a variety of settings, in particular when the role of country-level policy variables is under scrutiny, it has potential estimation problems, as discussed in Ciccone and Papaioannou (2010).

[19] In practice, the instrument consists in five categorical indicators of whether a country's legal system is based on British common law, on French, German, or Scandinavian civil law, or inherited from Soviet laws. As in Barseghyan (2008), we also experiment with another commonly used instrument in cross-country studies of the consequences of differences in institutional quality: geographic latitude (see Hall and Jones, 1999).

In both cases, the underlying assumption is that any economic mechanism inducing an effect of legal systems on AE that varies across industries as a function of exposure to the policy (firm turnover or layoff propensity) would operate through their effect on the regulation of product or labour markets.

In addition to the instrumental variable estimates outlined above, Section 6 contains a series of robustness checks to assess the sensitivity of our baseline estimates to the specification and sample used, to the inclusion of a number of other country-level structural variables (e.g. institutional quality, corruption, openness etc.) and to measurement issues, including computing AE based on alternative firm-level productivity indexes. Moreover, we deal with whether corrections are required when estimating the standard errors. In what follows our inference will be based on heteroskedasticity-robust standard errors, therefore implicitly assuming that shocks in (2) and (3) are not correlated (or equi-correlated) within industries and/or regions (or both). If this was not the case, then the estimated standard errors would be biased (although not necessarily downwards). However, clustering procedures might also introduce biases in the estimated standard errors in particular when the number of available units is small as in this application (see Thompson, 2011). We will return on this issue in section 6, showing that clustering the standard errors does not alter the main conclusions of the analysis.

## 5. EMPIRICAL RESULTS

This section discusses the results of the econometric analysis outlined in Section 4. We begin by discussing the core differences-in-differences results based on equation (2), which aims to assess the impact of the various public policy variables listed in Table A3 on AE. Section 5.1 also explores the extent to which the impact of policies on AE varies with the innovative capacity of the sector, in order to provide some further evidence on the channels through which policy distortions affect AE, and across common industry groupings. This analysis is based on 834 country-industry cells for the non-farm business sector. In turn, section 5.2 explores the impact of sector-specific regulations on AE in a smaller sub-set of services sectors (equation 3). In all these specifications AE will be based on labour productivity as measured by average turnover per worker. In Section 6.2 we will discuss results obtained using alternative productivity measures, such as value added per worker and TFP.[20]

### 5.1. The impact of policies on allocative efficiency and productivity

Table 1 reports the baseline estimates of equation 2 (panel A) as well as the coefficient obtained when the dependent variable is replaced by unweighted average productivity (the first term in decomposition (1); see panel B). The main results are summarized as follows.

Barriers in the product market (entry and exit) are negatively related to AE (columns 1 to 3 of Table 1; panel A). Results in columns 1 and 2 focus on economy-wide barriers to entry. The estimated negative coefficients indicate that lowering entry regulation would

[20] As discussed in the Data section, using turnover allows maximising the number of available firms and is less likely to be subject to measurement error. Use of firm level labor productivity is rather common and is theoretically motivated in studies on resource allocation in the recent work of Bartelsman et al (2013).

increase AE disproportionately more in highly exposed (i.e. high firm turnover) industries. To evaluate the relevance of the estimated effects, consider the difference in AE between a high turnover industry (such as Retail) and a low turnover industry (such as Rubber and plastic products). If we take the estimates from Table 1 as causal (causality will be discussed in more detail in Section 6), then according to the estimates from column 1, reducing Administrative burdens on start-ups from the high level of Italy to the lower level of Finland implies a gain in the above differential of more than 5 percentage points (recall that AE is measured in log points). Similar results are obtained focusing on exit policies (i.e. the effectiveness of bankruptcy regulation). Our estimate in column 3 implies that reducing the cost to close a business from the high levels of Hungary or Spain to those of Denmark or the Netherlands would improve the differential AE in high relative to low turnover industries by 4.8 percentage points.

Employment protection legislation also makes the reallocation of resources across heterogeneous firms less efficiency enhancing (columns 4 and 5). To appreciate the relevance of the estimated effect, consider the difference in AE between a high layoff industry (such as machinery and equipment) and a low layoff industry (such as Chemicals). Our estimates imply that reducing EPL from the high levels of Spain to the lower level of Japan implies a gain in the above differential in excess of 4.5 percentage points. To the extent that such policy-induced distortions are also likely to hinder the reallocation of resources across sectors, the above estimates of the impact of policies on AE may represent a lower bound, and the overall impact on resource allocation could be somewhat larger.[21]

The interaction of banking regulation (respectively, financial development) with external finance dependence is negatively (respectively, positively) correlated with AE. However, neither coefficient is statistically significant. This finding might be surprising in light of the well-established evidence of the positive link between financial development and economic growth and productivity. Notice however that the countries in our sample generally have a high level of financial development. This is important as several works adopting the Rajan–Zingales approach find that the empirical relevance of the finance-growth nexus loses statistical significance as developing countries are omitted (Carlin and Mayer, 2003; Manning, 2003).[22] Moreover, it might be that the bulk of the productivity gains traceable to finance arise due to the better selection of firms at entry rather than the improved allocation of resources, as highlighted by Midrigan and Xu (2013).

To test this latter hypothesis, and more generally to assess the extent to which the negative association of regulation with AE has a bearing for aggregate productivity, we run the same specifications replacing AE by unweighted industry productivity (the first term in decomposition (1)) as the dependent variable. The results, reported in Panel B of Table 1, show no significant effects of product market regulations, Bankruptcy legislation

[21] Table A5 reports standardized regression coefficients for the same specifications, which should be interpreted as the effects of increasing the "impact of regulation" (i.e. the product $Exp_j*Reg_c$). They suggest that employment protection is slightly more relevant than product market regulation.

[22] Consistently with this hypothesis, in a contemporaneous work Larrain and Stumpfner (2013) find a significant impact of financial liberalization on allocative efficiency (measured with the Olley-Pakes covariance term) in a sample of 10 Eastern European transition countries.

or EPL on the unweighted productivity term, suggesting that the estimated allocative gains from easing the regulatory burden are indeed reflected in higher productivity.

On the other hand, less stringent banking regulation (and greater financial development) is associated with higher unweighted average productivity (see columns 6-7, Table 1, Panel B).[23] Thus, even if they do not imply significant differences in employment allocation across units, well-functioning financing mechanisms would seem to raise productivity by lowering the proportion of low productivity firms relative to high productivity firms in the market.

The results for the most general specification where all of the policy interaction terms are included together are included in columns 8 and 9, and are broadly similar to the univariate regressions results.[24]

Within the same estimation framework, we explore whether the impact of policies on AE varies across: (i) common industry groupings (industry, energy and construction vs. services); and (ii) the distribution of patenting intensity, a potential indicator of the innovative capacity of the industry. Both exercises are performed by interacting each $Reg_c$*$Exp_s$ variable in equation (2) with dummies indicating whether the industry pertains to each of the relevant groups. This approach allows us to test whether the group-specific coefficients are (statistically) significantly different from one another.

We find very little evidence that the average effect of policies estimated in Table 1 differs in services vs. other industries in the economy (see Table 2, Panel A): the group-specific coefficients are very similar in magnitude and the differences are rarely statistically significant.

On the other hand, product and labour market regulations and bankruptcy codes are more relevant for AE in more innovative industries (see Table 2, Panel B). This is illustrated by the differences in coefficient test, which shows that the impact of the policy for the top quartile of the patenting intensity distribution is generally larger than the impact of policy for the bottom quartile of the patenting distribution at the 5% level of statistical significance. There are a number of reasons why innovative firms might be more sensitive to rigidities in the reallocation process (Andrews and Criscuolo, 2013). First, innovative firms will place a high option value on flexibility given their tendency to experiment with uncertain technologies. Second, innovative firms need to rapidly reallocate tangible resources in order to capture the value of their knowledge-based investments before other firms imitate, and rapidly scale down operations to facilitate exit and thereby release resources that can be used by other firms.

[23] Indeed, in a regression of weighted average productivity on financial development interacted with external finance dependency (henceforth FD*EFD) along with country and industry fixed effects, the coefficient on FD*EFD is about 0.06 and is statistically significant at around the 5% level. Consistent with the decomposition in Equation (1), this estimated coefficient is roughly equivalent to the sum of the coefficient on FD*EFD in the AE regression (0.15; see Table 1, Panel A and the FD*EFD coefficient in the unweighted productivity regression (0.045 see Table 1, Panel B). Performing the same exercise with BankingRegulation*EFD yields similar conclusions.

[24] In Panel A, the coefficient on the bankruptcy interaction becomes insignificant largely reflecting the high correlation between the entry and exit regulations (correlation = 0.62). Indeed, the bankruptcy interaction remains significant in a regression that controls for EPL and financial development interaction terms but not entry regulation.

**Table 1: Public policies and allocative efficiency across OECD countries and industries**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| Variables: | PMR and Bankruptcy | | | EPL | | Banking & finance | | All-in | |
| | PANEL A: ALLOCATIVE EFFICIENCY | | | | | | | | |
| BTE X turnover | -0.008** | | | | | | | -0.007* | -0.007** |
| | (0.003) | | | | | | | (0.004) | (0.003) |
| BTE2 X turnover | | -0.015** | | | | | | | |
| | | (0.006) | | | | | | | |
| Bankruptcy X turnover | | | -0.001* | | | | | -0.000 | -0.000 |
| | | | (0.000) | | | | | (0.000) | (0.000) |
| EPLR X layoff | | | | -0.052*** | | | | -0.051*** | -0.052*** |
| | | | | (0.015) | | | | (0.015) | (0.015) |
| EPLO X turnover | | | | | -0.017*** | | | | |
| | | | | | (0.005) | | | | |
| FinDev X ExtFinDep | | | | | | 0.015 | | 0.014 | |
| | | | | | | (0.020) | | (0.019) | |
| BankReg X ExtFinDep | | | | | | | -0.009 | | -0.010 |
| | | | | | | | (0.015) | | (0.014) |
| AdjR2 | 0.556 | 0.557 | 0.553 | 0.567 | 0.565 | 0.553 | 0.553 | 0.572 | 0.572 |
| | PANEL B: UNWEIGHTED PRODUCTIVITY | | | | | | | | |
| BTE X turnover | -0.000 | | | | | | | 0.002 | 0.002 |
| | (0.003) | | | | | | | (0.005) | (0.005) |
| BTE2 X turnover | | 0.006 | | | | | | | |
| | | (0.006) | | | | | | | |
| Bankruptcy X turnover | | | -0.000 | | | | | -0.000 | -0.000 |
| | | | (0.000) | | | | | (0.001) | (0.001) |
| EPLR X layoff | | | | 0.002 | | | | 0.002 | 0.000 |
| | | | | (0.027) | | | | (0.028) | (0.027) |
| EPLO X turnover | | | | | 0.003 | | | | |
| | | | | | (0.005) | | | | |
| FinDev X ExtFinDep | | | | | | 0.045** | | 0.045** | |
| | | | | | | (0.019) | | (0.019) | |
| BankReg X ExtFinDep | | | | | | | -0.025* | | -0.025* |
| | | | | | | | (0.014) | | (0.013) |
| AdjR2 | 0.876 | 0.876 | 0.876 | 0.876 | 0.876 | 0.879 | 0.878 | 0.879 | 0.878 |
| Observations | 834 | 834 | 834 | 834 | 834 | 828 | 828 | 828 | 828 |

Notes: In panel A, the dependent variable is allocative efficiency as defined (1), computed in 2005. In Panel B, the dependent variable is unweighted productivity as defined (1), computed in 2005.See Table A3 for definitions and sources of the explanatory interaction variables. All specifications include country and industry fixed effects. Observations are weighted by the number of firms used to compute the productivity decomposition in (1). Robust standard errors in parentheses; *** denotes statistical significance at the 1% level, ** significance at the 5% level, * significance at the 10% level.

**Table 2: Public policies and allocative efficiency across OECD countries and industries: heterogenous effects**

| VARIABLES | (1) BTE X turnover | (2) BTE2 X turnover | (3) Bankruptcy X turnover | (4) EPLR X layoff | (5) EPLO X turnover | (6) FinDev X ExtFinDep | (7) BankReg X ExtFinDep |
|---|---|---|---|---|---|---|---|
| PANEL A: EFFECTS OF POLICIES ON INDUSTRY AND SERVICES | | | | | | | |
| Effects on Industry | -0.010*** | -0.021*** | -0.001** | -0.060*** | -0.017*** | -0.032 | -0.003 |
| | (0.003) | (0.006) | (0.000) | (0.021) | (0.005) | (0.023) | (0.019) |
| Effects on Services | -0.008*** | -0.017*** | -0.001** | -0.070** | -0.017*** | 0.022 | -0.010 |
| | (0.003) | (0.005) | (0.000) | (0.034) | (0.005) | (0.022) | (0.017) |
| Test Industry-Services = 0 (p-value) | 0.193 | 0.0731 | 0.373 | 0.537 | 0.708 | 0.0453 | 0.777 |
| AdjR2 | 0.559 | 0.563 | 0.554 | 0.568 | 0.565 | 0.555 | 0.552 |
| PANEL B: EFFECTS OF POLICIES BY INDUSTRY PATENTING INTENSITY | | | | | | | |
| Effects on High Patenting (Q4) | -0.013*** | -0.022*** | -0.002*** | -0.062*** | -0.023*** | -0.000 | -0.014 |
| | (0.004) | (0.007) | (0.001) | (0.016) | (0.006) | (0.019) | (0.019) |
| Effects on Med-to-High Patenting (Q3) | -0.011*** | -0.019*** | -0.001** | -0.060*** | -0.022*** | -0.002 | -0.009 |
| | (0.004) | (0.007) | (0.001) | (0.016) | (0.006) | (0.029) | (0.017) |
| Effects on Med-to-Low Patenting (Q2) | -0.011*** | -0.019*** | -0.001** | -0.057*** | -0.020*** | 0.013 | -0.008 |
| | (0.003) | (0.006) | (0.000) | (0.016) | (0.005) | (0.028) | (0.015) |
| Effects on Low Patenting (Q1) | -0.009*** | -0.017*** | -0.001** | -0.051*** | -0.019*** | 0.033 | 0.010 |
| | (0.003) | (0.006) | (0.000) | (0.016) | (0.005) | (0.028) | (0.019) |
| Test High -Low = 0 (p-value) | 0.0169 | 0.0230 | 0.0359 | 0.151 | 0.0270 | 0.285 | 0.184 |
| AdjR2 | 0.560 | 0.560 | 0.554 | 0.570 | 0.572 | 0.553 | 0.558 |
| Observations | 834 | 834 | 834 | 834 | 834 | 828 | 828 |

Notes: Q4, Q3, Q2, and Q1 refer to the four quartiles of the patenting intensity distribution. See Table A3 for definitions and sources of the explanatory interaction variables. All specifications include country and industry fixed effects. Observations are weighted by the number of firms used to compute the productivity decomposition in (1). Robust standard errors in parentheses; *** denotes statistical significance at the 1% level, ** significance at the 5% level, * significance at the 10% level.

## 5.2. Assessing the aggregate impact of policy reforms

Guiso et al. (2004) proposed a methodology to simulate the aggregate impact of policy reforms using estimates from country-industry interaction models as (2). Their procedure can be applied to evaluate the potential gains in allocative efficiency from hypothetical product and labour market reforms, using the coefficients estimated in Table 2 (see Box 2 for details). In both cases, we will focus on an EU-wide policy lowering regulation in any of the member countries ($Reg_c$) to a target, or "best practice" level ($Reg_{BP} < Reg_c$).

**Box 2: a methodology to simulate the impact of reforms**

For each of the two policy reforms, applying the procedure proposed by Guiso et al. (2004) involves two steps.

First, we need to estimate the impact of the reform on allocative efficiency in country *c* and industry *j* ($\Delta\widehat{AE}_{jc}$). This requires multiplying the estimated coefficient $\hat{\beta}$ by the difference between best practice and country regulation ($Reg_{BP}$ - $Reg_c$), taking into account industry exposure to regulation:

$$\Delta\widehat{AE}_{jc} = \hat{\beta} Exp_j \left(Reg_{BP} - Reg_c\right).$$

For any given sector *j*, the countries whose efficiency benefits most are those with the largest gap in regulation ($Reg_{BP}$ - $Reg_c$). For any given regulation gap, the sectors that gain most are those with the highest exposure to regulation ($Exp_j$). The impact on a country's AE will therefore depend both on the extent of the reform and on its industrial specialization.

Second, we summarize the benefits of the deregulation policy at the country or industry level computing weighted averages of the expression above. More precisely, the impact of the policy reform on allocative efficiency in each country *c* is obtained as:

$$\Delta\widehat{AE}_c = \sum_j \left[\frac{L_{jc}}{\sum_j L_{jc}} \hat{\beta} Exp_j (Reg_{BP} - Reg_c)\right]$$

where $L_{jc}$ is employment in country *c*, industry *j*. Similarly, the impact of the policy reform on allocative efficiency in each industry *j* can be computed as:

$$\Delta\widehat{AE}_j = \sum_c \left[\frac{L_{jc}}{\sum_c L_{jc}} \hat{\beta} Exp_j (Reg_{BP} - Reg_c)\right]$$

Finally, country-specific gains can be further aggregated to obtain the impact of the reform on efficiency in the EU as a whole:

$$\Delta\widehat{AE}_{EU} = \sum_c \left(\frac{L_c}{L_{EU}} \Delta\widehat{AE}_c\right)$$

where $L_{EU}$ is overall employment in the European Union ($L_{EU} = \sum_j \sum_c L_{jc}$ ).

Two important caveats are in order before we proceed. First, the validity of this simulation procedure is conditional on a set of very restrictive assumptions, which are discussed in detail by Bassanini et al. (2009). Essentially, we need to assume that regulation has no, or negligible, aggregate (i.e. country-level) impact on allocative efficiency. If relevant, such effect should of course be taken into account in the simulation of the aggregate impact of the reform. However, its magnitude cannot be estimated in the context of model (2) due to the presence of country-specific fixed-effects. We also need to assume that the simulated policy would have no relevant consequences on the structure of production, that is, on industry employment shares within a country.[25] The second important caveat when interpreting these exercises is that, being purely cross-sectional, our analysis will be silent as to the time horizon in which the computed gains could materialize.

With these caveats in mind we applied Guiso et al. (2004) procedure to the cases of two policies reducing (i) administrative burdens on start-ups (the variable used in col. 1 of Table 1) and (ii) restrictions on individual dismissals (used in col. 3) in each country to the to the lowest level observed within the EU. In 2008, the latest year for which both regulation indexes are available, such "best practices" were observed, respectively, in Denmark and in the UK.

The results, reported in Table 3, suggest that the productivity gains from both reforms could be sizeable. Lowering product market regulation would raise allocative efficiency in the EU by around 15 log points, doubling the value that we estimated for the area using ORBIS data (14 log points, see Table A2; for comparison, overall allocative efficiency in the US is around 40 log points). In other words, in a reformed EU the index of labour productivity defined in (1) would be about 15 per cent higher owing to the improved efficiency of resource (i.e. employment) allocation across firms.[26]

This overall impact averages over heterogeneous country and industry effects. Looking at the country level gains, we found that the increase would be in excess of 30 log points in Poland and Greece (the two countries with lowest level of allocative efficiency, see Table A2) and above the EU average in troubled euro area countries such as Spain, Portugal and Italy. It would be lower than average in core countries as the UK, Germany and France. Looking across industries, lowering product market regulation would on average be more beneficial in service industries as retail trade, transport and communication, as well as the construction sector, than in manufactures. Interestingly, however, the highest gains would accrue to the traditional manufacturing industries (food, textiles and footwear) whose share is relatively large in highly regulated countries.

The impact of a reform lowering employment protection would be even higher. Averaging over all EU countries and industries, the implied increase of allocative efficiency is estimated around 23 log points. At the country level, the gains would be above the EU average not just in peripheral countries as Portugal, but also in core, highly regulated countries as Germany, Sweden and the Netherlands. Aggregating the effect across industries, the EPL reforms would be more beneficial in manufactures than in services, with allocative efficiency in Machinery and equipment estimated to increase in excess of 30 log points.

[25] Bassanini et al. (2009) show evidence supporting the second assumption in the case of EPL reforms in OECD countries.

[26] This is because, according to our results in Panel B of Table 1, such reform would have no significant impact on unweighted average productivity. Hence, changes in allocative efficiency translate into changes in industry (weighted) productivity.

**Table 3: Product and labour market reforms in the EU: potential gains in allocative efficiency by country and industry** (log points)

| Country | Simulated reform: Product market | Simulated reform: Labour market | Industry | Simulated reform: Product market | Simulated reform: Labour market |
|---|---|---|---|---|---|
| *European Union* | *0.15* | *0.23* | | | |
| Austria | 0.27 | 0.24 | **Manufactures:** | | |
| Belgium | 0.17 | 0.12 | Food, Textiles and Footwear | 0.20 | 0.26 |
| Czech Rep. | 0.25 | 0.38 | Wood, Paper and Printing | 0.15 | 0.24 |
| Germany | 0.01 | 0.36 | Chemical, Plastics and Fuel Prod. | 0.12 | 0.20 |
| Denmark | BP | 0.10 | Basic Metals and Fabr. Metal Prod. | 0.12 | 0.25 |
| Spain | 0.30 | 0.26 | Other Non-Metallic Mineral Products | 0.14 | 0.20 |
| Finland | 0.16 | 0.21 | Machinery and Equipment | 0.10 | 0.33 |
| France | 0.14 | 0.26 | Transport Equipment | 0.12 | 0.20 |
| United Kingdom | 0.03 | BP | **Services:** | | |
| Greece | 0.36 | 0.23 | Electricity, Gas, Water | 0.07 | 0.12 |
| Hungary | 0.37 | 0.16 | Retail Trade, Restaurants, Hotels | 0.16 | 0.19 |
| Italy | 0.19 | 0.13 | Transport, Storage, Communication | 0.18 | 0.20 |
| Netherlands | 0.13 | 0.31 | Real Estate, Business Activity | 0.13 | 0.22 |
| Poland | 0.43 | 0.18 | | | |
| Portugal | 0.21 | 0.63 | Construction | 0.19 | 0.31 |
| Slovak Rep. | 0.23 | 0.28 | | | |
| Sweden | 0.07 | 0.34 | | | |

Notes: The Table shows the potential increase in allocative efficiency at the country and industry level stemming from an EU-wide policy lowering administrative burdens on start-ups (PMR) and restrictions on individual dismissals (EPL) in each country in the European Union to the corresponding lowest level observed within the EU. The simulation uses the coefficients estimated in Columns 1 and 3 of Table 1, and the OECD regulation indexes as measured in 2008. "BP" identifies the best practice country.

### 5.3. Supporting evidence from a subset of service sectors

Table 4 shows the estimation results of equation (3) in a subset of service sectors. Lower anti-competitive regulations in services are associated with a significantly higher correspondence between firms' productivity and their employment share. For example, our baseline estimate in column 1 indicates that AE would increase by 11 percentage points if, holding constant other time-invariant country characteristics, $Reg_{s,c}$ was reduced by one point (corresponding to slightly less than the average within-country standard deviation of service regulation, or the difference between average service regulation in France and in the United Kingdom).

**Table 4: Public policies and allocative efficiency in the services sector**

| | (1) | (2) | (3) |
|---|---|---|---|
| | PANEL A: ALLOCATIVE EFFICIENCY | | |
| VARIABLES | Base | Base & FDI - 1 | Base & FDI - 2 |
| Service sector regulation | -0.111** | -0.130*** | |
| | (0.053) | (0.048) | |
| FDI restrictions | | -0.917** | -0.962* |
| | | (0.462) | (0.529) |
| Service sector regulation (including public ownership) | | | -0.118** |
| | | | (0.060) |
| AdjR2 | 0.624 | 0.629 | 0.593 |
| | PANEL B: UNWEIGHTED PRODUCTIVITY | | |
| VARIABLES | Base | Base & FDI - 1 | Base & FDI - 2 |
| Service sector regulation | -0.029 | -0.009 | |
| | (0.028) | (0.024) | |
| FDI restrictions | | -0.118 | -0.091 |
| | | (0.184) | (0.194) |
| Service sector regulation (including public ownership) | | | 0.006 |
| | | | (0.030) |
| AdjR2 | 0.954 | 0.964 | 0.964 |
| Observations | 174 | 152 | 152 |

Notes: In panel A, the dependent variable is allocative efficiency as defined in (1), computed in 2005. In Panel B the dependent variable is unweighted productivity as defined in (1), computed in the same years. See Table A3 for definition and sources of the explanatory variables. All specifications include country and industry fixed effects. Observations are weighted by the number of firms used to compute the productivity decomposition in (1). Robust standard errors in parentheses; *** denotes statistical significance at the 1% level, ** significance at the 5% level, * significance at the 10% level

Higher FDI restrictions are also associated with lower AE in services. The coefficient in column 2 implies that a reduction in FDI restrictions from the high level in Poland to the lower level in Germany would be associated with an increase in AE of nearly 5 percentage points. Both results are unaffected if we broaden the definition of $Reg_{s,c}$ to include the extent of public ownership (column 3).

As in the previous exercise we find that the estimated gains in AE largely translate into higher industry productivity. In a regression of unweighted productivity, the coefficient estimates are generally not statistically significant (see Table 4; Panel B).

If interpreted causally, these findings would imply a sizeable direct effect of lowering service regulation on aggregate service efficiency. One way to see this is to apply the simulation procedure introduced in the previous section to infer the aggregate gains of a

EU-wide service deregulation policy that reduces regulation in each service ($Reg_{c,s}$) to the lowest level observed within the EU (i.e. the "best practice", $Reg_{BP,s}$).[27]

Our estimates imply that, for the EU as a whole, allocative efficiency in the subset of regulated service industries would increase by around 17 log points.[28] Considering that regulated industries account for more than three quarters of overall service employment, this estimate translates into an increase of allocative efficiency in EU services of around 13 points. In other words, service labour productivity in the EU would be about 13 percentage points higher as a consequence of reforming the sub-set of regulated industries. This is a relevant finding as allocative efficiency in EU service industries is especially low (less than 4 log points in 2003, compared to 36 points in the US. See Table A2).

When looking at the gains in each regulated service separately, these would be higher than average in retail trade and in professional services, where allocative efficiency would increase by 20 and 16 log points, respectively.

While this figuring may provide a lower bound estimate to the extent that it does not account for the impact of regulation on resource flows across sectors, some caution is warranted in interpreting these effects owing to identification concerns and the lack of suitable instruments for industry level regulation across countries (see Section 4.3). These findings nonetheless support the core results in Section 5.1, and complement existing research which finds negative indirect effects of service regulation on AE in downstream industries, where endogeneity concerns are likely to less serious (see Arnold et al., 2008).

## 6. ROBUSTNESS CHECKS

### 6.1. Instrumental variables estimation to address the potential endogeneity of policies

The core results are robust to instrumental variable estimation, aimed at controlling for the potential endogeneity of product and labour market regulations. As outlined in Section 4.3, our main source of exogenous variation for governments' proclivity to intervene in the economy is based on the legal origins of commercial laws (see La Porta et al., 1999). We instrument barriers to entry with information on whether a country's legal system is based on British common law; French, German or Scandinavian civil law (see col. 1 of Table 5; in col.2, following Barseghyan (2008), we also account for geographic latitude). First stage estimates (reported in Table A6) suggest that countries

[27] These countries are the UK in the case of energy and land transport, Sweden in the case of retail and professional services, Finland for post and telecommunication and a group of countries in Air transport (for which the OECD regulation index is zero).

[28] In this case, the efficiency gains from lower regulation in each country-service industry are simply obtained as $\Delta\widehat{AE}_{cs} = \hat{\beta}\left(Reg_{BP,s} - Reg_{c,s}\right)$, where $\hat{\beta}$ is the coefficient estimated in Column 1 of Table 4. Sector specific effects ($\Delta\widehat{AE}_{cs}$) are then aggregated across EU countries for each regulated service $s$: $\Delta\widehat{AE}_{EU,s} = \sum_c \left(\frac{L_{c,s}}{L_{EU,s}} \Delta\widehat{AE}_{c,s}\right)$, where $L_{c,s}$ is employment in service industry $s$ and country $c$, and $L_{EU,s}$ is total EU employment in the regulated service $s$. Finally, the aggregate impact on efficiency in the EU is computed as: $\Delta\widehat{AE}_{EU,S} = \sum_s \left(\frac{L_{EU,s}}{L_{EU,S}} \Delta\widehat{AE}_{EU,s}\right)$, where $L_{EU,S}$ is total EU employment in regulated services.

with French and German legal origins are characterized by significantly higher administrative burdens than those with Scandinavian or British (the excluded category) origins. The second stage coefficients reinforce previous OLS findings, which uncover a negative relationship between entry regulation and AE. Similarly, the negative effect of EPL on AE is confirmed by the 2SLS estimation, which uses indicators regarding whether a country has a common law as opposed to a civil law system, and a refinement of this dichotomy identifying civil law systems with a single codified civil code, which are characterized by particularly constraining regulations (see Bassanini et al. 2009).

**Table 5: Public policies and allocative efficiency: instrumental variable estimates**

| | BTE | | EPL | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| VARIABLES | Legal origin | Legal origin and latitude | Common Law | Civil code |
| BTE X turnover | -0.019*** | -0.015*** | | |
| | (0.006) | (0.005) | | |
| EPLR X layoff | | | -0.081*** | -0.099*** |
| | | | (0.031) | (0.034) |
| R-squared | 0.579 | 0.584 | 0.595 | 0.588 |
| F-test on instruments | 188.6 | 130.4 | 292.3 | 29.8 |
| Overid.test (p-value) | | 0.07 | | |
| Observations | 834 | 834 | 834 | 834 |

Notes: 2SLS estimates of the OLS results in Columns 1 and 4 of Table 1. See Table A6 for the corresponding first stage estimates. The dependent variable is allocative efficiency as defined in (1), computed in 2005. See Table A3 for definition and sources of the explanatory interaction variables. All specifications include country and industry fixed effects. Observations are weighted by the number of firms used to compute the productivity decomposition in (1). Robust standard errors in parentheses; *** denotes statistical significance at the 1% level, ** significance at the 5% level, * significance at the 10% level.

## 6.2. Alternative productivity measures and clustered standard errors

While the use of (operating revenue) turnover-based labour productivity in measuring allocative efficiency is both common to the literature (see Bartelsman et al. 2013 for a discussion) and reflects data constraints (see Box 1), it may be problematic. A value added measure would be preferable, for example, to the extent that outsourcing is a relevant practice whose intensity responds to the strictness of some of the examined policies (e.g. EPL).[29] On the other hand, a TFP-based measure would allow, among other things, to insulate productivity from between firm differences in capital intensity. To address these issues, and despite the non-trivial implied reduction in the available number observations, we exploit the available data on value-added and capital stock to re-construct our dependent variable using a value-added measure of (log-)labour

[29] Outsourcing is likely to mechanically inflate turnover based labour productivity (relative to a value-added based measure) and reduce firm size (in terms of measured employment), inducing a fall in the observed within industry covariance between productivity and size – that is, AE. This will bias downwards our coefficient on, say, EPL if (within countries) firms with more intense reallocation needs are more prone to outsourcing (e.g. to exploit labour market flexibility in other countries).

productivity (AE-VA), and a Solow residual measure of (log-)TFP (AE-TFP), as outlined in Box 1. To provide an appropriate benchmark, for the same sub-sample of firms we also re-constructed an indicator of AE based on the operating revenue turnover measure of labour productivity used in the baseline estimates (AE-TURN). We then re-estimated our core model and compared the coefficient across specifications (see Table 6). While this exercise resulted in a significant reduction in both the sample size (685 country-sector cells compared to 834 cells in the core analysis), and the number of available firms per cell, there was little discernible difference in estimated coefficients – in terms of statistically significant and economic magnitude – between specifications.

Thus far, our estimates and inference are based on heteroskedasticity-robust standard errors, which might be underestimated in the case that shocks in (2) and (3) are correlated within industries and/or regions (or both). For the set of core results, Table 7 shows standard error estimates obtained allowing for 2-way clustering (by country and industry) following the procedures introduced by Cameron et al. (2006); for comparison, we also report the previous heteroskedasticty adjusted estimates. The estimated 2-way clustered standard errors in model (2) are very close and in many cases even lower than those adjusted for heteroskedasticity, suggesting that either there is negative intra-cluster correlation, or that the estimator is biased possibly due to the small number of clusters. Throughout specifications, the estimated coefficient remains statistically different from zero at conventional levels. Similar results are obtained when allowing for two-way clustering in the service regressions (specification (3)), except that the already weak evidence on the effects of barriers to foreign direct investment is no longer significant at conventional levels.[30]

[30] We also experimented with relaxing the assumption that observations in different clusters are independent, by defining broader groups to account for potential autocorrelation induced by geographical or technological proximity. Such an exercise is non-trivial with the data at hand, given that it implies a further reduction in the already low number of clusters. The geographical clusters were defined by experimenting with alternative spatial- and trade-based measures of proximity (i.e. grouping countries based on their main export or import partner); technological clusters were defined based on Input-Output linkages (i.e. grouping the original 42 industries based on having the same largest input-provider or output). The results were not significantly affected.

**Table 6: Public policies and allocative efficiency: robustness to alternative productivity measures**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| VARIABLES | BTE X turnover | BTE2 X turnover | Bankruptcy X turnover | EPLR X layoff | EPLO X turnover | FinDev X ExtFinDep | BankReg X ExtFinDep |
| | PANEL A: ALLOCATIVE EFFICIENCY – VALUE ADDED LABOUR PRODUCTIVITY | | | | | | |
| Policy X Exp | -0.008** | -0.014* | -0.001 | -0.021* | -0.008** | 0.016 | -0.017 |
| | (0.004) | (0.007) | (0.000) | (0.013) | (0.004) | (0.027) | (0.014) |
| AdjR2 | 0.430 | 0.429 | 0.426 | 0.428 | 0.428 | 0.425 | 0.427 |
| Observations | 685 | 685 | 685 | 685 | 685 | 682 | 682 |
| | PANEL B: ALLOCATIVE EFFICIENCY – VALUE ADDED TOTAL FACTOR PRODUCTIVITY (SOLOW RESIDUAL) | | | | | | |
| Policy X Exp | -0.011** | -0.018* | -0.001** | -0.039 | -0.012* | -0.019 | 0.026 |
| | (0.005) | (0.009) | (0.001) | (0.040) | (0.007) | (0.035) | (0.026) |
| AdjR2 | 0.353 | 0.353 | 0.352 | 0.353 | 0.352 | 0.348 | 0.352 |
| Observations | 688 | 688 | 688 | 688 | 688 | 688 | 688 |
| | PANEL C: ALLOCATIVE EFFICIENCY – TURNOVER LABOUR PRODUCTIVITY (BASELINE) | | | | | | |
| Policy X Exp | -0.010** | -0.016* | -0.001** | -0.026* | -0.018*** | 0.029 | -0.009 |
| | (0.004) | (0.009) | (0.000) | (0.016) | (0.006) | (0.031) | (0.017) |
| AdjR2 | 0.503 | 0.501 | 0.499 | 0.500 | 0.512 | 0.499 | 0.497 |
| Observations | 685 | 685 | 685 | 685 | 685 | 682 | 682 |

Notes: In panel A, the dependent variable is allocative efficiency as defined in (1), based on a value added based measure of labour productivity. In Panel B, a Solow residual measure of TFP and the firm's share of industry value added is used to compute allocative efficiency. Panel C shows the results for the baseline measure of allocative efficiency, which is based on a turnover based measure of labour productivity. See Table A3 for definitions and sources of the explanatory interaction variables. All specifications include country and industry fixed effects. Observations are weighted by the number of firms used to compute the productivity decomposition in (1). Robust standard errors in parentheses; *** denotes statistical significance at the 1% level, ** significance at the 5% level, * significance at the 10% level.

**Table 7: Clustering standard errors: consequences for core results**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| | | | PANEL A: FULL SAMPLE | | | | |
| VARIABLES | BTE X turnover | BTE2 X turnover | Bankruptcy X turnover | EPLR X layoff | EPLO X turnover | FinDev X ExtFinDep | BankReg X ExtFinDep |
| Policy X Exp | -0.0075** | -0.0149** | -0.0068*** | -0.0522*** | -0.0165** | 0.0150 | -0.0086 |
| | (0.0031) | (0.0059) | (0.0353) | (0.0153) | (0.0048) | (0.0202) | (0.0149) |
| | [0.0034] | [0.0067] | [0.0024] | [0.0092] | [0.0065] | [0.0254] | [0.0056] |
| AdjR2 | 0.556 | 0.557 | 0.553 | 0.567 | 0.565 | 0.553 | 0.553 |
| Observations | 834 | 834 | 834 | 834 | 834 | 828 | 828 |
| | | | PANEL B: SERVICES SECTOR | | | | |
| VARIABLES | Base | Base & FDI - 1 | Base & FDI - 2 | | | | |
| Service sector regulation | -0.1106** | -0.1295** | -0.1184* | | | | |
| | (0.0526) | (0.0483) | (0.0596) | | | | |
| | [0.0572] | [0.0533] | [0.0693] | | | | |
| FDI restrictions | | -0.9174 | -0.9616 | | | | |
| | | (0.4618) | (0.5293) | | | | |
| | | [0.649] | [0.740] | | | | |
| AdjR2 | 0.624 | 0.629 | 0.593 | | | | |
| Observations | 174 | 152 | 152 | | | | |

Notes: This table replicates the estimated coefficients shown in Table 1 (Panel A) and Table 4 (Panel A) but reports (in square brackets) estimates of two-way (country and industry level) clustered standard errors. For the purposes of this Table only, the bankruptcy X turnover variable has been divided by 10 in order to make the differences between the various standard errors clearer. Heteroscedasticity robust standard errors are also reported in parentheses. Statistical significance is computed based on the former. The dependent variable is allocative efficiency as defined in (1), computed in 2005. See Table A3 for definition and sources of the explanatory interaction variables. All specifications include country and industry fixed effects. Observations are weighted by the number of firms used to compute the productivity decomposition in (1). *** denotes statistical significance at the 1% level, ** significance at the 5% level, * significance at the 10% level

## 6.3. Sensitivity analysis

The baseline results reported in Table 1 are broadly robust to a number of sensitivity tests. First, while the core results are based on ORBIS data from 2005, they can be reproduced using data from 2006 and 2007 (the results are not reported for brevity). Second, the baseline coefficient estimates reported in Table 1 are broadly robust to excluding one country from the sample at a time (see Figure A1). While the estimated negative effect of EPL is always significant, there are a few instances where the barriers to entry and bankruptcy interactions are no longer significant at conventional levels when a given country is dropped from the regression. In both cases, however, the estimated coefficients are always significant at the 11-15 per cent level. Third, the results are also broadly robust to dropping one sector at a time. Again, the estimated coefficient on the EPL interaction is always significant, while the barriers to entry and bankruptcy interactions turn marginally insignificant at the 14 and 22 per cent levels respectively when the construction sector (NACE sector 45) is excluded.

Because the AE indexes capture within industry cross sectional covariances between size and productivity, it might be affected by the underlying (size and productivity) distributions reflecting, for example, differences in the minimum efficient scale of production or industry concentration. Unreported results show that accounting for this issue explicitly controlling for the quartiles of the firm size and the productivity distributions do not alter our estimates on policy variables which, if anything, turn out to be strengthened.

Finally, we test the sensitivity of our results to the inclusion of a host of other possibly relevant country-level structural variables such as institutional quality, corruption, reliance on professional management, openness to trade, infrastructure quality, education and country size. Each panel in Table A7 focuses on one of the policy variables for which we estimate significant coefficients in the paper when AE is the dependent variable (e.g. barriers to entry and exit, and EPL). In each column, we report the results obtained interacting each of the concurrent explanatory variables with the industry interaction term that is appropriate for the policy under scrutiny (e.g. layoffs in the case of EPL). Hence the coefficient is estimated exploiting cross-country variation in regulation that is not captured by the alternative policy variable. Broadly speaking, the results are not affected although the barriers to exit interaction term becomes marginally insignificant in a few instances.

## 6.4. Re-estimation with an alternative dependent variable

To further test the robustness of our core results, we re-estimate our models in a different sample of firms and with an alternate dependent variable. Since our measure of AE in Section 2 captures the extent to which the most productive firms in a given industry are also the largest, we would expect larger firms (e.g. firms with 250 or more employees) in a given industry to be relatively less productive in countries where framework policies hinder the efficient allocation of resources. While we focus on large firms because they have important implications for AE, they are also likely to be better represented in the ORBIS database (see Gonnard and Ragoussis 2012; Gal 2013). Accordingly, we estimate the following models using data for the full sample (4a) and services sample (4b):

$$P_{jc}^{250+} = \gamma \bar{P}_{jc} + \tilde{\beta}\left(Exp_j * Reg_c\right) + \tilde{\mu}_j + \tilde{\mu}_c + \tilde{\varepsilon}_{jc} \quad (4a)$$

$$P_{sc}^{250+} = \delta \bar{P}_{jc} + \tilde{\theta} Reg_{s,c} + \tilde{\eta}_s + \tilde{\eta}_c + \tilde{\epsilon}_{sc} \quad (4b)$$

where: $P^{250+}$ is the average labour productivity of firms with 250 or more employees in each country and industry; $\bar{P}_{jc}$ is average sector-wide labour productivity and the remaining terms are the same as described in equations (2) and (3). Controlling for average sectoral productivity is important as large firms might be less productive in a highly regulated economy simply because all firms in the industry have low productivity, but this would not necessarily have implications for AE.

The differences-in-differences estimation results (e.g. equation 4a) are contained in Table A8 and the service sector regressions (e.g. equation 4b) are stored in Table A9. Overall, these results are broadly consistent with those based on AE reported in Table 1 and Table 4, further highlighting the impact of framework policies on the efficiency of resource allocation.[31]

## 6.5. Unreported results on additional policies

Unreported results include additional explorations of the impact of policies such as various taxation variables (e.g. taxes on labour incomes, corporate profits and capital gains); tax incentives for research and development (R&D); intellectual property rights (IPR) regimes; supply of seed capital to GDP and quality of accounting standards.[32] These results were generally inconclusive. However, this does not necessarily imply that such policies have no effect on patterns of resource allocation. Instead, this ambiguity may reflect data limitations – our study focuses on the contribution of AE to the level of labour productivity at a given point in time and does not specifically model dynamic process, such as contribution of entry and exit to productivity growth.

[31] In the case of entry barriers and EPL, the diff-in-diff estimates are confirmed (if anything, slightly higher) adopting the IV approach described in section 6.1 (results are not reported for brevity).

[32] The impact of these country-level policies on allocative efficiency was modelled in a differences-in-differences framework. For example, the various taxation terms where interacted with sectoral measures of relative profitability and firm turnover; IPR regimes and R&D tax where interacted with sectoral measures of R&D and patenting intensity; seed capital and accounting standards where interacted with external finance dependency. Numerous other interactions were included but are not reported here for sake of brevity.

## 7. CONCLUSION AND POLICY IMPLICATIONS

Structural reforms have gained momentum in the aftermath of the Great Recession, largely because of market pressures and the co-ordinated efforts of multilateral agencies. A special emphasis has been placed on the functioning of product and labour markets, which – given the limited scope for demand side policies – are currently seen as the most important tools to strengthen productivity, competitiveness and ultimately economic growth in many OECD economies. The outcome of this reform momentum is most clearly illustrated with respect to EPL, with more than one-third of OECD countries having introduced, or passed, reforms that reduce its stringency, at least as this is captured by available indicators (OECD, 2013a). But efforts have also focused on improving financial stability through more efficient bank supervision, and on lowering the degree of product market regulation, particularly in those OECD countries that were more severely hit by the onset of the euro area crisis.

The emphasis on market regulation is largely motivated by the idea that more flexible economies would benefit from a more intense and rapid reallocation of resources towards their most productive use. Indeed, recent economic research showed that all major economies are characterized by such reallocation process, which plays a relevant role for aggregate productivity. The purpose of this work was to provide a comprehensive assessment of the importance of regulatory policies for productivity *via* the resource allocation channel. We exploited large harmonized firm level data set covering a cross-section of non-farm business industries and a large set of measures of market regulation to explore the impact of structural policies on the efficiency of resource allocation in a sample of 21 OECD countries.

Our results suggest that there is an economically and statistically robust negative relationship between policy-induced frictions and productivity, though the specific channel depends on the policy considered. In the case of employment protection legislation, product market regulations (including barriers to entry and bankruptcy legislation) and FDI restrictions, this is largely traceable to the worsening of allocative efficiency – that is, a reduced ability of an economy to channel resources to more productive firms. Moreover, regulations are more disruptive to AE and productivity in innovative sectors, which is consistent with growing body of evidence which suggests that poorly designed policies tend to penalise firms that operate closest to technology frontier (Aghion and Howitt, 2006; Arnold et al., 2008; Andrews and Criscuolo, 2013). By contrast, financial market under-development tends to be associated with a lower (un-weighted) average productivity. In line with recent model based evidence (Midrigan and Xu, 2013), this finding suggests that the bulk of aggregate productivity losses traceable to financial frictions arise due to a worse selection of firms at entry rather than to resource misallocation.

Our simulated “scenarios” of the gains from plausible EU-wide reforms of product and labour markets indicate a substantial boost to productivity. For example, using the latest available regulation data (referred to 2008), we estimate that a policy that were to lower entry barriers in each country to the lowest level observed within the EU could double allocative efficiency in the entire area. This would account for a large part of the gap relative to the United States and imply that labor productivity in the EU (as defined in section 2) would be 15 per cent higher owing to a more efficient allocation of resources. The gains would be particularly high for the countries with lowest (in fact, highly negative) estimated allocative efficiency, Poland and Greece, and higher than average in troubled euro area countries such as Portugal and Spain. Significant productivity gains are also predicted by analogous exercises simulating an EU wide reduction of anti-competitive barriers in regulated services (retail trade, energy, transport and communications, and professional services) or the reform of employment protection legislation on worker dismissal.

Hence, our paper suggests that several of the reforms recently undertook in many advanced economies might have a positive impact on productivity through their effect on the efficiency of resource allocation. Indeed, such reforms were particularly evident in countries, as those of Southern Europe, where regulation was most stringent before the onset of the crisis, and allocative efficiency, at least as measured in our data, was generally low. And yet, there seems to be room for further action, as the policy recommendations of multilateral agencies demonstrate.[33] One reason for this is that structural reforms in many areas often take place gradually, with incremental policy changes introduced in sequential rounds. Furthermore, countries often prove unable to implement and enforce legislated reforms in a coherent and consistent fashion. This is likely to occur when the legislation is fragmented, poorly drafted, or ambiguous; when local regulation is inconsistent with national legislation; when implementation or enforcement is ineffective because of slow moving courts, inefficiency of the public administration and corruption. While these dimensions of structural reforms are not always captured by the indicators used in the present analysis, they highlight the fact that the effectiveness of reforms goes beyond the legislative step, which suggests that ongoing qualitative structural surveillance exercises by multilateral organisations are an important compliment to the analysis conducted in this paper.

Because it represents a major shock to most economies, the current recession provides important opportunities to boost long-term productivity via the reallocation of resources away from inefficient firms and business activities towards more productive ones. Based on our findings, recent structural reforms should enhance the likelihood that these gains are realised and in a timely fashion. And yet, the crisis also highlights that reallocation entails costs for firms, workers and governments, which are not considered in the current paper but should nonetheless be a concern for policymakers. For example, the adverse consequences of reforms to job protection legislation could be mitigated if accompanied by broader mechanisms that insure workers against labour market risk, such as well-designed social safety nets and portable health and pension benefits.

[33] See, for example, the OECD's recent stock take of structural reforms (OECD, 2013b). Progress in the policy action was also required in the latest Country Specific Recommendations issued by the European Commission within the surveillance mechanism called Macroeconomic Imbalance Procedure (MIP); and World Bank indicators highlight the need for further reforms in Bankruptcy legislation (World Bank, 2013).

Given the limitations of our approach, more research into the links between framework policies and resource allocation is required. Further progress could be made with more comprehensive data, which would allow, in particular, a more accurate measure of productivity. Moreover, we focus only on the correlation between productivity and size at a single point in time (i.e. static allocative efficiency). While this metric should in principle reflect patterns of resource reallocation across incumbent firms and firm turnover (i.e. entry and exit) in preceding periods, more direct evidence on the influence of policies on dynamic allocative efficiency would be desirable. Given the limitations of ORBIS in reliability identifying entrants and exits over time, however, further evidence on the link between policies and firm turnover will depend on the availability of administrative data from national sources (i.e. business registers).[34] From a policy perspective, a more direct focus on firm specific distortions (such as taxes and subsidies) would be desirable, both to extend the analysis to other relevant policy measures, and to infer an average or direct impact of policies, which is prevented by our differences-in-differences framework.

[34] Existing data do nonetheless allow analysing the impact of framework policies on other aspects of dynamic resource allocation, such as reallocation across incumbent firms (Andrews and Criscuolo, 2013). Indeed, to the extent that cross-country differences in the post-entry performance of firms tends to be more marked than differences in entry and exit patterns this would seem like a worthy endeavour.

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