FACULTEIT ECONOMIE
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TWEEKERKENSTRAAT 2
B-9000 GENT
Tel. : 32 - (0)9 – 264.34.61
Fax. : 32 - (0)9 – 264.35.92

WORKING PAPER

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Tim Buyse
Freddy Heylen
Ruben Schoonackers

First version: September 2015
This version: December 2015
2015/911

D/2015/7012/13

On the role of public policies and wage formation for business investment in R&D: a long-run panel analysis*

Tim Buyse[1], Freddy Heylen[1] and Ruben Schoonackers[2†‡]

[1] *Ghent University and SHERPPA*

[2] *National Bank of Belgium, Ghent University*

December 21, 2015

Abstract

This paper studies the drivers of business funded and performed R&D in a panel of 14 OECD countries since 1981. More specifically, we investigate the effects of public R&D related policies and wage formation. Following Pesaran (Econometrica, 2006) and Kapetanios et al. (Journal of Econometrics, 2011), our empirical strategy allows for cross-sectionally correlated error terms due to the presence of unobserved common factors, which are potentially non-stationary. We find that tax incentives are effective. Public funding (subsidization) of R&D performed by firms can also be effective if subsidies are not too low, neither too high. R&D performed within the government sector and within institutions of higher education is basically neutral with respect to business R&D. We find no evidence for crowding out, nor for complementarity. The higher education sector may, however, indirectly be of great significance. Our results reveal human capital accumulation at the tertiary level as a key driver of business R&D in the OECD during the last decades. As to the impact of wage formation, using an indicator for wage pressure developed by Blanchard (Economic Policy, 2006), we find that wage moderation may contribute to innovation, but only in fairly closed economies and in economies with flexible labour markets. In highly open economies and economies with rigid labour markets rather the opposite holds. In these economies high wage pressure may enhance creative destruction and force firms to innovate as a competitive strategy. Our results show that a careful treatment of the properties of the data is crucial.

JEL Classification: E22, J30, O31, O38, O57

Keywords: R&D, technology policy, wage formation, panel cointegration

*We would like to thank Gerdie Everaert, Glenn Rayp, Markus Eberhardt, Tino Berger, Emmanuel Dhyne and Stefan Van Parys for constructive comments and suggestions. Any remaining errors are ours.

†Corresponding author. Research Department, National Bank of Belgium, Berlaimontlaan 14, B-1000 Brussels, Belgium. Email: Ruben.Schoonackers@nbb.be. Phone: +32 2 221 51 44.

‡The opinions expressed are strictly those of the authors and do not necessarily reflect the views of the National Bank of Belgium.

1 Introduction

Ageing and rising pressure on the welfare state force all OECD countries to develop effective employment and growth policies. When it comes to long-run growth, both the theoretical and empirical literature recognize investment in research and development (R&D) as a major factor (see Romer, 1990; Aghion and Howitt, 1992; Coe and Helpman, 1995; Coe, Helpman, and Hoffmaister, 2009). Numerous studies have therefore investigated the determinants of business investment in R&D in many countries, both at the micro and the macro level. Guellec and Van Pottelsberghe (1997, 2003) were the first to provide an explanation at the macro level in a panel of 17 OECD countries. In their seminal paper, they paid particular attention to the role of public policies organized to stimulate private R&D investment i.e. tax incentives, public funding of R&D projects in the business sector, expenditures on R&D within the government sector and R&D spending in institutions of higher education.

Our research is inspired by two gaps in the empirical macro literature on the drivers of business R&D. A first one relates to the impact of wage formation. Today, OECD countries are not only called upon to develop effective growth policies, but also to create jobs and to raise employment rates. To reach this goal, many countries adopt outspoken wage moderation policies. Interestingly, these policies also affect incentives and available resources for firms to innovate and invest in R&D. On the employer side, it is often argued that wage moderation is an important factor to maintain firm profitability, which is a key condition for investment in R&D. Several researchers have, however, argued that an excessive focus on wage moderation may kill incentives to innovate (e.g Kleinknecht, 1998). Wage moderation may for example increase the survival probability of the least innovative firms and retard the process of creative destruction. Weighing on the purchasing power of households, outspoken wage moderation may also lead to lower demand-driven innovations as demand for new products and services falls. Conversely, high wage pressure may force firms to innovate as a key element in their competitive strategy. To the best of our knowledge, despite its theoretical importance, rigorous cross-country empirical work on these conflicting hypotheses has never been done.

A second gap in the existing empirical macro literature on the determinants of R&D investment is methodological. A key characteristic of new technology and knowledge is that they may spill over to other firms and countries, so that all may benefit from an improvement in the world level of technology, although not necessarily to the same extent (Coe, Helpman, and Hoffmaister, 2009; Everaert, Heylen, and Schoonackers, 2015). Eberhardt, Helmers, and Strauss (2013) have shown that these spillovers affect firms' private returns to R&D and therefore business R&D investment. A crucial econometric issue, however, follows from the fact that the world level of technology and knowledge is largely unobserved. Technology spillovers will then manifest themselves in standard panel R&D regressions as cross-sectional dependence in the error terms, induced by an unobserved common factor. Guellec and Van Pottelsberghe (1997, 2003) and subsequent macro research (e.g. Falk, 2006; Westmore, 2014) have neglected this issue. If omitted common factors are correlated with the included explanatory variables, estimated parameters will be biased and inconsistent. Even worse, when unobserved common factors are non-stationary, standard estimators yield spurious results.

Our contribution in this paper is to study the determinants of business investment in R&D in 14 OECD countries in the period 1981-2012, with a special focus on the role of wage formation and by adopting an empirical strategy that deals with cross-sectionally correlated error terms due to the presence of unobserved common factors. Figure 1 shows the data. To be precise, they include the expenditures on R&D performed *and* financed by the business sector. They are expressed in real per capita terms and in 2010 PPP dollars. Further in this paper we characterize this variable as $BERD$, briefly defined as business R&D investment. Huge cross-country differences stand out, both in the level and in the evolution of R&D, making an empirical analysis highly relevant. To quantify wage formation, we follow Blanchard (2006) and use insights from growth theory. The approach is to compare actual (growth of) real wages with the so-called 'warranted' real wage (growth). The latter is determined by the rate of Harrod-neutral technical progress. In growth theory, this is the rate of real wage growth consistent with stable employment along a balanced growth path. We will speak of high wage pressure when actual real wage growth is higher than the rate of technical progress. A positive and increasing wage gap will then arise. We speak of wage moderation when actual real wage growth is lower than the rate of technical progress. The wage gap then declines and may turn negative. Next to the role of wage pressure, we also test the impact of public policies organized to stimulate business R&D investment, in line with Guellec and Van Pottelsberghe (2003). To estimate our model, we use the common correlated effects pooled (CCEP) estimator of Pesaran (2006). This estimator controls for unobserved common factors by adding cross-sectional averages of the data. As shown by Kapetanios, Pesaran, and Yamagata (2011), this approach is also valid in a non-stationary panel context.

Our main findings are the following. First of all, we learn from our results that a careful treatment of the properties of the data is crucial. The empirical analysis reveals significant cross-sectional correlation in levels and in first-differences for most variables. All variables are also found to be non-stationary. For most variables the non-stationarity is induced by an (unobserved) common factor. The use of the CCEP estimator is therefore highly justified. Second, the effects of wage pressure are significant but not uniform. We find that in economies where firms face relatively little (foreign) competition and dispose of flexibility to adjust their employed labour force because employment protection legislation is soft, high wage pressure has negative effects on business R&D investment. In open economies where firms face sharp (foreign) competition and run their activities in a rather rigid and regulated labour environment, however, the opposite seems to happen. In such economies - think of many European economies - firms that do not innovate cannot survive when wage pressure is high. Rising wages thus enhance creative destruction and force all firms to innovate as a competitive strategy. Third, our empirical analysis reveals various ways in which governments can effectively promote business R&D investment. We observe that both tax incentives and public funding (subsidization) of R&D projects in the business sector can work, if chosen carefully. This condition applies in particular to public funding. For this policy instrument, we confirm an earlier finding of Guellec and Van Pottelsberghe (2003) that the relationship between subsidization and business R&D investment is inverted U-shaped. That is, subsidies encourage private firms to raise their own R&D spending if these subsidies are not too low neither too high. The optimal subsidization rate (at the macro level) may be somewhere between 6 % and 10 %. The results also show that the available stock of high-skilled human capital is an important driver of business R&D investment implying that governments should invest in schooling in order to increase the percentage of the population with a higher degree. Finally, we find that R&D investment within the government sector and within universities will also have positive effects on aggregate R&D spending. Most of our results predict a one-to-one effect from higher spending within the public sector to aggregate R&D. In other words, neither the idea that public R&D would crowd out private R&D spending, nor the idea of complementarity between the two, find support in our results.

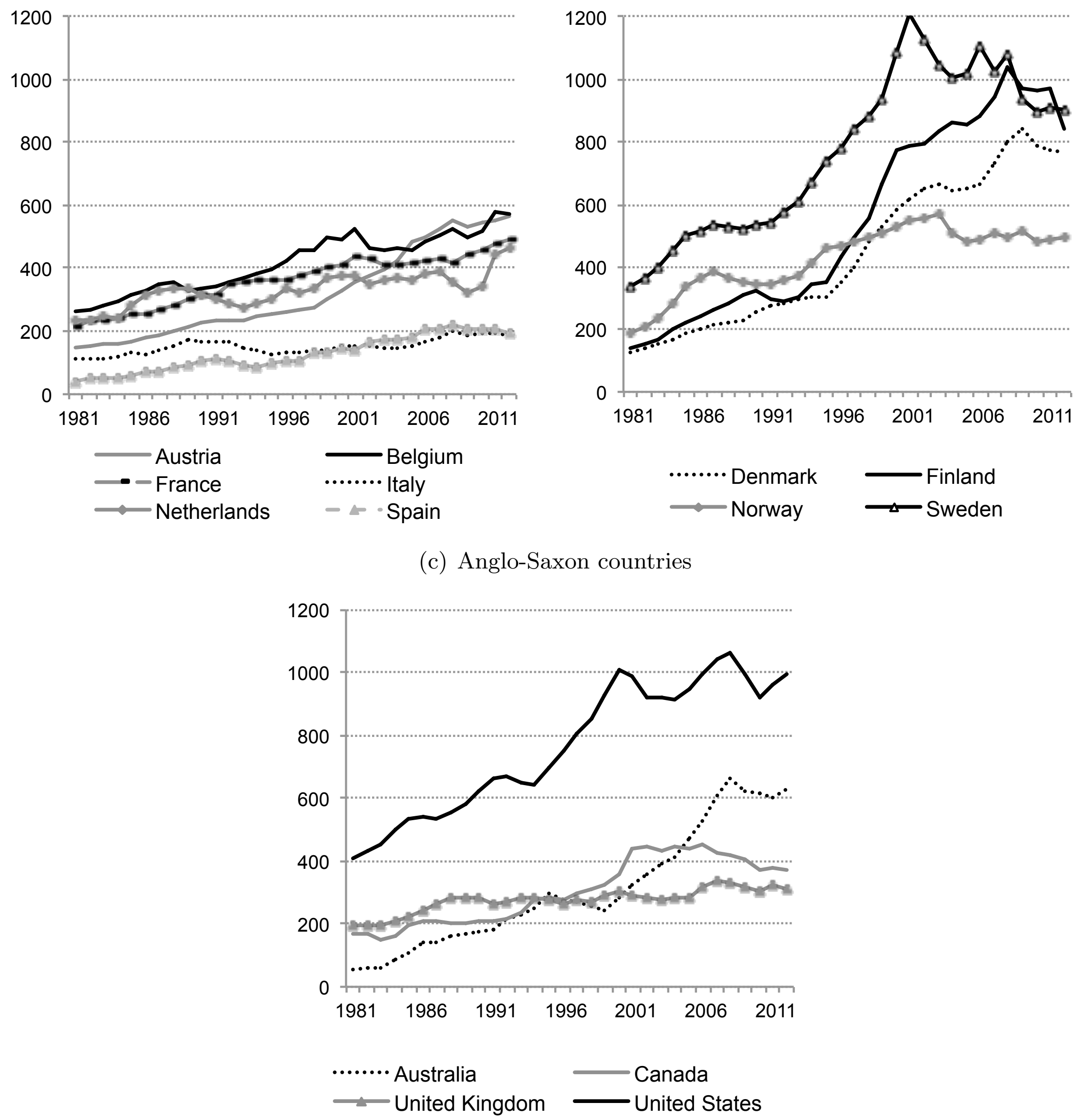
Our focus on aggregate business R&D investment in this paper is not common in the literature. In comparative perspective, many more studies have investigated R&D expenditures at the firm or the industry level, see e.g. the surveys in David, Hall, and Toole (2000) and Becker (2015). Yet, there are very good reasons why an analysis of macroeconomic data is important. A first one relates to the indirect effects or externalities of policies. For example, if individual firms benefit from R&D investment subsidies, this may boost their innovation activity. At the same time, however, also other firms may be affected. Competing firms may suffer because of the advantage given to a direct competitor. Due to falling rates of return they may reduce their R&D investment. On the other hand, downstream customers in the supply chain may benefit from knowledge spillovers induced by the innovating firm. They may raise their R&D investment. Similar externalities can occur between industries (Guellec and Van Pottelsberghe, 2003). The potential presence of these external effects makes the case for an empirical analysis at the macro level. A second reason follows from the observation that (firms in) different industries may react differently to changes in the drivers of R&D, for example because market environment and institutions are different. In that sense, the response of R&D investment to rising wage pressure may be different in manufacturing sectors than in services. For policy makers it will be highly interesting also to know what the response is at the aggregate level.

Figure 1: Business financed and performed R&D expenditures ($BERD$) in 14 OECD countries (real per capita, 2010 PPP dollars)

(a) Euro area countries

(b) Nordic countries

(c) Anglo-Saxon countries



Note: further details and data sources are described in Appendix C

The remainder of this paper is structured as follows. Section 2 contains a brief survey of the literature on public policy instruments to encourage business R&D investment, and on their effects. This section also reviews the conflicting hypotheses regarding the influence of wage formation on innovation. Section 3 discusses important properties of the data, sets out the empirical model and discusses the econometric methodology. In Section 4 we report our estimation results. Section 5 concludes the paper.

2 Drivers of business R&D intensity: literature

Boosting R&D intensity is one of the top priorities of OECD countries today. The Europe 2020 targets include that 3% of EU GDP has to be invested in R&D and innovation (public and private combined) by 2020. To stimulate private R&D intensity, governments have different instruments at their disposal. These instruments are used to offset market failure in the allocation of resources to long-term and risky investment, which are key characteristics of R&D investment. As a result, private investment in R&D is mostly lower than socially optimal, thus justifying government support.

Section 2.1 discusses existing public policy instruments and some of the empirical evidence on their impact. In Section 2.2. we review the literature regarding the effects of wage formation and some underlying labour market characteristics on R&D investment and innovation. Various countries have institutionalized wage moderation or wage control mechanisms in the second half of the 1980s or early 1990s. Other countries have decentralized wage bargaining and introduced legislation to reduce union power, also contributing to wage moderation. While most will agree that these policies have positive effects on employment and competitiveness, at least in the short run, their possible long-run effects on a country's innovative capacity occur much less clear. In our discussion of the arguments for and against wage moderation, we also pay attention to the potential impact of the institutional environment within which wage formation takes place. We end with a brief explanation of the role of product market characteristics.

2.1 Public policy instruments

Traditionally, R&D policy can be subdivided in direct support (such as public sector R&D and direct R&D subsidies) and indirect support (such as R&D tax incentives). In addition, governments may also provide support for the university research system and the formation of high-skilled human capital as for formal R&D cooperation between institutions. In this section, we point at existing, mostly empirical, evidence on the impact of policy support measures on private R&D expenditures.

Public sector R&D and government funding of R&D in the business sector

Among the most frequently used public policy instruments to support R&D are public sector R&D and government funding of private investment in R&D. The former refers to direct R&D expenditures by public research institutions (intramural) and universities. The latter may either take the form of grants or subsidies, where the results of the R&D belong to the private performer, or it may concern funding aimed at the procurement of R&D, where the results belong to a recipient that is not necessarily the performer. An important question in the literature is whether these instruments are effective tools to stimulate private investment in R&D, or not. Effects may be positive when public sector involvement reduces the cost and risk of research for the industry. One way to achieve this is by conducting basic or fundamental research (where the wedge between private and social returns is probably the highest) and by making its results publicly available. Effects may also be positive when public resources lift potential cash constraints in private firms or provide a buffer when high financial risk is involved. Guellec and Van Pottelsberghe (2003), however, see three reasons why one may question the effectiveness of public spending on R&D. As a worst case scenario, public spending may even crowd out private R&D. First, government spending on R&D may increase the demand for researchers, which may raise these researchers' wages and make private R&D investment more expensive. This potential source of crowding out is most likely to occur if there is a shortage in the most decisive factor of the R&D process. That is if high-skilled labour is scarce. Second, public sector money can act as a substitute to private money. In other words, governments may execute or subsidize projects that would have been implemented anyway such that the same investment is performed with public instead of private money, without any increase in total R&D. Third, the allocation of funds by the government generally occurs less efficiently than by market forces, thereby distorting competition and resource allocation.

As to the empirical evidence on the effects of R&D in the public sector, Goolsbee (1998), for the United States, finds evidence of crowding out of private funding through raising wages of scientists and engineers. Guellec and Van Pottelsberghe (2003) (their Table III) report results for a panel of 17 OECD countries that are consistent with this observation. According to their findings, a one euro increase in R&D expenditures within the government sector tends to imply a 0.38 euro decline in business expenditures in the long run. Although this supports the hypothesis of crowding out, the net aggregate effect of intramural government R&D would still seem to be positive. That is, crowding out is only partial. As to R&D expenditures in universities, Guellec and Van Pottelsberghe (2003) find an effect on private spending that is basically zero, leaving an aggregate net effect of 1. Falk (2006), on the other hand, finds indications of a significant positive impact of R&D in the higher education sector on business R&D.

When it comes to the effects of direct funding by the government of R&D in the private sector, David, Hall, and Toole (2000) report that one third of available, mostly firm-level, studies find substitution effects. Overall the authors conclude that the empirical literature is inconclusive about the net impact of public R&D subsidies. Falk (2006) and Bassanini and Ernst (2002) are also inconclusive or report negligible effects. By contrast, Guellec and Van Pottelsberghe (2003) find that the net long-run impact of R&D subsidies on private R&D investment is positive. A one euro increase in government funded R&D in the business sector would induce an additional 0.7 euro of private spending. Lach (2002) also finds that public R&D subsidies stimulate private R&D expenditures in the long run. So does most of the more recent research. While Westmore (2014) finds positive effects of public R&D subsidies in a macro panel of OECD countries, Becker (2015), in her survey, includes many micro based studies that support the idea of additionality (see for instance Duguet, 2004; Carboni, 2011; Czarnitzki and Hussinger, 2004; Aerts and Schmidt, 2008; Hussinger, 2008; Cerulli and Poti, 2012; Oezcelik and Taymaz, 2008; Bloch and Graversen, 2012).

The effects of R&D subsidies need not be homogeneous, however. For instance, Jaumotte and Pain (2005) show that at a firm level the positive effect of R&D subsidies is more pronounced when firms are cash-constrained. In fact, there is broader empirical evidence that public subsidies are more effective drivers of R&D in small (financially constrained) firms. In the same spirit Czarnitzki and Ebersberger (2010) underscore the importance of aimed targeting of subsidies. These authors observe that in many cases most funding is awarded to larger firms that would have performed the R&D even in the absence of the public subsidy. Some studies also report heterogeneity in effects depending on the size of public subsidies. Guellec and Van Pottelsberghe (2003), for instance, find an inverted U-shape, where the strongest positive effects on private R&D can be observed for public subsidy rates of $4 - 11$ %, while rates that are too high (>20%) tend to generate negative (substitution) effects. Gorg and Strobl (2007) confirm these findings. Becker (2015) concludes that this non-linear effect suggests that it could be more effective to provide intermediate support levels to a larger number of firms than a large amount of support to fewer firms.

R&D tax incentives

The policy mix aimed at stimulating business R&D and innovation has seen growing use of R&D tax incentives. Such measures are indirect since the decision to use them, and the decision on how to use them, remains with the company. They are thus considered to be more market-oriented than for instance direct subsidies. Companies investing in R&D are eligible to claim tax reductions against their payable tax (Warda, 2001). As such, R&D tax incentives reduce the marginal cost of R&D spending and are also more neutral (i.e. less distortive) than direct R&D subsidies. In general, while direct subsidies are more targeted towards long-term research, R&D tax schemes are more likely to encourage short-term applied research and boost incremental innovation rather than radical breakthroughs (EC, 2003; OECD, 2014).

Fiscal incentives for R&D may take on various forms such as R&D tax credits, which are present in countries such as France, Belgium and the UK (OECD, 2014; EC, 2003). These tax credits are deducted from the corporate income tax and are applicable either to the level of R&D expenditures or to the increase in these expenditures with respect to a given base. Alternatively, some countries, such as Canada, Denmark and the UK, allow for the immediate or accelerated depreciation of investment in machinery, equipment, and buildings devoted to R&D activities (Warda, 2013; Falk, 2006). Finally, tax incentives do not only find application in the corporate income tax, but may also apply to the personal income tax, as in the Netherlands and Belgium, or to the value added tax (or other taxes such as consumption, land or property) (OECD, 2014).

An often used indicator reflecting the overall generosity of R&D tax incentives in a country is the so-called B-index (Warda, 2001). It is a composite index that is computed as the present value of income before taxes necessary to cover the initial cost of R&D investment and to pay the corporate income tax so that it becomes profitable to perform research activities (Warda, 2001). Algebraically, the B-index is equal to the after-tax cost of a one euro expenditure on R&D divided by one minus the corporate income tax rate. The after-tax cost is the net cost of investing in R&D, taking account of all available tax incentives (corporate income tax rates, R&D tax credits and allowances, depreciation rates). The more favourable a country's tax treatment of R&D investment, the lower its B-index.

Hall and Van Reenen (2000) find that most studies in the pre 2000 literature show positive effects of fiscal incentives on R&D expenditures. More recent research into the effectiveness of tax credits is even more unanimous in concluding that there are positive R&D effects (Becker, 2015). For instance, both Bloom, Griffith, and Van Reenen (2002) and Guellec and Van Pottelsberghe (2003) find significant negative coefficients on the B-index in their regressions explaining business R&D expenditures. Bloom, Griffith, and Van Reenen (2002) estimate that a 10% tax cut induced fall in the cost of R&D induces just over a 1% rise in the level of R&D in the short run, and just under a 10% rise in R&D in the long run. That is, they find a long-run elasticity of R&D with respect to the user cost of just below 1 in absolute value. Long-run elasticities vary between modest estimates of -0.14 (Bernstein and Mamuneas, 2005; Baghana and Mohnen, 2009) and strong ones of about -1.5 (as in Harris, Li, and Trainor, 2009; Parisi and Sembenelli, 2003). Most studies find elasticities in between these extremes (Lokshin and Mohnen, 2012; Koga, 2003; Mulkay and Mairesse, 2013).

Knowledge spillovers from the university research system and the formation of high-skilled human capital

Governments may resort to other than the traditional policy instruments to support private R&D expenditures. Some recent studies indicate the relevance of knowledge spillovers from university research to firms, enhancing technological opportunities and the productivity of private R&D, for example through personal interactions, university spin-offs and consultancy. Most empirical studies on this topic indeed find positive (geographically localized) knowledge externalities from university research to private R&D (see for instance Jaffe, 1989; Autant-Bernard, 2001; Karlsson and Andersson, 2009). Policies may thus aim to facilitate and support the formation of regional clusters of university and private R&D activity to exploit agglomeration economies. An important role in this context is played by the (increased) availability of high-skilled personnel trained by universities. Some studies do indeed find important positive R&D effects of high-skilled human capital resources[1]. Education policies and human capital investment thus also have a role in increasing private R&D.

2.2 Wage formation, labour and product market characteristics and innovation

The monitoring of wage formation is an important feature of many OECD countries' economic policy as it has a direct impact on employment and a country's competitiveness. Expected positive effects on employment generally underlie arguments in favour of wage moderation (see e.g. Bovenberg, 1997). Lower wages may increase firm profitability, generating more resources for investment. They may improve the competitiveness of domestic firms and raise exports. And they may make production more labour intensive. It then comes as no surprise that in many European countries wage moderation policies have become institutionalized. Germany's success is currently often taken as guiding inspiration (Heylen and Buyse, 2012).

An important additional element, especially from a long-run perspective, is the possible impact of wage formation on a country's innovative capacity. If high (excessive) wages reduce R&D investment, their negative effects on employment and competitiveness would be multiplied. On the other hand, if wage pressure promotes innovation, negative effects on competitiveness would be limited to the short run, whereas in the long run competitiveness

[1] Variables that are considered are the availability of highly qualified scientists and engineers (Adams, Chiang, and Starkey, 2001; Adams, Chiang, and Jensen, 2003; Becker and Pain, 2008), the share of workers with higher education in the total number of workers (Garcia and Mohnen, 2010), the share of the population with tertiary education in the total working age population (Wang, 2010) and the years of formal schooling (Kanwar and Evenson, 2003).

and employment would rise. In the literature both theoretical cases have been made. The first one goes as follows. If a focus on wage restraint is missing, rents from innovation may be appropriated by unions through higher wage claims. This may reduce firms' willingness and resources to innovate. An early statement of this argument was the so-called *hold-up problem* under incomplete contracts (Grout, 1984; Menezes-Filho and Van Reenen, 2003). In more recent work, Ulph and Ulph (1994) confirm this argument in a right-to-manage model where unions and firms bargain only over the wage. The main factor driving firms in their innovation efforts in their model is the expected difference between the profits that the firm can earn once it has successfully innovated and the profits that it would earn otherwise. In this setup high (excessive) wages represent a 'tax' that unions impose on the investment and the success of the firm. Lower R&D investment would be the result. Conversely, a focus on wage moderation would imply higher R&D. Other authors, however, have challenged this expectation (see e.g. Kleinknecht, 1994, 1998; Kleinknecht and Naastepad, 2004). One of their main arguments is that long-lasting wage moderation raises the survival probability of low-productive firms and non-innovators, slowing down the process of creative destruction. In a regime of wage increases and wage pressure, by contrast, the balance would shift and lack of innovation would no longer - or much less - be an option. In the framework of Ulph and Ulph (1994), this argument would imply that high wage pressure no longer reduces, but raises the profit differential between innovating and not innovating. The explanation is the very negative outcome (failure of the firm) in the non-innovating case. Intuitively, this idea raises a number of interesting extensions. One would expect this positive effect of high wage pressure to exist mainly in a very competitive environment and when firms lack the flexibility to adjust their (expensive) labour force. What we have in mind are very open economies and/or economies with highly deregulated product markets, but a very regulated labour market (e.g. extensive employment protection legislation). It will be exactly in such an environment that high wages and lack of innovation imply huge losses and the risk of bankruptcy. In these economies innovation will be firms' only possible competitive strategy.

Theory being inconclusive, what do we know about the impact of wage moderation on innovation and R&D empirically? First of all, it must be said that existing empirical work directly relating wage formation and innovation is very scarce. Most studies that analyse the effect of labour markets on innovation focus on aspects of numerical flexibility, such as the existence of flexible employment contracts, or functional flexibility such as the possibility of outsourcing or temporary employment. For instance, Bassanini and Ernst (2002) have estimated the impact of labour market regulation on an industry's R&D intensity in a cross-section of 18 manufacturing industries and 18 OECD countries. More recently, Murphy, Siedschlag, and McQuinn (2012) examined the impact of the strictness of employment protection legislation on innovation intensity in the OECD. Univocal results are hard to find. Observed effects depend on the system of industrial relations and the characteristics of industries. We know of only one study that has directly analyzed the impact of wage changes on innovation. Pieroni and Pompei (2008) find, for a panel of Italian manufacturing industries, that wage increases are positively related to the number of patents (their proxy for innovation). However, the authors only look at absolute wages and do not include an adequate measure of wage pressure (wage moderation) as we will do (See Section 3.1.1).

Next to the impact of labour market institutions, a growing number of researchers have studied the role of product market characteristics (in particular product market competition) on innovation. In a highly cited contribution, Aghion, Bloom, Griffith, and Howitt (2005) put forward an inverted U-shaped relationship between the degree of competition and investment in innovation. The argument goes as follows. When competition is low to begin with, the economy is expected to consist of a higher fraction of sectors with 'neck-and-neck' competing firms. Product market deregulation will induce these neck-and-neck firms to innovate in order to escape competition, since the incremental value of getting ahead rises in the degree of competition. When competition is high to begin with, however, the economy will have a higher fraction of sectors with one technological leader and many laggards. Further deregulation then has negative effects on innovation. Since more competition reduces the net rent that can be captured by laggards who succeed in catching up, the incentives for them to try will get weaker. This is the Schumpeterian effect of more competition. Although our focus in this paper is not on product market characteristics, we will control for them in our empirical work. Moreover, as we have mentioned above, the degree of product market competition may also be a factor that changes the effect of wage pressure on firms' investment in R&D.

3 Empirical analysis

Our empirical analysis follows Guellec and Van Pottelsberghe (2003) and relies on a simple R&D investment model that considers real per capita business funded and performed R&D investment ($BERD_{it}$) to be a function of a mix of policy instruments ($POLICY_{it}$), discussed in Section 2, and of real per capita value added generated by the business sector (VA_{it}). A set of other possible determinants of business R&D investment are included in (Z_{it}). Finally, we explicitly investigate the possible impact of wage formation ($WAGE_{it}$) on $BERD_{it}$,

$$BERD_{it} = f(VA_{it}, POLICY_{it}, Z_{it}, WAGE_{it}), \tag{1}$$

where subscripts i and t respectively denote the ith country and tth period. The exact functional form for equation (1) will depend on the discussion of the properties of the data in Section 3.1.3.

3.1 A first look at the data

3.1.1 Data and sources

We analyse the determinants of real per capita business R&D for a group of 14 OECD countries[2] using yearly data over the period 1981-2012. An overview of the construction of all data and their sources can be found in Appendix C.

Figure 1 reported wide variation across the countries in our sample, both in the level and the evolution of business expenditure on R&D. Policy instruments included in $POLICY_{it}$ are real per capita government intramural expenditure on R&D ($GOVERD_{it}$) and real per capita expenditure on R&D in the higher education sector ($HERD_{it}$). As a measure for direct R&D subsidies ($SUBS_{it}$) we include real per capita government funded expenditure on R&D performed in the business sector. A final measure included in $POLICY_{it}$ is the B-index ($BINDEX_{it}$), which captures direct R&D tax incentives[3]. In our empirical analysis, $VA_{it}, BERD_{it}$ and all variables in $POLICY_{it}$ will be expressed in logarithms.

Regarding the variables in Z_{it}, we focus on three possible determinants of business sector R&D, i.e. the degree of openness of the economy ($OPEN_{it}$), the available stock of high-skilled human capital ($HCAP_{it}$) in a country and the degree of product market regulation (PMR_{it}). The degree of openness is included to account for international trade, which is an important channel of knowledge and technology transfers across countries raising the return to domestic business R&D investment (e.g. Coe and Helpman, 1995; Coe, Helpman, and Hoffmaister, 2009; Acharya and Keller, 2009). Based on this argument, we expect a positive effect from a higher degree of openness on $BERD$. The stock of high-skilled human capital is considered because of its potential double impact on business R&D investment. First, human capital is an important determinant of the absorptive capacity of an economy with regards to international technology and knowledge (see amongst others Nelson, Denison, Sato, and Phelps, 1966; Coe, Helpman, and Hoffmaister, 2009). Second, and more directly, the fraction of highly educated people in the economy is a key determinant of the supply of scientists and researchers, and therefore a central factor in the R&D production function. As to product market regulation, it would be our basic position to expect a U-shaped relationship with R&D investment, in line with the arguments raised by Aghion, Bloom, Griffith, and Howitt (2005)

[2]These countries are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Italy, Netherlands, Norway, Spain, Sweden, UK and US. The selection of countries has been driven by data availability.

[3]See Section 2.1 for more details.

that we discussed in section 2.2. We measure $OPEN_{it}$ as the sum of imports and exports of goods and services as a percentage of GDP. As a proxy for the stock of human capital, we use the percentage of the population aged 15 and over that has completed tertiary schooling. To capture PMR_{it}, the OECD economy-wide product market regulation index is employed.

As a final determinant of business R&D investment, we introduce an indicator for wage pressure. Its construction is discussed in Section 3.1.2.

3.1.2 An appropriate wage indicator

To assess the impact of wage formation and wage pressure on business R&D investment, we follow Blanchard (2006) and use insights from growth theory. The approach is to compare actual (growth of) real wage costs with the so-called 'warranted' real wage (growth). The latter is determined by the rate of Harrod-neutral technical progress. In growth theory, this is the rate of real wage growth consistent with stable employment along a balanced growth path. Blanchard (2006) constructs the rate of Harrod-neutral technical progress using the Solow residual, and dividing it by the labour share. More formally, let W_{it} represent real hourly labour cost in country i at time t and let A_{it} be a measure of labour efficiency driven by technological progress. The underlying CRS production function is

$$Y_{it} = K_{it}^{\alpha} G_{it}^{\beta} (A_{it} L_{it})^{(1-\alpha-\beta)}, \tag{2}$$

with Y_{it} real output, K_{it} the stock of real private physical capital, G_{it} the stock of real public capital, L_{it} total hours worked, and $A_{it}L_{it}$ effective labour in hours. Labour efficiency can then be computed as:

$$\ln A_{it} = \frac{1}{1-\alpha-\beta} \left[\ln Y_{it} - \alpha \ln K_{it} - \beta \ln G_{it} - (1-\alpha-\beta) \ln L_{it}\right] \tag{3}$$

Following Blanchard's reasoning, a suitable wage gap or wage pressure indicator will then be defined as real hourly labour cost per efficiency unit of labour, $\frac{W_{it}}{A_{it}}$. In our empirical analysis, we will express this indicator in logs, such that we get

$$\ln WAGE_{it} = \ln \frac{W_{it}}{A_{it}} = \ln W_{it} - \ln A_{it} \tag{4}$$

As to data, W_{it} represents real compensation of employees per hour. To compute $\ln A_{it}$, we estimate the production function in (2) for the same panel of countries that we study in our empirical analysis of private R&D investment. In line with, amongst others, Costantini and Destefanis (2009), Eberhardt and Teal (2013) and Everaert, Heylen, and Schoonackers (2015), we account for the presence of unobserved common factors that are potentially non-stationary. Estimation of this production function yields a share of private capital in total income (α) of 0.20, a share of public capital (β) of 0.14, and a labour share ($1-\alpha-\beta$) of 0.66. Our estimate for β is very close to the results reported by Bom and Ligthart (2014). Building on a meta-regression analysis, they put forward 0.11 as long-run output elasticity of public capital. Using the Blanchard indicator has the additional advantage that it is not (directly) affected by endogenous adjustment of labour productivity, as is the case for more traditional indicators that measure the wage gap by relating real labour cost to labour productivity i.e. output per hour or per worker. Such indicators will give the wrong sign when firms adjust capital intensity in response to wage changes. For example, excessive wage increases may induce firms to substitute capital for labour. The productivity of labour will then rise and excessive wage pressure may no longer show up in the data, implying measurement error.

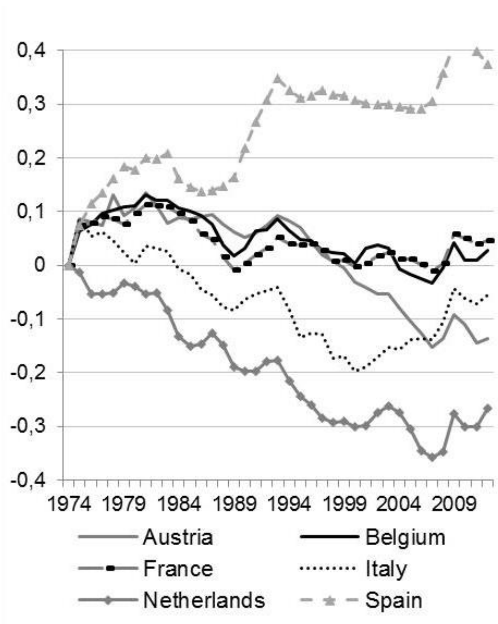
Figure 2 shows our indicator for wage pressure ($\ln WAGE_{it}$) in three groups of countries: six euro area countries, four Nordic countries and four Anglo-Saxon countries. Note that for each country, we normalized the wage gap to zero in 1974. Although this is obviously somewhat arbitrary, the idea is that in the early 1970s about all countries were close to full employment, so that wages must have been more or less at their 'warranted' level[4]. All in all, our indicator is very similar to the real wage gap of Arpaia and Pichelmann (2007), which is also based on the Blanchard approach.

Wage pressure increased strongly in most countries throughout the second half of the 1970s, with a peak around 1982. From then onwards, the trend in the wage gap was negative in most countries. Many countries, such as Belgium, Italy and Sweden, institutionalized mechanisms of wage moderation or wage control to bring (and keep) the evolution of wages more in line with their warranted level. Other countries, like the UK, decentralized wage bargaining, and introduced tough legislation to reduce union power. Only in the early 1990s and in the first years after the recent financial crisis, we observe a temporary resurgence of wage pressure. The main exceptions to this overall pattern are the US, the Netherlands, Canada and Spain. The evolution of wages was exceptional in the US in that we see no excess wage growth in the 1970s. Moreover, since 1980, wage growth in the US has only been slightly smaller than its warranted level, keeping the wage gap between 0 and -8 % all of the time. The Netherlands, by contrast, shows a steady decline of wage pressure throughout almost the entire period under consideration. This confirms the strong focus on wage moderation as an important policy instrument in this country. Very influential in this respect was the so-called *Wassenaar agreement* of 1982, which initiated a series of national social contracts to restrain wage growth. Unions were convinced of the need to restrain inflationary pressure

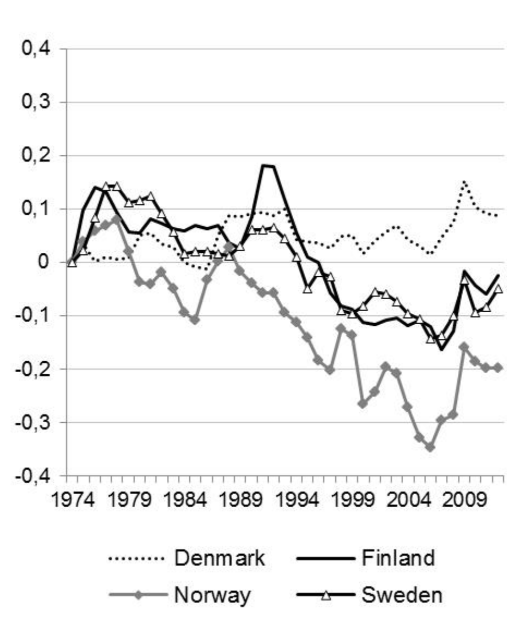
[4]Even if this assumption were wrong for some countries, it will not affect our estimation results in Section 4, since we control for unobserved country fixed effects. What matters is the evolution of $\ln WAGE$ over time, not its initial level.

Figure 2: Indicator of wage pressure ($\ln WAGE_{it}$) for three groups of countries

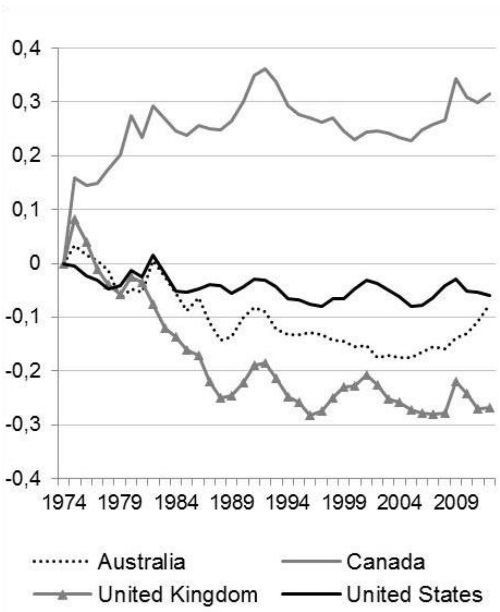
(a) Euro area countries



(b) Nordic countries



(c) Anglo-Saxon countries



in the labour market and co-ordinated action was introduced to bring this about. Canada and Spain differ in the sense that we see no wage moderation in these countries during the last three decades.

In our regressions in Section 4 we will at first introduce $\ln WAGE$ as a separate variable. Building on our discussion in Section 2.2, however, we will soon add interaction terms with context variables that may tilt the effect of wage pressure on R&D investment. The degree of openness ($OPEN$) and the degree of product market regulation (PMR), already discussed in Section 3.1.1, affect the strength of the competition that firms experience. The degree of employment protection legislation (EPL) determines the difficulty that firms may face to adapt by changing (expensive) labour. All three context variables $OPEN$, PMR and EPL affect the impact of wage pressure on the difference between the profits that firms may expect to earn when they innovate and when they do not innovate.

3.1.3 Properties of the data

As a guide to selecting the most appropriate estimation method in Section 3.3 and to determine the optimal functional form for equation (1), we first look at two important properties of the data: the degree of cross-sectional dependence and the order of integration.

Cross-sectional dependence

Recently, the panel data literature has seen an increasing interest in models with unobserved, time-varying heterogeneity that may stem from omitted (and unobserved) common variables or global shocks that affect all units, but perhaps to a different degree (see e.g. Coakley, Fuertes, and Smith, 2002; Eberhardt and Teal, 2011; Everaert and Pozzi, 2014). These omitted common variables induce error cross-sectional dependence and may lead to inconsistent estimates if they are correlated with the explanatory variables and to a spurious regression problem if they are non-stationary.

At the macroeconomic level, cross-sectional dependencies are rather the rule than the exception because countries are interconnected through trade, geography, international relations etc. (Westerlund, 2008). When considering the potential determinants of business R&D intensity across OECD countries, unobserved common variables are also likely to be present. A first potential common factor is a global business cycle, which results from the increased business cycle synchronization across countries. Changes in this global business cycle affect the financial constraints of both the government and the business sector and will thus have an impact on business R&D intensity (Guellec and Van Pottelsberghe, 2003). A second, and probably more important unobserved common factor is the world level of technology and knowledge. A key characteristic of new technology is that it may spill over to other firms and countries. Eberhardt, Helmers, and Strauss (2013) have shown that these spillovers affect firms' private returns to R&D. Depending on the extent to which firms and countries enjoy these spillovers, the world level of technology will be an important (but unobserved) factor driving business R&D expenditures.

If these unobserved common factors have indeed an impact on business R&D, this should show up as strong cross-sectional dependence in the data. Table 1 therefore reports the average pairwise correlation coefficient ($\bar{\hat{\rho}}$) and the cross-sectional dependence (CD) test of Pesaran (2004). As all series are potentially non-stationary, we also report results for the first-differenced data to avoid spurious nonzero correlation. To assess if common factors are really influencing business R&D, especially the cross-sectional dependence in $\ln BERD_{it}$ is important. For completeness, we also report the test results for each of the explanatory variables.

The results in Table 1 show that all variables except one exhibit considerable positive cross-sectional correlation in levels and in first differences. $\ln SUBS_{it}$ is the exception as the null hypothesis of no cross-sectional dependence is not rejected for the variable in levels, but is rejected for the data in first differences. The finding of significant cross-sectional dependence in $\ln BERD_{it}$ implies that we need to take this into account when choosing our econometric methodology and estimating our empirical model.

Table 1: Cross-sectional dependence in the data

Sample period: 1981-2012, 14 OECD countries

	Levels			First-differences				Levels			First-differences		
	$\bar{\hat{\rho}}$	CD		$\bar{\hat{\rho}}$	CD			$\bar{\hat{\rho}}$	CD		$\bar{\hat{\rho}}$	CD	
$\ln BERD_{it}$	0.881	47.55	[0.00]	0.194	10.277	[0.00]	$\ln BINDEX_{it}$	0.190	10.255	[0.00]	0.037	1.965	[0.05]
$\ln VA_{it}$	0.926	49.955	[0.00]	0.575	30.544	[0.00]	$OPEN_{it}$	0.701	37.830	[0.00]	0.669	35.507	[0.00]
$\ln GOVERD_{it}$	0.051	2.745	[0.01]	0.054	2.87	[0.01]	$HCAP_{it}$	0.930	50.185	[0.00]	0.05	2.656	[0.01]
$\ln HERD_{it}$	0.961	51.868	[0.00]	0.089	4.771	[0.00]	$\ln WAGE_{it}$	0.415	2.379	[0.00]	0.447	23.745	[0.00]
$\ln SUBS_{it}$	0.027	1.468	[0.14]	0.043	2.262	[0.02]	PMR_{it}	0.958	51.738	[0.00]	0.191	10.147	[0.00]

Notes: The average cross-correlation coefficient $\bar{\hat{\rho}} = (2\,/N\,(N-1))\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}$ is the average of the country-by-country cross-correlation coefficients $\hat{\rho}_{ij}$ (for $i \neq j$). CD is the Pesaran (2004) test defined as $\sqrt{2T\,/N\,(N-1)}\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}$, which is asymptotically standard normal under the null of cross-sectional independence. p-values are reported in square brackets.

Time series properties

An analysis of the time series properties of each variable in our empirical model requires a panel unit root test allowing for cross-sectional dependence. Such panel unit root tests have been proposed by, most notably, Pesaran (2007), Moon and Perron (2004) and Bai and Ng (2004). These tests are similar in that they all assume an observed variable x_{it} to have the following common factor structure

$$x_{it} = d_{it} + f_t\pi_i + \xi_{it}, \tag{5}$$

where f_t is an $r \times 1$ vector of r common factors with country-specific factor loadings π_i, ξ_{it} is an idiosyncratic error term and d_{it} is a deterministic component which can be (i) zero, $d_{it} = 0$, (ii) an idiosyncratic intercept, $d_{it} = d_{0i}$, or (iii) an idiosyncratic intercept and idiosyncratic linear trend $d_{it} = d_{0i} + d_{1i}t$. Cross-sectional dependence stems from the component $f_t\pi_i$ which is correlated over countries as it includes the common factors f_t. The series x_{it} is non-stationary if at least one of the common factors in f_t is non-stationary, or the idiosyncratic error ξ_{it} is non-stationary, or both. The above mentioned panel unit root tests differ in the allowed number and order of integration of the unobserved common factors and in the way these factors are eliminated.

The most general panel unit root test allowing for cross-sectional dependence is the PANIC unit root test of Bai and Ng (2004) as this is the only one that allows for non-stationarity in either the common factors, or in the idiosyncratic errors, or in both. Rather than testing the order of integration using the observed data, x_{it} is first decomposed according to the structure in equation (5). By applying the method of principal components to the first-differenced data, the common and idiosyncratic components in first-differences can be estimated consistently, irrespectively of their orders of integration. Next, these components are accumulated to obtain the corresponding level estimates $\widehat{f}_t^{pc}$ and $\widehat{\xi}_{it}^{pc}$. These components can then be tested separately for unit roots. When there is only one factor, testing for a unit root in $\widehat{f}_t^{pc}$ can be done using a standard augmented Dickey-Fuller (ADF)-type test (with deterministic terms according to the specification of d_{it}). For multiple common factors, the $MQ_c^{c,\tau}$ and $MQ_f^{c,\tau}$ statistics (see Bai and Ng, 2004, for details) are designed to determine the number of independent stochastic trends $r_1 \leq r$ in $\widehat{f}_t^{pc}$. As under the appropriate choice for the number of common factors, $\widehat{\xi}_{it}^{pc}$ by design satisfies the cross-sectional independence assumption required for pooling, the Maddala and Wu (1999) (MW) panel unit root test can be used on $\widehat{\xi}_{it}^{pc}$. This test consists of combining p-values for the ADF tests (with no deterministic terms) on the idiosyncratic error $\widehat{\xi}_{it}^{pc}$. The relevant distributions for the ADF tests on $\widehat{f}_t^{pc}$ and $\widehat{\xi}_{it}^{pc}$, for the intercept only and the linear trend model, can be found in Bai and Ng (2004).

Monte Carlo simulation results in Bai and Ng (2004), for samples as small as (T=100, N=40), and in Gutierrez (2006), for samples as small as (T=50, N=20), show that the PANIC approach performs well in small samples. The ADF test on the common factor and the MW test on the idiosyncratic error terms both have an actual size close to the 5% nominal level and adequate power. Applications of the PANIC approach to unit root testing using a similar data span as ours (T=32, N=14) can be found in, among others, Byrne, Fiess, and Ronald (2011), Costantini, Demetriades, James, and Lee (2013) and Everaert, Heylen, and Schoonackers (2015).

Table 2: PANIC unit root tests

Sample period: 1981-2012, 14 OECD countries

		$\widehat{f}_t^{pc}$		$\widehat{\xi}_{it}^{pc}$			$\widehat{f}_t^{pc}$		$\widehat{\xi}_{it}^{pc}$
	Det	r	r_1	MW-test		Det	r	r_1	MW-test
$\ln BERD_{it}$	d_{it}	1	1	37.907 [0.10]	$\ln BINDEX_{0i}$	d_{it}	0	0	17.45 [0.93]
$\ln VA_{it}$	d_{it}	3	3	24.854 [0.64]	$OPEN_{it}$	d_{it}	2	2	12.364 [1.00]
$\ln GOVERD_{it}$	d_{0i}	1	1	12.056 [1.00]	$HCAP_{it}$	d_{it}	5	5	42.238 [0.04]
$\ln HERD_{it}$	d_{it}	0	0	27.91 [0.47]	$\ln WAGE_{it}$	d_{0i}	3	3	21.597 [0.80]
$\ln SUBS_{it}$	d_{it}	1	1	27.868 [0.47]	PMR_{it}	d_{it}	3	3	24.572 [0.65]

Notes: 'Det' indicates the deterministic component of the model, i.e. d_{0i} for the intercept only model and $d_{it} = d_{0i} + d_{1i}t$ for the linear trend model. The number of common factors is estimated using the BIC_3 of Bai and Ng (2002) with a maximum of 5 factors. When $r = 1$, the number of non-stationary factors r_1 is determined using the ADF-GLS test of Elliott, Rothenberg, and Stock (1996) with deterministic terms according to the specification of d_{it}. When $r > 1$, r_1 is determined using the MQ_c^c (intercept only model) or MQ_c^τ (linear trend model) statistic of Bai and Ng (2004). The panel unit root test on the estimated idiosyncratic errors is the Maddala and Wu (1999) (MW) test (with no deterministic terms). The null hypothesis for each of these tests is that the series has a unit root. p-values are reported in square brackets.

In Table 2 we report the results of the PANIC unit root tests. For each of the variables the number of common factors r is estimated using the BIC_3 information criterion suggested by Bai and Ng (2002). Their simulation results, as well as those of Moon and Perron (2007), show that the BIC_3 outperforms other information criteria in small samples like ours. The specification of the deterministic component d_{it} is chosen from the observed trending behaviour of the variables. Results show that all variables are found to be non-stationary at the 5 % level of significance. For all but two variables, the non-stationarity is induced by both the common component and idisoyncratic errors. For the variable $HCAP_{it}$ non-stationarity only stems from the presence of a set of unobserved common factors while for $\ln HERD_{it}$ non-stationarity comes from the idiosyncratic component as this variable is found to have no common factor according to the BIC_3 information criterion. When focusing on the main variable of interest, $\ln BERD_{it}$, the Bai and Ng (2002) test to determine the number of common factors shows the presence of 1 non-stationary common factor.

3.2 Empirical model

Our empirical analysis shares the macro focus of existing research by Guellec and Van Pottelsberghe (2003), Falk (2006) and Westmore (2014). While these authors study both the long-run and the short-run relationship between privately-funded business R&D and its drivers, our focus is on the long-run cointegrating relationship only. An important contribution of this paper, however, is that we fully take into account (and deal with) the two key properties of the data that we described in the previous section, i.e. the significant degree of cross-sectional dependence due to the presence of unobserved common factors and the non-stationarity of the variables considered. We consider as our basic specification the following long-run relationship for $\ln BERD_{it}$,

$$\ln BERD_{it} = \gamma_i + X_{it}\beta + \mu_{it}. \tag{6}$$

where $X_{it} = (\ln VA_{it}, \ln POLICY_{it}, Z_{it}, \ln WAGE_{it})$ and $\beta' = (\beta_1, \beta_2, \beta_3, \beta_4)$. In this equation, the individual effect γ_i captures unobserved time-invariant heterogeneity.

To deal with cross-sectionally correlated errors (see Section 3.1.3) we adopt a multi-factor error structure, where cross-sectional dependence is modelled to arise from unobserved common factors (see e.g. Eberhardt and Teal, 2011; Everaert, Heylen, and Schoonackers, 2015):

$$\mu_{it} = \lambda_i' f_t + \epsilon_{it}, \tag{7}$$

where f_t is an rx1 vector of unobserved common factors and λ_i an rx1 country-specific vector of factor loadings. The generality of the error structure in (7) is an advantage as it allows for an unknown (but fixed) number of unobserved common components with heterogeneous factor loadings (heterogeneous cross-sectional dependence). It thus also nests common time effects (homogeneous cross-sectional dependence) as a special case and controls for possible spatial spillovers (Pesaran and Tosetti, 2011). This last element could be important as in a recent paper Montmartin and Herrera (2015) point to the importance of spatial dependence between private R&D activities in OECD countries.

In the empirical analysis we will focus on determining the long-run drivers of business sector R&D by estimating equation (6). Note that when estimating this equation it is important to deal appropriately with the multi-factor error structure in (7) as ignoring the presence of unobserved common factors leads to inconsistent estimates if the unobserved factors are correlated with the explanatory variables and to a spurious regression problem if they are non-stationary. Finally, as all variables have a unit root we test for the existence of a cointegration relationship between the variables in (6).

3.3 Econometric methodology

In line with Pesaran (2006) and Kapetanios, Pesaran, and Yamagata (2011), the set of unobserved common factors f_t is identified from the cross-sectional dimension of the data. Taking cross-sectional averages of the model represented by equations (6)-(7) yields

$$\overline{y}_t = \overline{\gamma} + \overline{\lambda} f_t + \overline{X}_t \beta + \overline{\epsilon}_t, \tag{8}$$

where $y_{it} = \ln BERD_{it}$ and where $\overline{y}_t = \frac{1}{N}\sum_{i=1}^{N} y_{it}$ and similarly for $\overline{\gamma}$, $\overline{\lambda}$, $\overline{X}_t$ and $\overline{\epsilon}_t$. For notational convenience we assume a single common factor ($r = 1$) but the results straightforwardly generalize to multiple factors (see Pesaran, 2006). Equation (8) can then be solved for f_t as

$$f_t = \frac{1}{\overline{\lambda}}\left(\overline{y}_t - \overline{\gamma} - \overline{X}_t \beta - \overline{\epsilon}_t\right), \tag{9}$$

which yields $\widehat{f}_t^{ca}$

$$\widehat{f}_t^{ca} = \frac{1}{\overline{\lambda}}\left(\overline{y}_t - \overline{\gamma} - \overline{X}_t \beta\right), \tag{10}$$

as a proxy for f_t. Under the assumption that ϵ_{it} is a zero mean stationary error term which is uncorrelated over cross-section units, implying that $\operatorname*{plim}_{N\to\infty} \overline{\epsilon}_t = 0$ for each t, we have that $\widehat{f}_t^{ca} \xrightarrow{p} f_t$ for $N \to \infty$. This is the main result in Pesaran (2006) that the cross-sectional averages of the observed data can be used as observable proxies for f_t. Although the construction of $\widehat{f}_t^{ca}$ as a consistent estimator for f_t in equation (10) requires knowledge of the unknown underlying parameters, Pesaran (2006) shows that these parameters can be estimated from an augmented model obtained by replacing the unobserved f_t in equation (7) by the cross-sectional averages of the observed data using equation (9)

$$\begin{aligned} y_{it} &= \gamma_i + \left(\overline{y}_t - \overline{\gamma} - \overline{X}_t \beta - \overline{\epsilon}_t\right)\frac{\lambda_i}{\overline{\lambda}} + X_{it}\beta + \varepsilon_{it}, && (11)\\ &= \gamma_i^{+} + \overline{y}_t \lambda_{i1} + \overline{X}_t \lambda_{i2} + X_{it}\beta + \epsilon_{it}^{+}, && (12) \end{aligned}$$

where $\gamma_i^{+} = \gamma_i - \overline{\gamma}\lambda_i \big/ \overline{\lambda}$, $\lambda_{i1} = \lambda_i \big/ \overline{\lambda}$, $\lambda_{i2} = -\lambda_i \big/ \overline{\lambda}\,\beta$ and $\epsilon_{it}^{+} = \epsilon_{it} - \lambda_i \big/ \overline{\lambda}\,\overline{\epsilon}_t$. Since $\epsilon_{it}^{+} \xrightarrow{p} \epsilon_{it}$ for $N \to \infty$, the augmented model in equation (12) - ignoring any parameter restrictions - can be estimated with least squares (LS), an approach referred to as the CCEP estimator.[5] Pesaran (2006) shows that, under appropriate regularity conditions, the CCEP estimator is consistent and asymptotically normal in stationary panel regressions. Kapetanios, Pesaran,

[5] Although equation (12) is derived, for notational convenience, under the assumption of a single factor, exactly the same augmented form is obtained for multiple common factors (see Pesaran, 2006).

and Yamagata (2011) show that these asymptotic results continu to hold in non-stationary panels provided that the idiosyncratic error term ϵ_{it} is stationary. This requires that there is cointegration (i) between (y_{it}, X_{it}) if $f_t \sim I(0)$ or (ii) between (y_{it}, X_{it}, f_t) if $f_t \sim I(1)$.

As our empirical analysis involves testing for cointegration, we need an appropriate panel cointegration test based on the CCEP estimator. These kind of tests have been suggested by Banerjee and Carrion-i-Silvestre (2011) and Everaert (2014). Banerjee and Carrion-i-Silvestre (2011) show that under the null of no cointegration, the linear CCEP estimator allows for consistent estimation of the homogeneous coefficients β but not for the heterogeneous coefficients (γ_i, λ_i). Given this result, they suggest to obtain a consistent estimate for the composite error term $e_{it} = \gamma_i + \lambda_i f_t + \epsilon_{it}$ as

$$\widehat{e}_{it} = y_{it} - X_{it}\widehat{\beta} = (\gamma_i + \widehat{\lambda_i f_t} + \epsilon_{it}), \tag{13}$$

and test for cointegration using a panel unit root test on $\widehat{e}_{it}$ that takes into account the cross-sectional dependence induced by the set of unobserved common factors f_t. To this end, they suggest to use the cross-section augmented ADF (CADF) panel unit root test of Pesaran (2007). Although this approach can effectively sweep out a single common factor, f_t is restricted to have the same order of integration as the idiosyncratic error term ϵ_{it}. This rules out that $f_t \sim I(1)$ and $\epsilon_{it} \sim I(0)$, i.e. cointegration between (y_{it}, x_{it}, f_t). Since the structure of the composite error term $e_{it} = \gamma_i + \lambda_i f_t + \epsilon_{it}$ aligns with the general factor structure of equation (5), an obvious alternative to the CADF test is to apply the PANIC approach of Bai and Ng (2004).[6] This allows to consistently decompose $\widehat{e}_{it}$ in a set of common factors, denoted $\widehat{f}_t^{pc}$, and an idiosyncratic error term, labeled $\widehat{\epsilon}_{it}^{pc}$, which can then be separately tested for unit roots (see PANIC approach outlined in Section 3.1.3). The main advantage of this approach is that the test whether the idiosyncratic errors ϵ_{it} are stationary or not does not depend on the order of integration of f_t. As such, testing for cointegration from the CCEP estimation results boils down to testing whether there is a unit root in $\widehat{\epsilon}_{it}^{pc}$, for which the MW panel unit root test can be used. Note that although cointegration only requires the idiosyncratic errors to be $I(0)$, the integration properties of the common factors provide additional interesting information, i.e. when $f_t \sim I(0)$ there is cointegration between (y_{it}, X_{it}) while for $f_t \sim I(1)$ there is cointegration between (y_{it}, X_{it}, f_t). In a simulation exercise both Everaert (2014) and Everaert, Heylen, and Schoonackers (2015) show that a PANIC on the composite error term

[6] Using the PANIC approach to testing for panel cointegration in the presence of common factors has also been suggested by Gengenbach, Palm, and Urbain (2006), Banerjee and Carrion-i-Silvestre (2006) and Bai and Carrion-i-Silvestre (2013). The main difference between these approaches and ours lies in the estimation of the unknown coefficients in the cointegrating relation, for which we use the CCEP estimator while the above references estimate a model in first-differences with the common factors and factor loadings estimated using principal components.

$\widehat{e}_{it}$ is an appropriate approach to test for common-factor augmented panel cointegration, even in small samples as ours.

4 Estimation results

4.1 Main results

The main estimation results are reported in Table 4. As mentioned before, our dependent variable is the log of real per capita R&D investment financed and performed by the business sector ($\ln BERD_{it}$). We estimate 10 different specifications. We start in column (1) by considering the standard set of variables that Guellec and Van Pottelsberghe (2003) introduce in their regressions. Next to value added in the business sector ($\ln VA_{it}$), there are four policy variables: public funding of R&D projects in the business sector ($\ln SUBS_{it}$), the B-index reflecting a country's tax treatment of R&D investment ($\ln BINDEX_{it}$), direct 'intramural' government expenditures on R&D ($\ln GOVERD_{it}$) and expenditures on R&D by higher education institutions ($\ln HERD_{it}$). In columns (2)-(4) we respectively extend the set of explanatory variables by the degree of openness ($OPEN_{it}$), the stock of high-skilled human capital ($HCAP_{it}$) and our wage pressure indicator ($\ln WAGE_{it}$). Column (5) further controls for a non-linear impact of the amount of public subsidies whereas columns (6)-(10) test for non-linear and/or heterogeneous effects of wage pressure.

In a first step each specification is tested for the existence of a cointegration relationship using the PANIC approach of Bai and Ng (2004), which requires determining the number of unobserved common factors in $\ln BERD_{it}$. The analysis in Table 2 points to the existence of 1 common factor in $\ln BERD_{it}$. As an additional check, Table 3 reports the cross-sectional correlation in $\ln BERD_{it}$ and in the CCEP composite error term $\widehat{e}_{it}$ after taking out the contribution of $r = (0, 1, 2, 3)$ common factors. For $r = 0$, this is the cross-sectional correlation in the original series, while for $r > 0$ this is the cross-sectional correlation in the idiosyncratic part calculated using PANIC with $r = (1, 2, 3)$. The results confirm the presence of one common factor as this seems sufficient to remove the cross-sectional dependence from $\ln BERD_{it}$ and the CCEP composite error term.

FE results

To highlight the importance of dealing with cross-sectional dependence for the estimation results, we first ignore any unobserved common factors and estimate the empirical model using a standard FE estimator. The results can be found in Appendix A. Using the FE estimator, we cannot reject the null of no cointegration in any specification. The PANIC cointegration

Table 3: Determining the number of relevant common factors

Sample period: 1981-2012, 14 OECD countries

	Cross-sectional correlation left after taking out r factors								
	$r=0$	$r=1$	$r=2$	$r=3$		$r=0$	$r=1$	$r=2$	$r=3$
$\ln BERD_{it}$	0.881	-0.053	-0.063	-0.055	$\Delta \ln BERD_{it}$	0.1935	-0.048	-0.06	-0.059
$\widehat{e}_{it}^{S}1$	0.549	-0.015	-0.017	-0.0384	$\Delta\widehat{e}_{it}^{S}1$	0.086	-0.02	-0.013	-0.041
$\widehat{e}_{it}^{S}2$	0.487	-0.028	-0.041	-0.024	$\Delta\widehat{e}_{it}^{S}2$	0.096	-0.032	-0.034	-0.039
$\widehat{e}_{it}^{S}3$	0.194	-0.044	-0.044	-0.063	$\Delta\widehat{e}_{it}^{S}3$	0.112	-0.038	-0.032	-0.056
$\widehat{e}_{it}^{S}4$	0.574	-0.026	-0.040	-0.043	$\Delta\widehat{e}_{it}^{S}4$	0.092	-0.025	-0.028	-0.039
$\widehat{e}_{it}^{S}5$	0.131	-0.058	-0.062	-0.066	$\Delta\widehat{e}_{it}^{S}5$	0.147	-0.045	-0.048	-0.058
$\widehat{e}_{it}^{S}6$	0.086	-0.047	-0.039	-0.056	$\Delta\widehat{e}_{it}^{S}6$	0.128	-0.047	-0.038	-0.049
$\widehat{e}_{it}^{S}7$	0.089	-0.042	-0.044	-0.052	$\Delta\widehat{e}_{it}^{S}7$	0.116	-0.039	-0.04	-0.054
$\widehat{e}_{it}^{S}8$	0.127	-0.052	-0.063	-0.053	$\Delta\widehat{e}_{it}^{S}8$	0.154	-0.047	-0.055	-0.06
$\widehat{e}_{it}^{S}9$	0.817	-0.047	-0.047	-0.052	$\Delta\widehat{e}_{it}^{S}9$	0.117	-0.040	-0.036	-0.054
$\widehat{e}_{it}^{S}10$	0.315	-0.051	-0.043	-0.054	$\Delta\widehat{e}_{it}^{S}10$	0.103	-0.038	-0.027	-0.049

Note: $\widehat{e}_{it}^{S}1$, $\widehat{e}_{it}^{S}2$,..., $\widehat{e}_{it}^{S}8$ are the CCEP composite error terms, defined in equation (13) taken from specification (1),(2),...,(10) respectively. We report the average cross-correlation $\overline{\widehat{\rho}}$ (see Table 1 for the definition) after taking out r common factors using PANIC.

test at the bottom of Table 7 shows that both the common factor and the idiosyncratic error term are non-stationary at the 5% level of significance. This is problematic as Urbain and Westerlund (2011) show that the standard result in Phillips and Moon (1999) that panel regressions yield consistent results even if there is no cointegration, does no longer hold when the non-stationarity in the error term is induced by a common factor. This implies that the results from the FE estimator, which ignores the presence of non-stationary common factors, are spurious. As such we do not interpret these results.

CCEP results

When we use the CCEP estimator and so control for unobserved common factors, we can reject the null hypothesis of no cointegration at the 10% level or better for all specifications containing our wage indicator. Table 4 reports the results. We obtain the best test results, i.e. rejection of the null of no cointegration at the 1% level, in specifications (6) and (7). These specifications do not only include the wage gap, but also allow its effect on business R&D investment to depend on the institutional context as reflected by $OPEN_{it}$ or EPL_{it}. In this sense, our results are supportive to the overall hypothesis on non-uniform wage effects that we formulated in Section 3. Looking in detail at the PANIC cointegration test results, the time series properties of the unobserved common factor f_t reveal that this variable is part of the cointegration relationship. So there is cointegration between (y_{it}, X_{it}, f_t).

Regarding the estimated coefficients, the effect of total value added on R&D investment in the business sector is robustly positive and statistically significant in all our regressions. The estimated (partial) long-run elasticity varies between 0.43 and 0.75, the median being 0.65. As to public policies, our estimation results reveal various ways in which governments can effectively promote R&D investment in a country. One approach is to give tax incentives or subsidies and grants to private investors. Another is to spend more on R&D within the public sector itself if this does not crowd out private spending. Our evidence suggests that both options can work, if chosen appropriately.

Let us start with the former. In a majority of our regressions, we observe a negative and statistically significant effect on the B-index of about -0.18, supporting the hypothesis that tax incentives encourage business R&D investment. This result is clearly in line with most of the literature that we summarized in Section 2.1. In some of our regressions, though, the observed negative effect is not statistically significant. For public funding of investment in the business sector ($\ln SUBS_{it}$) we always obtain positive but mostly highly insignificant elasticities. Only in specifications (9) and (10) the long-run elasticity varies around 0.045 and is significant at the 5% level. At first sight, our results therefore seems to indicate that private firms are not encouraged to raise their own R&D expenditures and undertake additional investments when some of their projects are publicly funded. Neither, however, do they cut back on their own spending. The observed positive coefficient on $\ln SUBS_{it}$ clearly challenges the hypothesis that subsidized private firms would just substitute public money for their own. Additional analysis, however, as in specification (5), reveals a much richer reality behind this general result. When we follow Guellec and Van Pottelsberghe (2003) and allow for different effects from public funding on business R&D expenditures depending on the level of the subsidization percentage, we find both at low subsidization rates (i.e. below 4%) and at high subsidization rates (above 11%) a negative elasticity of public funding[7]. At intermediate subsidization rates, however, we find this elasticity to be positive (0.076) and statistically significant. We conclude that direct government funding can be effective in promoting business R&D investment, but this funding should not be too low, neither too high. In the former case support may be too weak to help firms overcome the risks and uncertainties involved in innovation projects. In the latter case, support may be larger than

[7] We follow Guellec and Van Pottelsberghe (2003) and use the share of government funded R&D in total business performed R&D as a proxy for the subsidization rate. We find this rate to be low (< 4 % on average over the sample period) in Australia and Finland, and high (> 11%) in France, Italy, Norway, Spain, UK and US. The other countries take intermediate positions.

the number of (new) projects that firms can develop, so that in the end they simply use public resources to finance projects that would have been done anyway. In this sense we confirm earlier evidence by Guellec and Van Pottelsberghe (2003)[8].

Results on the effects of R&D spending within the government sector ($\ln GOVERD_{it}$) and within institutions of higher education ($\ln HERD_{it}$) all go in the same direction. The effect is positive in almost all cases but small and mostly insignificant. Although this may sound poor from a statistical perspective, it is not unimportant economically. It means that each euro that the government spends on 'intramural' R&D or on R&D within universities adds one euro to aggregate spending on R&D. Our findings therefore go against the hypothesis of (weak) crowding out, for which Guellec and Van Pottelsberghe (2003) found evidence, as well as against the hypothesis of complementarity between public and private spending, as suggested by Falk (2006). Only for $\ln GOVERD_{it}$ in columns (6) and (7) we may find some weak indications in favour of complementarity. The regressions in these columns yield a long-run positive elasticity of about 0.1[9].

Important is also that governments can stimulate private R&D investment by encouraging high-skilled human capital formation. This is confirmed by our empirical results which yield very robust and significant positive estimates on the stock of high-skilled human capital ($HCAP_{it}$). Considering the lack of consistent findings in the existing literature (see for example Falk, 2006), this is an interesting result. We also find a positive effect from the degree of openness of the economy on business R&D spending. In column (2) this positive effect is not statistically significant. In specification (6) it is. This may point to the importance of international transfer of technology and knowledge for business R&D. However, in line with results that we discuss below for the wage gap, a complementary interpretation could be that a more open economy raises the degree of competition that firms face. Facing more competitors then seems to encourage firms to innovate.

An important potential determinant of business R&D investment is wage pressure. Theory being inconclusive, what do we learn from our results on its impact on innovation? When analyzing the basic effect in specifications (4) and (5) we do not find any significant effect from wage formation on business R&D investment. In column (4) the effect is insignificantly negative whereas in column (5) it is insignificantly positive. However, a much more detailed

[8]Obviously, considering our use of macro data, the non-linear relationship that we find is also a macro feature. Low or high macro rates of subsidies may hide large firm-level heterogeneity in the subsidies.

[9]The observed elasticities allow us to compute the marginal effect on business financed R&D and on aggregate R&D spending (business + public) induced by one euro spent by the government. Considering that $BERD_{it}$ relates to $GOVERD_{it}$ as 5 to 1 and to $SUBS_{it}$ as 13 to 1 on average over all countries considered in our analysis, elasticities of 0.126 for $GOVERD_{it}$ and 0.076 for $SUBS_{it}$ (the highest we observe in our results) would imply that $\frac{\Delta BERD_{it}}{\Delta GOVERD_{it}} = 0.63$ and $\frac{\Delta BERD_{it}}{\Delta SUBS_{it}} = 0.99$. In the case of $GOVERD_{it}$, aggregate R&D spending would thus rise by 1.63 euro (1 euro public + 0.63 euro private), in the case of $SUBS_{it}$ by 1.99 euro (1 euro public + 0.99 euro private).

analysis, based on the theoretical arguments in Section 2.2, gives a clearer view. As wage pressure has possibly positive effects in a very competitive environment, we allow in specification (6) for interaction between $\ln WAGE_{it}$ and $OPEN_{it}$. The basic impact of $\ln WAGE_{it}$ is now negative and highly significant, with an estimated long-run coefficient equal to -1.1. If wage pressure increases with 1 %point, this implies that, on average, business R&D investment drops with 1.11 %. Higher wage pressure thus seems to undermine business R&D expenditures. An obvious explanation, and in line with Ulph and Ulph (1994), would be that higher wages reduce the profit differential between innovating and not innovating. However, the basic hypothesis only seems to survive in economies where firms face relatively little (foreign) competition. In a competitive environment the wage effect may be tilted. From the interaction term in specification (6) we learn that in countries with a degree of openness higher than 80 % the global impact of wage pressure becomes positive, meaning that higher wages encourage private R&D investment. Specification (7) differentiates the effect of $\ln WAGE_{it}$ according to the level of employment protection legislation (EPL_{it})[10]. In countries with low average EPL a significant negative effect of wage pressure emerges, again indicating that higher wages reduces the incentive to innovate. In (very) regulated labour markets the negative impact disappears as for the other two groups of countries we observe (insignificant) positive effects of wage pressure on business R&D investment. As an additional check, the possible impact of openness and labour market characteristics are integrated in specification (8). We distinguish three groups of countries. The first group of Anglo-Saxon countries is characterized by a relatively low degree of openness and low employment protection legislation. The estimated effect from $\ln WAGE_{it}$ is clearly negative in this group (although significant only at 20%). The second group of euro area countries is characterized by rather the opposite of a high degree of openness and rigid labour markets. Here we observe a significant positive coefficient on $\ln WAGE_{it}$. The arguments raised by Kleinknecht (1998) and co-authors that an excessive focus on wage moderation could be harmful to innovation, would thus seem to find support for this group. The third group of Nordic countries takes an intermediate position.

[10] As time variation in EPL_{it} is too limited, we cannot interact $\ln WAGE_{it}$ with EPL_{it}. As a solution, we differentiate the impact of $\ln WAGE_{it}$ amongst three groups of countries with different average EPL_{it}

Table 4: CCEP regression results

Dependent variable: $\ln BERD_{it}$

Sample period: 1981-2012, 14 OECD countries

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Coefficient estimates										
Explanatory variables										
$\ln VA_{it}$	0.644***	0.709***	0.646***	0.459**	0.663***	0.659***	0.497**	0.742***	0.428*	0.512**
	(0.18)	(0.194)	(0.187)	(0.242)	(0.227)	(0.257)	(0.249)	(0.222)	(0.230)	(0.240)
$\ln BINDEX_{it}$	-0.186^{**}	-0.114	-0.187^{**}	-0.182^{**}	-0.170^{**}	-0.103	-0.10	-0.189^{**}	-0.229^{***}	-0.142^{*}
	(0.087)	(0.10)	(0.089)	(0.086)	(0.084)	(0.093)	(0.092)	(0.080)	(0.086)	(0.082)
$\ln SUBS_{it}$	0.004	0.01	0.017	0.000		0.028	0.031	0.034	0.046**	0.044**
	(0.002)	(0.025)	(0.024)	(0.023)		(0.024)	(0.024)	(0.022)	(0.023)	(0.022)
$\ln GOVERD_{it}$	0.055	0.053	0.055	0.064	0.059	0.087*	0.126**	0.021	0.023	0.020
	(0.044)	(0.045)	(0.045)	(0.044)	(0.043)	(0.046)	(0.05)	(0.042)	(0.053)	(0.049)
$\ln HERD_{it}$	0.071	0.063	-0.017	0.096	-0.057	0.042	0.032	0.074	-0.018	0.022
	(0.069)	(0.072)	(0.071)	(0.068)	(0.064)	(0.071)	(0.071)	(0.066)	(0.065)	(0.067)
$HCAP_{it}$			0.051***		0.093***	0.06***	0.068***	0.092***	0.047***	0.067***
			(0.010)		(0.011)	(0.011)	(0.01)	(0.011)	(0.012)	(0.010)
$OPEN_{it}$		0.002				0.014***				
		(0.001)				(0.005)				
$\ln WAGE_{it}$				-0.205	0.127	-1.108^{**}			-0.085	0.383
				(0.211)	(0.197)	(0.463)			(0.202)	(0.445)
PMR_{it}									0.419**	-0.112
									(0.196)	(0.154)
PMR^2_{it}									-0.046	
									(0.036)	
$\ln SUBS_{it} * low$					-0.053					
					(0.075)					
$\ln SUBS_{it} * medium$					0.076**					
					(0.034)					
$\ln SUBS_{it} * high$					-0.087^{**}					
					(0.040)					
$\ln WAGE_{it} * OPEN_{it}$						0.014**				
						(0.006)				
$\ln WAGE_{it} * PMR_{it}$										-0.295^{*}
										(0.172)
$\ln WAGE_{it} * epl_{low}$							-1.205^{***}			
							(0.447)			
$\ln WAGE_{it} * epl_{middle}$							0.134			
							(0.474)			
$\ln WAGE_{it} * epl_{high}$							0.092			
							(0.257)			
$\ln WAGE_{it} * anglo$								-0.639		
								(0.464)		
$\ln WAGE_{it} * euro$								1.331***		
								(0.338)		
$\ln WAGE_{it} * nordic$								0.27		
								(0.305)		
Panic Cointegration test (one common factor)										
ADF-GLS on $\widehat{f}^{pc}_t$	-0.699	-1.392	-0.931	-0.797	-1.455	-2.105	-1.55	-2.168	-0.073	-0.344
	[0.97]	[0.84]	[0.94]	[0.96]	[0.82]	[0.52]	[0.79]	[0.49]	[0.99]	[0.99]
MW on $\widehat{\epsilon}^{pc}_{it}$	1.23	1.37*	1.086	1.603**	1.295*	2.446***	2.918***	1.661**	11.460*	2.085**
	[0.11]	[0.09]	[0.14]	[0.05]	[0.10]	[0.01]	[0.00]	[0.05]	[0.07]	[0.02]

Notes: standard errors are in parentheses. *, ** and *** indicate significance at the 10%, 5% and 1% level respectively. For the panel cointegration test results, the unit root test on the common factor $\widehat{F}_t$ is a ADF-GLS test for a model with constant. The corresponding (simulated) p-values are reported in square brackets. The unit root test on the estimated idiosyncratic errors $\widehat{\epsilon}^{pc}_{it}$ is a MW test. The corresponding p-values are reported in square brackets

Finally, in specifications (9) and (10) we analyse the direct impact of product market (de)regulation, PMR_{it}, and its possible effect on the relation between wage pressure and innovation. Following Aghion, Bloom, Griffith, and Howitt (2005), a U-shaped relationship should be expected between PMR_{it} and $\ln BERD_{it}$. In specification (9) we do not find evidence for this U-shaped effect. On the contrary, results show that more regulated product markets increase firms' incentive to invest in R&D. In this view, higher product market regulation, and thus lower competition, increases firms' rents when investing in R&D. When we also take into account the possible impact of PMR_{it} on the effect of wage pressure on business R&D in specification (10), the direct impact of PMR_{it} is somewhat different. Now, higher product market regulation has a negative but insignificant impact on the amount of business R&D expenditures. More interesting is the interaction between PMR_{it} and $\ln WAGE_{it}$. This interaction effect confirms our earlier finding. In a less competitive environment (high PMR_{it}), higher wage pressure reduces the incentives of firms to invest in R&D. When product markets become more deregulated, the basic negative effect disappears and in these circumstances wage pressure can even stimulate private R&D investment.

4.2 The importance of economic and policy related variables in explaining private investment in R&D

Our empirical results in Table 4 help us to understand and explain important differences in the level and evolution of real business R&D in the OECD during the last decades. In what follows, we discuss the explanatory power of our estimated empirical model and conduct a counterfactual analysis. The latter allows us to assess the contribution of changes since 1981 in public policy, wage pressure and human capital to the evolution of business R&D. What fraction of the total change in $BERD$ between 1981 and 2012 can these explanatory variables explain? Which was more important, which was less important? Are the results the same for all countries/country groups?

Explanatory power

Figure 3 demonstrates the capacity of our empirical model to explain the variation in business R&D investment across countries and over time. We use the regression result in specification (6). The upper panel in Figure 3 (panel a) relates our model's prediction (economic explanation) for the *level* of business R&D expenditures in 2006-2007 to the true observation[11]. Both prediction and true observation are represented as log deviations from their overall country

[11] We choose these years as they are the last before the outbreak of the global financial crisis. Severe shocks to firms' investment decisions during this crisis imply that the data are much less likely to match the long-run equilibrium relationship that our model captures.

averages. The lower panel (panel b) relates predicted and observed *changes* in business R&D between 1981 and 2007. We emphasize that our predictions in both panels have been obtained solely from using the 'economic' and 'policy related' parts of the estimated equation. They do not include the country-specific fixed effects nor the approximations for the country-specific factor loadings and unobserved common factors.

Correlation in panel (a) is 0.50. Our model correctly predicts far above average business R&D investment in 2006-2007 in the US and (far) below average R&D investment in Italy and France. The model's prediction of close to average performance in Australia, Austria, Belgium, Denmark and Norway is also quite well in line with the facts. On the other hand, using only the economic and policy related explanatory variables in the model, it is harder to match the high level of business R&D investment in 2006-2007 in Finland and Sweden. So it is to match relatively low investment in Spain. It is clear that for these countries the unobserved common factor was more important than for other countries. In this respect, our results are in line with those of Everaert, Heylen, and Schoonackers (2015). Studying the drivers of TFP, they find for Finland and Sweden a relatively strong and rising absorptive capacity to the (unobserved and common) world level of technology. Stronger international technology spillovers, and their effects on the private return to R&D, may explain an important part of the above average business investment in innovation in these countries. The opposite may explain weaker investment in Spain. Clearly, the observation that the common factor plays an important role, at least for some countries, is fully in line with our earlier finding that this factor belongs to the cointegration relationship. It supports (again) our choice for the CCEP estimator.

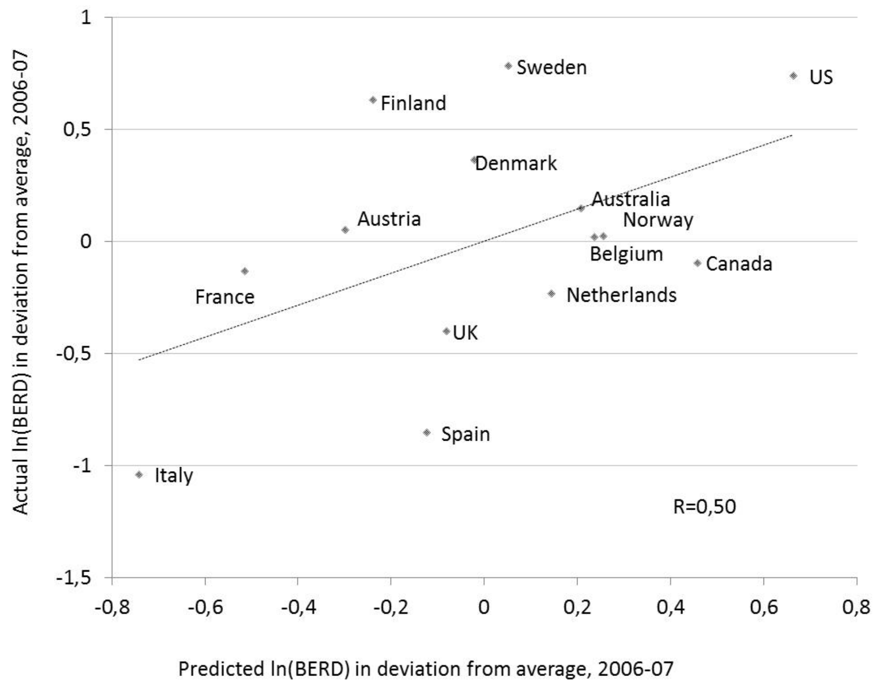
Correlation in panel (b) is 0.34. The model again seems to have the main drivers of business R&D investment right for the US, France and Italy. It also explains quite well the change in business R&D investment over time in countries like Austria and Norway. Finland is again by far the largest outlier. On the basis of (changes in) economic and policy related variables it is impossible to explain the strong actual rise in $BERD$ since 1981 in this country. Dropping Finland, correlation in panel (b) rises to 0.50.

Counterfactual analysis

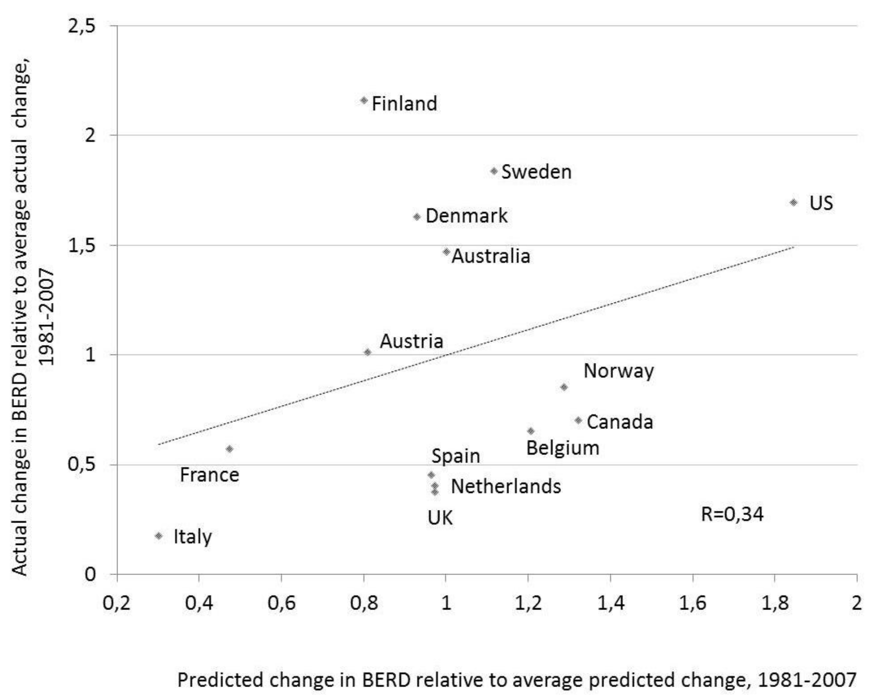
Figure 4 reveals the estimated size of the estimated effects on business R&D expenditures of changes in public innovation policy, wage pressure, human capital and real value added in the business sector since 1981 in the US, an average of five EU countries and the Nordic countries in our sample. Each graph compares the model's fitted value for these countries with (i) the simulated value if all policy variables (spending, taxes) had remained at their 1981 level, (ii)

Figure 3: Actual and predicted business R&D expenditure (Table 4, specification 6)

(a) Actual and predicted business R&D expenditure levels, 2006-2007



(b) Actual and predicted change in business R&D expenditure, 2007 versus 1981



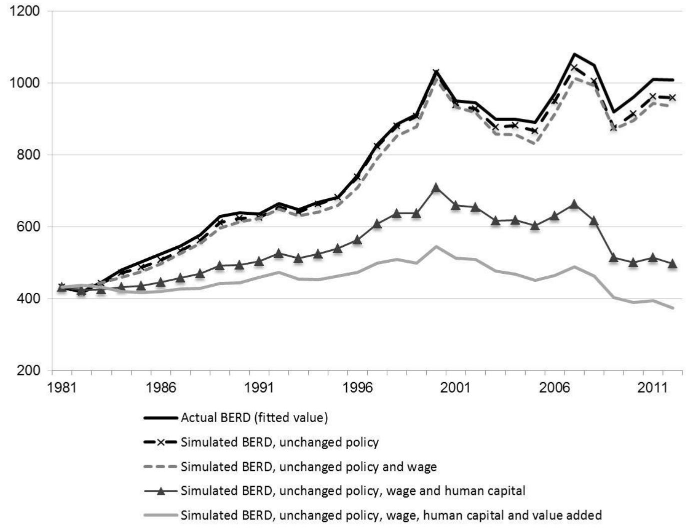
the simulated value if all policy variables and wage pressure had remained at their 1981 level, (iii) the simulated value if all policy variables, wage pressure and human capital had remained at their 1981 level, and (iv) the simulated value if all policy variables, wage pressure, human capital and real value added had remained at their 1981 level. Individual country graphs are available upon request.

All three graphs underscore the importance of public policy, wage pressure, human capital and real value added. As indicated by the lower lines in Figure 4a-c, business R&D investment in the US and the core EU5 would have remained fairly flat between 1981 and 2012 if these variables had remained unchanged. The Nordic countries would then have seen an increase in $BERD$ of about 75%. In this sense, Figure 4 is fully consistent with our earlier findings in Figure 3. The observed policy and other variables were very important for the evolution of business R&D investment in most countries. It seems again, though, that in comparative perspective the Nordic countries also benefited strongly from the evolution of the unobserved common factor. In the US and the core EU its impact was minimal.

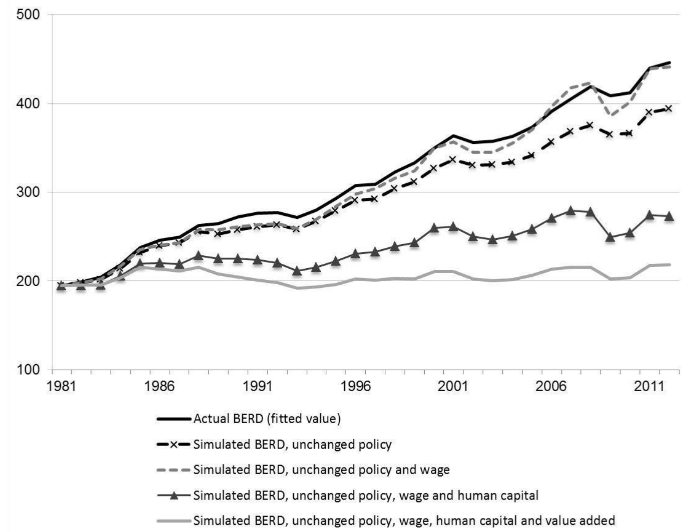
As to the relative contribution of the set of policy variables, the wage gap, value added and high-skilled human capital for the evolution of $BERD$, Figure 4 leaves no doubt that the latter was the most important. In the US the increase of human capital contributed to more than 75% of the total increase in $BERD$. In the core EU that was almost 70%, in the Nordic countries about 40%. Value added comes second in line as a driver of business R&D investment in each of the three country groups. Public innovation policies come third. In the core EU changes in policy accounted for a little less than 20% of the observed increase in $BERD$ since 1981. In the Nordic countries the contribution of policy changes was about 13%, in the US about 9%. Finally, changes in the wage gap may have contributed the least to change in business R&D investment. This conclusion holds in particular for the US and the Nordic countries where fairly limited changes in wage pressure, on average, had rather neutral effects. Changes in the wage gap did matter, however, in the core EU5. The focus on wage moderation in countries like the Netherlands, Austria and Italy in particular had an important negative impact on $BERD$ over time. Figure 4 shows that, for the evolution of $BERD$, the stimulating effect of public policy in the core EU5 countries was entirely neutralized by the negative effects of wage moderation.

Figure 4: Counterfactual analysis: fitted and simulated model (Table 4, specification 6)

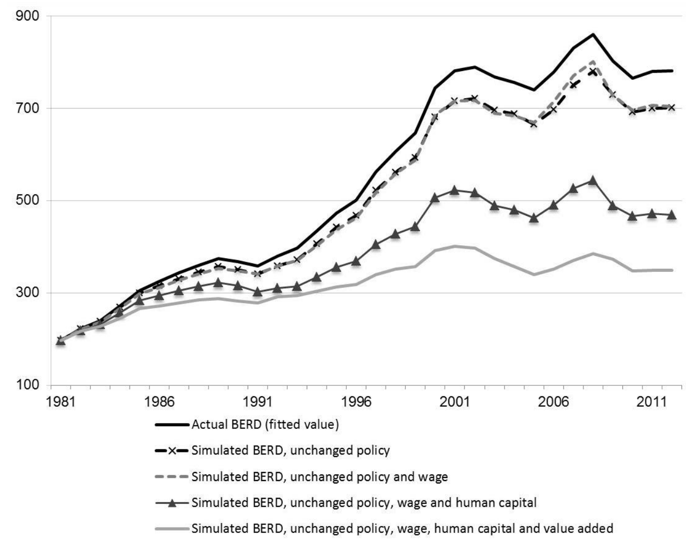
(a) United States



(b) Core-EU5 (Austria, Belgium, France, Italy, Netherlands)



(c) Nordic-4



4.3 Robustness test: alternative specification of the wage indicator

To construct our wage indicator, $\ln WAGE$, in Section 3.1.2, we used data on the share of private capital (α), public capital (β) and labour ($1 - \alpha - \beta$) in total income. Estimation of a basic production function for the set of countries in our empirical analysis gave us the required information. However, to show that our results are not sensitive to the exact choice of the income shares, we also constructed an alternative wage indicator based on different (but evenly realistic) output elasticities. More specifically, we set α and β equal to respectively 0.30 and 0.10, which is in line with the results reported by Bom and Ligthart (2014), and the resulting labour share to 0.6. Using these elasticities, we recalculated $\ln WAGE$ and re-estimated all related specifications with the CCEP estimator. Results can be found in Appendix B. For all specifications, results are similar to the ones in Table 4 and the same conclusions apply. As an additional robustness check we took into account the empirical observation that bargained wages tend to be lower the higher the unemployment rate (see e.g. Blanchflower and Oswald, 1994; Nickell, Nunziata, Ochel, and Quintini, 2003). This may somewhat bias our wage gap indicator. As a second robustness test, we controlled for this by following Arpaia and Pichelmann (2007) and added the unemployment rate to our wage indicator. Again, results are very similar to the ones reported in Table 4[12].

4.4 Direction of causation

The empirical results in Table 4 give proof of a long-run relationship between $\ln BERD_{it}$ and its determinants. To provide evidence that the long-run coefficients in Table 4 can be interpreted as empirical causal effects, we apply a test for the direction of caution based on the approach of Eberhardt and Teal (2013) which builds on the discussion in Canning and Pedroni (2008). From the Granger Representation Theorem (Engle and Granger, 1987) we know that if there exists a cointegration relationship between the variables in the model, these series can be represented in the form of a dynamic error correction model.

Equations (14)-(15) formalize this relationship for our empirical model. For notational convenience, this is done under the assumption of 1 ($r = 1$) common factor, f_t, and of 1 variable included in X_{it}.

[12]Results using this alternative wage indicator are not reported here, but are available upon request.

$$\Delta \ln BERD_{it} = \kappa_{1i} + \alpha_1 \widehat{\varepsilon}_{i,t-1} + \rho_{1i}\Delta f_t + \sum_{j=1}^{J} \phi_{11j}\Delta \ln BERD_{i,t-j} + \sum_{j=1}^{J} \phi_{12j}\Delta X_{i,t-j} + \nu_{1it}, \tag{14}$$

$$\Delta X_{it} = \kappa_{2i} + \alpha_2 \widehat{\varepsilon}_{i,t-1} + \rho_{2i}\Delta f_t + \sum_{j=1}^{J} \phi_{21j}\Delta \ln BERD_{i,t-j} + \sum_{j=1}^{J} \phi_{22j}\Delta X_{i,t-j} + \nu_{2it}. \tag{15}$$

where $\widehat{\varepsilon}_{i,t-1}$ represents the 'disequilibrium term'. The cointegration test results from Table 4 show that there is cointegration between (y_{it}, X_{it}, f_t). This implies that the 'disequilibrium term' is constructed as $\widehat{\varepsilon}_{it} = \ln BERD_{it} - \gamma_i - X_{it}\beta - \lambda_i f_t$ and that f_t is included in equations (14)-(15). As a proxy for f_t and for $\widehat{\varepsilon}_{it}$, we use the results of the PANIC cointegration testing procedure which provides us respectively with $\widehat{f}_t^{pc}$ and $\widehat{\epsilon}_{it}^{pc}$. Equations (14)-(15) further include lagged differences of the observable variables in the cointegrating relationship.

For a long-run relationship to exist between $\ln BERD_{it}$, X_{it} and f_t, α_1 or α_2 must be nonzero . If $\alpha_1 \neq 0$ then X_{it} has a causal impact on $\ln BERD_{it}$; if $\alpha_2 \neq 0$ then $\ln BERD_{it}$ has a causal impact on X_{it}. If both α_1 and α_2 are non-zero, X_{it} and $\ln BERD_{it}$ determine each other jointly. In the above example there are only two equations, as we have two variables in the cointegration relationship. In our empirical analysis we will have $k+1$ equations, with k being the number of variables in X_{it}. Empirical estimates for α_1, α_2,...,α_{k+1} are investigated using standard t-ratios, given that all variables in the ECM regression are stationary[13].

Results are presented in Table 5 for two lags ($J = 2$), but the same conclusions can be drawns for one lag ($J = 1$). Due to the limited time series dimension of our data, we do not consider extra lags. In Table 5, the first row of each specification refers to the estimation of $\widehat{\alpha}_1$, while for all other rows, $\widehat{\alpha}_2$ is estimated with the dependent variable the variable mentioned in the column 'Variable'. Table 5 shows that for each specification that we have estimated, X_{it} has an impact on $\ln BERD_{it}$. This can be seen from the estimation of equation (14) in row 1 for each specification in Table 5. To be sure that the estimated impact is causal, equation (15) is estimated for each element in X_{it} as a dependent variable. If the error correction term of these equations is zero, then the corresponding x-variable has a causal impact on $\ln BERD_{it}$. The results in Table 5 show that all results can be interpreted as empirical causal effects as

[13]The disequilibrium term $\widehat{\varepsilon}_{it}$ constructed from specifications (1), (2), (3) and (5) is not stationary at the 5% significance level but still stationary at the 10% level for specifications (2) and (5). This implies that for the 'direction of causation' test based on specification (1) and (3) we should employ simulated critical values. However, in our analysis we still use standard t-ratios with the reason being that the p-values of stationarity of the disequilibrium term are very close to 10 % and that we are mainly interested in the specifications that include our wage measure.

no error correction term is significant when estimating equation (15).

As an additional check we allow the short term coefficients (ϕ_{11j}, ϕ_{12j}, ϕ_{21j} and ϕ_{22j}) and the error correction terms (α_1 and α_2) in equations (14) and (15) to vary across countries. Results for the 'direction of causation' test when allowing for this heterogeneity can be found in Table 6, where the mean group results are reported. When allowing for short-term heterogeneity across countries, conclusions on the direction of causation are somewhat different. First, there is clear evidence that $\ln HERD_{it}$ and $\ln BERD_{it}$ determine each other jointly. This implies that the coefficient on $\ln HERD_{it}$ in our specifications should be interpreted as a correlation and not as a causal effect. Evidence of reverse causality is also present for $\ln VA_{it}$ as is shown by the test results of specification (3), (5),(7) and (8). For PMR_{it}, the test results of specification (10) indicate a possible problem of reverse causality, although only at the 10 % significance level. Finally, it is also important to note that only in specification (7) there is some indication that the coefficients on $\ln WAGE_{it}$ could not be interpreted as causal. However, this is only the case at the 10% level of significance. Moreover, all other specifications show that the estimated effect of $\ln WAGE_{it}$ is causal.

Table 5: Test for direction of causation

Variable	$\widehat{\alpha}$	std	Verdict	Variable	$\widehat{\alpha}$	std	Verdict	Variable	$\widehat{\alpha}$	std	Verdict
Specification (1)				**Specification**(2)				**Specification** (3)			
$\ln BERD_{it}$	-0.182^{***}	(0.041)	$x \rightarrow y$	$\ln BERD_{it}$	-0.196^{***}	(0.049)	$x \rightarrow y$	$\ln BERD_{it}$	-0.229^{***}	(0.053)	$x \rightarrow y$
$\ln VA_{it}$	0.001	(0.023)		$\ln VA_{it}$	-0.003	(0.025)		$\ln VA_{it}$	-0.011	(0.028)	
$\ln BINDEX_{it}$	0.036	(0.037)		$\ln BINDEX_{it}$	0.042	(0.041)		$\ln BINDEX_{it}$	0.34	(0.045)	
$\ln SUBS_{it}$	-0.179	(0.125)		$\ln SUBS_{it}$	-0.169	(0.137)		$\ln SUBS_{it}$	-0.194	(0.152)	
$\ln GOVERD_{it}$	0.052	(0.062)		$\ln GOVERD_{it}$	0.042^{*}	(0.068)		$\ln GOVERD_{it}$	0.060	(0.076)	
$\ln HERD_{it}$	0.050^{*}	(0.237)		$\ln HERD_{it}$	0.039	(0.047)		$\ln HERD_{it}$	0.054	(0.052)	
				$OPEN_{it}$	0.010	(0.039)		$HCAP_{it}$	0.097	(0.167)	
Specification (4)				**Specification** (5)				**Specification** (6)			
$\ln BERD_{it}$	-0.188^{***}	(0.048)	$x \rightarrow y$	$\ln BERD_{it}$	-0.171^{**}	(0.078)	$x \rightarrow y$	$\ln BERD_{it}$	-0.237^{***}	(0.095)	$x \rightarrow y$
$\ln VA_{it}$	0.001	(0.025)		$\ln VA_{it}$	0.001	(0.039)		$\ln VA_{it}$	-0.011	(0.042)	
$\ln BINDEX_{it}$	0.034	(0.041)		$\ln BINDEX_{it}$	0.017	(0.064)		$\ln BINDEX_{it}$	0.027	(0.069)	
$\ln SUBS_{it}$	-0.182	(0.137)		$\ln GOVERD_{it}$	0.048	(0.109)		$\ln SUBS_{it}$	-0.116	(0.239)	
$\ln GOVERD_{it}$	0.053	(0.069)		$\ln HERD_{it}$	0.040	(0.074)		$\ln GOVERD_{it}$	0.081	(0.118)	
$\ln HERD_{it}$	0.046	(0.046)		$HCAP_{it}$	0.254	(0.231)		$\ln HERD_{it}$	0.031	(0.082)	
$\ln WAGE_{it}$	-0.002	(0.023)		$\ln Wage_{it}$	0.013	(0.037)		$HCAP_{it}$	0.202	(0.262)	
				$\ln SUBS_{it} * low$	-0.020	(0.083)		$OPEN_{it}$	0.032	(0.068)	
				$\ln SUBS_{it} * medium$	-0.102	(0.166)		$\ln WAGE_{it}$	0.009	(0.039)	
				$\ln SUBS_{it} * high$	-0.106	(0.132)		$\ln WAGE_{it} * OPEN_{it}$	-0.026	(0.078)	
Specification (7)				**Specification** (8)				**Specification** (9)			
$\ln BERD_{it}$	-0.227^{***}	(0.081)	$x \rightarrow y$	$\ln BERD_{it}$	-0.176^{**}	(0.076)	$x \rightarrow y$	$\ln BERD_{it}$	-0.263^{***}	(0.091)	$x \rightarrow y$
$\ln VA_{it}$	-0.021	(0.042)		$\ln VA_{it}$	-0.001	(0.038)		$\ln VA_{it}$	-0.020	(0.045)	
$\ln BINDEX_{it}$	0.020	(0.067)		$\ln BINDEX_{it}$	0.016	(0.062)		$\ln BINDEX_{it}$	0.034	(0.078)	
$\ln SUBS_{it}$	-0.131	(0.231)		$\ln SUBS_{it}$	-0.049	(0.216)		$\ln SUBS_{it}$	-0.096	(0.264)	
$\ln GOVERD_{it}$	0.078	(0.115)		$\ln GOVERD_{it}$	0.064	(0.106)		$\ln GOVERD_{it}$	0.089	(0.131)	
$\ln HERD_{it}$	0.052	(0.078)		$\ln HERD_{it}$	0.066	(0.072)		$\ln HERD_{it}$	0.062	(0.088)	
$HCAP_{it}$	0.080	(0.253)		$HCAP_{it}$	0.185	(0.286)		$HCAP_{it}$	0.134	(0.286)	
$\ln WAGE_{it} * epl_{low}$	-0.004	(0.016)		$\ln WAGE_{it} * anglo$	-0.009	(0.015)		$\ln WAGE_{it}$	0.022	(0.043)	
$\ln WAGE_{it} * epl_{middle}$	0.019	(0.023)		$\ln WAGE_{it} * euro$	0.028	(0.019)		PMR_{it}	0.212	(0.143)	
$\ln WAGE_{it} * epl_{high}$	0.009	(0.029)		$\ln WAGE_{it} * nordic$	-0.002	(0.028)		PMR_{it}^2	1.214	(0.757)	
Specification (10)											
$\ln BERD_{it}$	-0.252^{***}	(0.002)	$x \rightarrow y$								
$\ln VA_{it}$	-0.006	(0.046)									
$\ln BINDEX_{it}$	0.022	(0.077)									
$\ln SUBS_{it}$	-0.049	(0.265)									
$\ln GOVERD_{it}$	0.102	(0.131)									
$\ln HERD_{it}$	0.047	(0.089)									
$HCAP_{it}$	0.233	(0.287)									
$\ln WAGE_{it}$	0.009	(0.043)									
PMR_{it}	0.223	(0.147)									
$\ln WAGE_{it} * PMR_{it}$	-0.147	(0.145)									

Notes: The first row for each specification presents the results for $\widehat{\alpha}_1$ (see eq. (14)). For all other rows 'Variable' refers to the ECM or dynamic regression with 'Variable' on the left hand side. In the first column we report the $\widehat{\alpha}_2$ coefficient as is described in eq. (15). The second and third column respectively denote the standard deviation of $\widehat{\alpha}$ and the final conclusion. For the first row, '$x \rightarrow y$' means that X_{it} has an impact on $\ln BERD_{it}$. For the rows where the column 'Verdict' is empty, this implies that the corresponding x-variable has a causal impact on $\ln BERD_{it}$. The rows where the 'Verdict' column indicates '$\longleftrightarrow$' point to reverse causality between the corresponding x-variable and $\ln BERD_{it}$ and results for this explanatory variable should thus be interpreted carefully. *, ** and *** indicate significance at the 10%, 5% and 1% level respectively.

Table 6: Test for direction of causation: heterogeneous short-term effects

Variable	$\widehat{\alpha}$	std	Verdict	Variable	$\widehat{\alpha}$	std	Verdict	Variable	$\widehat{\alpha}$	std	Verdict
Specification (1)				**Specification** (2)				**Specification** (3)			
$\ln BERD_{it}$	-0.180^{***}	(0.036)	$x \to y$	$\ln BERD_{it}$	-0.169^{***}	(0.045)	$x \to y$	$\ln BERD_{it}$	-0.220^{***}	(0.064)	$x \to y$
$\ln VA_{it}$	0.011	(0.024)		$\ln VA_{it}$	0.017	(0.027)		$\ln VA_{it}$	0.078^{**}	(0.029)	$\longleftrightarrow$
$\ln BINDEX_{it}$	0.055	(0.041)		$\ln BINDEX_{it}$	0.095	(0.056)		$\ln BINDEX_{it}$	0.134^{**}	(0.061)	$\longleftrightarrow$
$\ln SUBS_{it}$	-0.241^{**}	(0.111)	$\longleftrightarrow$	$\ln SUBS_{it}$	-0.327	(0.197)		$\ln SUBS_{it}$	-0.251	(0.196)	
$\ln GOVERD_{it}$	0.060	(0.080)		$\ln GOVERD_{it}$	0.057	(0.078)		$\ln GOVERD_{it}$	0.052	(0.089)	
$\ln HERD_{it}$	0.149^{*}	(0.078)	$\longleftrightarrow$	$\ln HERD_{it}$	0.143	(0.085)		$\ln HERD_{it}$	0.145	(0.091)	
				$OPEN_{it}$	0.046	(0.043)		$HCAP_{it}$	0.157	(0.236)	
Specification (4)				**Specification** (5)				**Specification** (6)			
$\ln BERD_{it}$	-0.211^{***}	(0.033)	$x \to y$	$\ln BERD_{it}$	-0.173^{*}	(0.095)	$x \to y$	$\ln BERD_{it}$	-0.230^{*}	(0.128)	$x \to y$
$\ln VA_{it}$	0.035	(0.032)		$\ln VA_{it}$	0.069^{*}	(0.038)	$\longleftrightarrow$	$\ln VA_{it}$	0.087	(0.056)	
$\ln BINDEX_{it}$	0.069	(0.046)		$\ln BINDEX_{it}$	0.070	(0.223)		$\ln BINDEX_{it}$	0.215^{**}	(0.086)	$\longleftrightarrow$
$\ln SUBS_{it}$	-0.162	(0.111)		$\ln SUBS_{it}$	-0.189	(0.109)		$\ln SUBS_{it}$	-0.460	(0.325)	
$\ln GOVERD_{it}$	0.046	(0.082)		$\ln GOVERD_{it}$	-0.065	(0.099)		$\ln GOVERD_{it}$	-0.146	(0.131)	
$\ln HERD_{it}$	0.134^{*}	(0.071)	$\longleftrightarrow$	$\ln HERD_{it}$	0.253^{**}	(0.102)	$\longleftrightarrow$	$\ln HERD_{it}$	0.283^{**}	(0.127)	$\longleftrightarrow$
$\ln WAGE_{it}$	-0.016	(0.027)		$\ln HCAP_{it}$	0.369^{*}	(0.18)	$\longleftrightarrow$	$HCAP_{it}$	0.170	(0.392)	
				$\ln WAGE_{it}$	-0.028	(0.033)		$OPEN_{it}$	0.112	(0.100)	
								$\ln WAGE_{it}$	-0.063	(0.056)	
								$\ln WAGE_{it} * OPEN_{it}$	-0.118	(0.083)	
Specification (7)				**Specification** (8)				**Specification** (9)			
$\ln BERD_{it}$	-0.220^{*}	(0.109)	$x \to y$	$\ln BERD_{it}$	-0.283^{***}	(0.096)	$x \to y$	$\ln BERD_{it}$	0.655^{***}	(0.196)	$x \to y$
$\ln VA_{it}$	0.111^{**}	(0.046)	$\longleftrightarrow$	$\ln VA_{it}$	-0.001	(0.038)	$\longleftrightarrow$	$\ln VA_{it}$	-0.067	(0.051)	
$\ln BINDEX_{it}$	0.134	(0.079)		$\ln BINDEX_{it}$	0.061^{*}	(0.034)		$\ln BINDEX_{it}$	-0.015	(0.165)	
$\ln SUBS_{it}$	-0.131	(0.231)		$\ln SUBS_{it}$	-0.195	(0.234)		$\ln SUBS_{it}$	-0.288	(0.456)	
$\ln GOVERD_{it}$	0.0854	(0.088)		$\ln GOVERD_{it}$	-0.084	(0.092)		$\ln GOVERD_{it}$	-0.194	(0.195)	
$\ln HERD_{it}$	0.314^{**}	(0.108)	$\longleftrightarrow$	$\ln HERD_{it}$	0.224^{**}	(0.084)	$\longleftrightarrow$	$\ln HERD_{it}$	0.265^{*}	(0.137)	$\longleftrightarrow$
$HCAP_{it}$	-0.094	(0.214)		$HCAP_{it}$	0.088	(0.210)		$HCAP_{it}$	0.641	(0.631)	
$\ln WAGE_{it}$	-0.096^{*}	(0.047)	$\longleftrightarrow$	$\ln WAGE_{it}$	-0.030	(0.033)		$\ln WAGE_{it}$	0.065	(0.056)	
								PMR_{it}	-0.168	(0.871)	
								PMR_{it}^2	0.401	(3.941)	
Specification (10)											
$\ln BERD_{it}$	-0.530^{***}	(0.178)	$x \to y$								
$\ln VA_{it}$	-0.031	(0.084)									
$\ln BINDEX_{it}$	0.130	(0.109)									
$\ln SUBS_{it}$	0.160	(0.534)									
$\ln GOVERD_{it}$	-0.017	(0.166)									
$\ln HERD_{it}$	0.076	(0.102)									
$HCAP_{it}$	-0.355	(0.371)									
$\ln WAGE_{it}$	-0.010	(0.070)									
PMR_{it}	0.542^{*}	(0.253)	$\longleftrightarrow$								
$\ln WAGE_{it} * PMR_{it}$	-0.512^{**}	(0.210)	$\longleftrightarrow$								

Notes: The first row for each specification presents the results for $\sum_{i=1}^{N} \widehat{\alpha}_{1i}$ (see eq. (14)). For all other rows 'Variable' refers to the ECM or dynamic regression with 'Variable' on the left hand side. In the first column we report the $\sum_{i=1}^{N} \widehat{\alpha}_2$ coefficient as is described in eq. (15). The second and third column respectively denote the standard deviation of $\widehat{\alpha}$ and the final conclusion. For the first row, '$x \to y$' means that X_{it} has an impact on $\ln BERD_{it}$. For the rows where the column 'Verdict' is empty, this implies that the corresponding x-variable has a causal impact on $\ln BERD_{it}$. The rows where the 'Verdict' column indicates '$\longleftrightarrow$' point to reverse causality between the corresponding x-variable and $\ln BERD_{it}$ and results for this explanatory variable should thus be interpreted carefully. *, ** and *** indicate significance at the 10%, 5% and 1% level respectively.

5 Conclusion

The wedge between private and social returns to the creation of knowledge and technology justifies government involvement in the area of research and development. Nevertheless, in the current environment of restoring sustainability of public finances and search for an increasing efficiency of public policy, the question arises which policy options are most effective in stimulating business R&D investment. This paper therefore analyses the effects of different policies on aggregate business funded and performed R&D investment in a panel of 14 OECD countries since 1981. Concerning traditional policy options, we find that tax incentives are effective. Public funding (subsidization) of R&D performed by firms can also be effective if subsidies are not too low nor too high. The optimal subsidization rate may be somewhere between 6 and 10%. R&D performed within the government sector and within institutions of higher education is basically neutral with respect to business R&D. We find no evidence for crowding out nor for complementarity, which implies that each euro spent on R&D within the government feeds through one-to-one in aggregate R&D. The higher education sector may, however, indirectly be of great significance. This paper revealed human capital accumulation at the tertiary level as the most important driver of business R&D investment in the OECD during the last decades.

One of the main contributions of this paper is its attention to the impact of wage formation on business R&D investment. Conflicting hypotheses have been introduced in the literature, but not yet systematically analyzed. One hypothesis is that innovation and investment in R&D benefit from low or moderate wages, since these are important for firm profitability, which is a key condition for investment. Wage restraint is also important to convince firms that rents from innovation will not be appropriated by the unions through higher wages. The opposite hypothesis is that an excessive focus on wage moderation may kill incentives to innovate. Wage moderation may for example increase the survival probability of the least innovative firms and retard the process of creative destruction. Conversely, according to this hypothesis, higher wage pressure may force firms to innovate as a key element in their competitive strategy. Our empirical analysis favours the first hypothesis in fairly closed economies and in economies with flexible labour markets. The Anglo-Saxon countries may be the closest to this type. In highly open economies and economies with rigid labour markets, however, rather the opposite holds and high wage pressure may encourage innovation. Many European countries are more likely to match this type.

Our paper may also contribute to the macro R&D literature methodologically. More than existing studies, we pay particular attention to the time series properties of the data. As most variables in our empirical model are found to be non-stationary, we estimate a cointegrating relationship. Moreover, we also take into account the presence of cross-sectionally correlated error terms, which we find to be induced by an unobserved (non-stationary) common factor that drives business R&D spending. A sensible interpretation is that this common factor reflects the worldwide level of technology and knowledge. To capture this, we adopt the CCEP estimator of Pesaran (2006). We find that the standard fixed effects estimator yields spurious results.

The policy implications of our results include a warning against excessive wage moderation in highly open economies with rigid labour markets. Using insights from long-run growth theory, an appropriate guideline is that actual real wage growth matches the rate of Harrod-neutral technical progress. Keeping real wage growth below the rate of technical progress for an extended period of time, as more than a few European countries have been doing during the last decades, may promote employment in the short run, but it can also hurt the economy's innovative capacity and productivity in the long run. The fairly poor growth of business R&D investment in a country like the Netherlands may illustrate this long-run disadvantage. Conversely, however, in our view our findings provide no argument in favour of excessive wage pressure. In rigid labour markets the loss of employment that excessive wages may cause in the short run, may persist in the longer run due to for example hysteresis effects in bad times. If promotion of business investment in R&D is the objective, our paper suggests better alternatives, in particular tax incentives, well-chosen innovation subsidies and the development of high-skilled human capital.

Appendix A Fixed effects regression results

Table 7: FE regression results

Dependent variable: $\ln BERD_{it}$
Sample period: 1981-2012, 14 OECD countries

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Coefficient estimates										
Explanatory variables										
$\ln VA_{it}$	1.069^{***}	1.014^{***}	1.080^{***}	1.114^{***}	1.103^{***}	1.030^{***}	1.058^{***}	0.963^{***}	1.120^{***}	1.013^{***}
	(0.124)	(0.129)	(0.127)	(0.164)	(0.167)	(0.161)	(0.174)	(0.172)	(0.175)	(0.166)
$\ln BINDEX_{it}$	0.264^{**}	0.275^{***}	0.256^{**}	0.260^{**}	0.129	0.189^{*}	0.252^{***}	0.291^{***}	0.150	0.127
	(0.104)	(0.104)	(0.106)	(0.105)	(0.111)	(0.102)	(0.107)	(0.105)	(0.108)	(0.105)
$\ln SUBS_{it}$	0.243^{***}	0.237^{***}	0.240^{***}	0.244^{***}		0.201^{***}	0.241^{***}	0.240^{***}	0.265^{***}	0.255^{***}
	(0.023)	(0.023)	(0.024)	(0.023)		(0.024)	(0.025)	(0.024)	(0.025)	(0.024)
$\ln GOVERD_{it}$	-0.182^{***}	-0.179^{***}	-0.179^{***}	-0.183^{***}	-0.115^{***}	-0.079^{*}	-0.163^{***}	-0.176^{***}	-0.216^{***}	-0.197^{***}
	(0.041)	(0.041)	(0.041)	(0.041)	(0.041)	(0.042)	(0.042)	(0.042)	(0.042)	(0.040)
$\ln HERD_{it}$	0.528^{***}	0.516^{***}	0.544^{***}	0.517^{***}	0.465^{***}	0.541^{***}	0.537^{***}	0.551^{***}	0.434^{***}	0.481^{***}
	(0.048)	(0.048)	(0.060)	(0.054)	(0.065)	(0.060)	(0.063)	(0.062)	(0.068)	(0.066)
$HCAP_{it}$			-0.003		-0.002	-0.004	-0.002	-0.000	-0.008	-0.012^{*}
			(0.007)		(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
$OPEN_{it}$		0.002				0.028^{***}				
		(0.001)				(0.004)				
$\ln WAGE_{it}$				0.076	0.186	-1.738^{***}			0.078	1.302^{***}
				(0.179)	(0.188)	(0.340)			(0.179)	(0.282)
PMR_{it}									0.123	-0.456^{***}
									(0.095)	(0.076)
PMR^2_{it}									-0.043^{**}	
									(0.018)	
$\ln SUBS_{it} * low$					0.453^{***}					
					(0.049)					
$\ln SUBS_{it} * medium$					0.124^{***}					
					(0.031)					
$\ln SUBS_{it} * high$					0.106^{***}					
					(0.035)					
$\ln WAGE_{it} * OPEN_{it}$						0.031^{***}				
						(0.005)				
$\ln WAGE_{it} * PMR_{it}$										-0.446^{***}
										(0.086)
$\ln WAGE_{it} * epl_{low}$							-0.689^{*}			
							(0.415)			
$\ln WAGE_{it} * epl_{middle}$							0.129			
							(0.228)			
$\ln WAGE_{it} * epl_{high}$							0.124			
							(0.124)			
$\ln WAGE_{it} * anglo$								-0.809^{**}		
								(0.410)		
$\ln WAGE_{it} * euro$								0.456^{**}		
								(0.202)		
$\ln WAGE_{it} * nordic$								-0.460^{*}		
								(0.245)		
Panic Cointegration test (one common factor)										
ADF-GLS on $\widehat{f}_t^{pc}$	-2.541	-2.635	-2.580	-2.538	-2.877	-2.909	-2.608	-2.736	-2.437	-2.589
	[0.31]	[0.27]	[0.29]	[0.31]	[0.18]	[0.17]	[0.28]	[0.23]	[0.35]	[0.29]
MW on $\widehat{e}_{it}^{pc}$	-0.619	-1.325	-0.534	-0.492	-0.489	-1.469	-0.383	0.654	0.366	1.366^{*}
	[0.73]	[0.91]	[0.70]	[0.69]	[0.69]	[0.93]	[0.65]	[0.26]	[0.35]	[0.09]

Notes: standard errors are in parentheses. *, ** and *** indicate significance at the 10%, 5% and 1% level respectively. For the panel cointegration test results, the unit root test on the common factor $\widehat{F}_t$ is a ADF-GLS test for a model with constant. The corresponding (simulated) p-values are reported in square brackets. The unit root test on the estimated idiosyncratic errors $\widehat{e}_{it}^{pc}$ is a MW test. The corresponding p-values are reported in square brackets

Appendix B CCEP results with an alternative wage indicator based on different production function elasticities

Table 8: CCEP regression results for (4), (5), (6), (7), (8), (9) and (10) for alternative calculation of TFP and the wage gap

Dependent variable: $\ln BERD_{it}$
Sample period: 1981-2012, 14 OECD countries

	(4')	(5')	(6')	(7')	(8')	(9')	(10')
Coefficient estimates							
Explanatory variables							
$\ln VA_{it}$	0.375	0.597**	0.554**	0.395	0.655***	0.314	0.357
	(0.249)	(0.238)	(0.267)	(0.258)	(0.230)	(0.240)	(0.248)
$\ln BINDEX_{it}$	−0.171**	−0.160*	-0.093	-0.085	−0.188**	−0.217***	−0.146*
	(0.086)	(0.084)	(0.093)	(0.092)	(0.022)	(0.087)	(0.081)
$\ln SUBS_{it}$	0.002		0.029	0.031	0.034	0.049**	0.050**
	(0.023)		(0.024)	(0.024)	(0.022)	(0.023)	(0.022)
$\ln GOVERD_{it}$	0.067	0.068	0.092**	0.143***	0.020	0.023	0.033
	(0.044)	(0.044)	(0.047)	(0.051)	(0.042)	(0.053)	(0.050)
$\ln HERD_{it}$	0.100	-0.046	0.047	0.039	0.077	-0.015	0.007
	(0.068)	(0.064)	(0.071)	(0.071)	(0.066)	(0.065)	(0.066)
$HCAP_{it}$		0.092***	0.058***	0.065***	0.093***	0.044***	0.062***
		(0.012)	(0.011)	(0.010)	(0.011)	(0.012)	(0.010)
$OPEN_{it}$			0.012***				
			(0.004)				
$\ln WAGE_{it}$	-0.289	0.011	−1.182***			-0.206	0.269
	(0.195)	(0.186)	(0.413)		(0.192)	(0.404)	
PMR_{it}						0.419**	-0.111
						(0.198)	(0.135)
PMR^2_{it}						-0.045	
						(0.037)	
$\ln SUBS_{it} * low$		-0.055					
		(0.076)					
$\ln SUBS_{it} * medium$		0.070**					
		(0.035)					
$\ln SUBS_{it} * high$		−0.080**					
		(0.040)					
$\ln WAGE_{it} * OPEN_{it}$			0.013**				
			(0.005)				
$\ln WAGE_{it} * PMR_{it}$							−0.308**
							(0.151)
$\ln WAGE_{it} * epl_{low}$				−1.127***			
				(0.391)			
$\ln WAGE_{it} * epl_{middle}$				0.042			
				(0.432)			
$\ln WAGE_{it} * epl_{high}$				-0.126			
				(0.239)			
$\ln WAGE_{it} * anglo$					-0.551		
					(0.413)		
$\ln WAGE_{it} * euro$					1.215***		
					(0.322)		
$\ln WAGE_{it} * nordic$					0.143		
					(0.282)		
Panic Cointegration test (one common factor)							
ADF-GLS on $\widehat{f}^{pc}_t$	-0.812	-1.498	-2.043	-1.606	-2.254	-0.091	-0.289
	[0.95]	[0.81]	[0.55]	[0.77]	[0.44]	[0.99]	[0.99]
MW on $\widehat{e}^{pc}_{it}$	1.949**	1.423*	2.524***	2.635***	1.408*	1.757**	3.125***
	[0.03]	[0.08]	[0.01]	[0.00]	[0.08]	[0.04]	[0.00]

Notes: standard errors are in parentheses. *, ** and *** indicate significance at the 10%, 5% and 1% level respectively. For the panel cointegration test results, the unit root test on the common factor $\widehat{F}_t$ is a ADF-GLS test for a model with constant. The corresponding (simulated) p-values are reported in square brackets. The unit root test on the estimated idiosyncratic errors $\widehat{e}^{pc}_{it}$ for different number of common factors r=1,2 is a MW test. The corresponding p-values are reported in square brackets

Appendix C Construction of data and data sources

Table 9: Construction of data and data sources

Name	Notation	Construction	Data Sources
Real per capita business sector funded and performed R&D	$BERD_{it}$	$\frac{BERDVOL_{it}}{POP1564_{it}}$	
Real business sector funded and performed R&D	$BERDVOL_{it}$	$\frac{BERDVAL_{it}}{DEFL_{it}}$	
Business sector funded and performed R&D, value	$BERDVAL_{it}$	$BERDMSTI_{it} * PERCIND_{it}$	
Business sector expenditure on R&D, value	$BERDMSTI_{it}$	Original data	OECD, Main Science and Technology Indicators
Percentage of business sector expenditure on R&D financed by industry	$PERCIND_{it}$	Original data	OECD, Main Science and Technology Indicators
Working-age population between 15 and 64 years	$POP1564_{it}$	Original data	OECD, Economic Outlook, No 95, May 2014
GDP deflator, market prices	$DEFL_{it}$	Original data	OECD, Economic Outlook, No 95, May 2014
Real per capita value added in the business sector	VA_{it}	$\frac{VAVOL_{it}}{POP1564_{it}}$	
Real value added in the business sector	$VAVOL_{it}$	$\frac{VAVAL_{it}}{DEFL_{it}}$	
Value added in the business sector, value	$VAVAL_{it}$	Original data	OECD, Main Science and Technology Indicators
Real per capita government funded expenditure on R&D performed in the business sector	$SUBS_{it}$	$\frac{SUBSVOL_{it}}{POP1564_{it}}$	
Real government funded expenditure on R&D in the business sector	$SUBSVOL_{it}$	$\frac{SUBSVAL_{it}}{DEFL_{it}}$	
Government funded expenditure on R&D in the business sector, value	$SUBSVAL_{it}$	$PERCGOV_{it} * BERDTOTAL_{it}$	
Business enterprise expenditure on R&D, value	$BERDTOTAL_{it}$	Original data	OECD, Main Science and Technology Indicators
Percentage of $BERDTOTAL_{it}$ financed by the government	$PERCGOV_{it}$	Original data	OECD, Main Science and Technology Indicators

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Construction of data and data sources

Name	Notation	Construction	Data Sources
Real per capita government intramural expenditure on R&D	$GOVERD_{it}$	$\frac{GOVERDVOL_{it}}{POP1564_{it}}$	
Real government intramural expenditure on R&D	$GOVERDVOL_{it}$	$\frac{GOVERDVAL_{it}}{DEFL_{it}}$	
Government intramural expenditure on R&D, value	$GOVERDVAL_{it}$	Original data	OECD, Main Science and Technology Indicators
Real per capita expenditure on R&D in the higher education sector	$HERD_{it}$	$\frac{HERDVOL_{it}}{POP1564_{it}}$	
Real expenditure on R&D in the higher education sector	$HERDDVOL_{it}$	$\frac{HERDVAL_{it}}{DEFL_{it}}$	
expenditure on R&D in the higher education sector, value	$HERDVAL_{it}$	Original data	OECD, Main Science and Technology Indicators
B-index	$BINDEX_{it}$	Original data	OECD Science and Technology industry Outlook 2014, based on Warda (2013)
Percentage of population aged 15 and over that has completed tertiary schooling	$HCAP_{it}$	Data are available for 1980, 1985,1990,...2010. Data for the intermediate years are calculated by interpolation and data is extrapolated for 2011 and 2012	Barro and Lee (2013)
Degree of openness in the economy	$OPEN_{it}$	$\frac{IMPORTS_{it}+EXPORTS_{it}}{GDPVAL_{it}} * 100$	
Imports of goods and services, value, national accounts basis	$IMPORTS_{it}$	Original data	OECD, Economic Outlook, No 95, May 2014
Exports of goods and services, value, national accounts basis	$EXPORTS_{it}$	Original data	OECD, Economic Outlook, No 95, May 2014
Gross domestic product, value, market prices	$GDPVAL_{it}$	Original data	OECD, Economic Outlok, No 95, May 2014

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Construction of data and data sources

Name	Notation	Construction	Data Sources
Wage pressure indicator	$\ln WAGE_{it}$	$\ln W_{it} - \ln A_{it}$	
Real compensation of employees per hour	W_{it}	$\frac{\frac{Wval_{it}}{DEFL_{it}}}{L_{it}}$	
Compensation of employees, value	$Wval_{it}$	Original data	OECD, Economic Outlook, No 95, May 2014
Total hours worked in the economy	L_{it}	Original data	The Conference Board, Total Economy Database, January 2014
Harrod-neutral technical progress (in logs)	$\ln A_{it}$	$\frac{1}{1-\alpha-\beta}\left[\ln Y_{it} - \alpha \ln K_{it} - \beta \ln G_{it} - (1-\alpha-\beta)\ln L_{it}\right]$	For more details, see Section 3.1.2
GDP, volume, market prices	Y_{it}	Original data	OECD, Economic Outlook, No 95, May 2014
Private, non-residential net capital stock	K_{it}	See Everaert, Heylen, and Schoonackers (2015)	Everaert, Heylen, and Schoonackers (2015)
Real government net capital stock	G_{it}	See Everaert, Heylen, and Schoonackers (2015)	Everaert, Heylen, and Schoonackers (2015)
Product market regulation index	PMR_{it}	Data are available for 1998, 2003, 2008 and 2013. Data for the intermediate years are calculated by interpolation. Data for the years before 1998 are retropolated using growth rates of PMR data for seven network industries. This data is also taken from Koske, Wanner, Bitetti, and Barbiero (2015)	Koske, Wanner, Bitetti, and Barbiero (2015)
Low average subsidization rate	*low*	a dummy variable equal to one for the countries for which on average $\frac{SUBSVAL_{it}}{BERDTOTAL_{it}} \leq 4\%$	based on Guellec and Van Pottelsberghe (2003)
Medium average subsidization rate	*medium*	a dummy variable equal to one for the countries for which on average $4\% \leq \frac{SUBSVAL_{it}}{BERDTOTAL_{it}} \leq 11\%$	based on Guellec and Van Pottelsberghe (2003)
High average subsidization rate	*high*	a dummy variable equal to one for the countries for which on average $\frac{SUBSVAL_{it}}{BERDTOTAL_{it}} > 11\%$	based on Guellec and Van Pottelsberghe (2003)

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Construction of data and data sources

Name	Notation	Construction	Data Sources
Low employment protection legislation	epl_{low}	a dummy variable equal to one for the countries for which on average $EPL_{it} \leq 1.10$.	
Medium employment protection legislation	epl_{middle}	a dummy variable equal to one for the countries for which on average $1.10 \leq EPL_{it} \leq 2.07$	
High employment protection legislation	epl_{high}	a dummy variable equal to one for the countries for which on average $EPL_{it} > 2.07$	
Employment protection legislation index	EPL_{it}	Average of the index for legislation on temporary contracts and on regular contracts. Annual time series data for 1985-2013. Data before 1985 is retropolated based on the data from Berger and Heylen (2011)	OECD, Employment Outlook, 2013
Anglo-Saxon countries	$anglo$	a dummy variable equal to one for the Anglo-Saxon countries in our sample i.e. Australia, Canada, UK and US	
Core European countries	$euro$	a dummy variable equal to one for the core European countries in our sample i.e. Austria, Belgium, France, Netherlands, Italy and Spain	
Nordic countries	$nordic$	a dummy variable equal to one for the Nordic countries in our sample i.e. Denmark, Finland, Norway and Sweden	

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