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Three essays on the determinants of labor market dynamics

PhD dissertation

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Contents

Introduction and main results

This section provides an introduction to this PhD thesis in three parts: a general motivation for the study of labor market macroeconomics, a description of the scientific relevance of the topics studied, an overview of the main results, and finally, some policy implications of these results.

The rest of the dissertation is structured as follows. Three essays on the determinants of labor market dynamics are presented as three main chapters. Each one of these essays has a research paper structure: introduction and motivations, theoretical underpinnings, empirical model and results, discussion or further exploration of the results, and concluding remarks. This introductory section also operates as a brief summary of the results obtained and discussed throughout the dissertation.

1 Motivation

The labor market recovery from the massive economic recession unleashed in 2008 has shown to be tricky. The U.S. is a good example, where the employment rate as of mid-2013 had still not reached its pre-crisis level. Also Spain, where unemployment is still scaling on sky-high levels of unemployment (26% in the third quarter of 2013[1]). But this is not the case everywhere. Some countries, like Germany, are experiencing growth in the employment rate for a few years now, with a remarkable performance during the turmoil of the Great Recession.

This dissertation brings some new insights into why we observe these highly different paths in labor market outcomes across countries. In previous decades, there was an "unemployment problem" in Europe. Analysts wondered what was different in the U.S. than in Europe to explain such persistent gap on the time-paths of the unemployment rate. Unemployment was steadily higher in Europe, in contrast to the U.S., apparently because of the institutional set up. In other words, the European welfare state and its employment protection mechanisms were keeping unemployment higher than it could be at a given level of technology and productivity. Labor market institutions became the usual suspects of Europe's unemployment problem, and then the standard knowledge became that the reduction in employment protection and benefits would provide better incentives, allowing employment to rise.

The behavior of labor markets in the aftermath of the Great Recession is putting this standard policy advice to trial. It looks like the paradigm may change. Recent work by Freeman (2013) shows how the flexible job machine that the U.S. had been for more

[1]Source: INE (www.ine.es).

than a decade is being closely examined since it is not performing as it should. The job-less recovery in the U.S. comes across as unexpected. The diagnostics get even worse when comparing with the "German miracle" or other "less flexible" and more "institution-oriented" economies and their noteworthy labor market performance in the last few years.

As stressed by Schettkat (2010), new questions are arising. Markets were supposed to be fully efficient and the private sector to be wiser and outperform the public sector, monetary policy was supposed to be neutral to the real economy, and expansive fiscal policy was supposed to be ineffective because consumers' rationale implies full transparent knowledge that public debt is followed by tax increases (i.e. what is known as the Ricardian equivalence). However, we find ourselves with growing labor market problems such as high and resistant unemployment rates.

This dissertation is motivated from the scientific urge to learn more about the dynamics behind labor market outcomes. From this perspective we take into account some of the questions in today's economic debate and bring them to the analysis. We observe different time-paths of unemployment. It is clear that among developed countries labor market performances are heterogeneous. Then, what are the forces behind unemployment that may help in explaining these differences? Should we also question some of the standard assumptions in economic analysis? What can we learn from the experience? Maybe economics should be based more on actual facts and less on beliefs. In words of Nobel laureate Robert Solow: "It ain't the things you don't know that hurt you, it's the things you know that ain't so" (Solow, 1997).

2 Relevance of this study

This PhD thesis presents empirical investigations and international comparison of labor market performances. The scenario just described calls for the study of the forces underlying unemployment dynamics. The unemployment rate is the final output of all labor market kinematics where these forces take place. In a nutshell, the main objective throughout the thesis is to learn more about how labor markets work.

To that end it is crucial to look into the evolution of employment. On one hand, this dissertation studies the forces behind wage determination, a crucial variable to labor demand since the real wage affects hiring and firing decisions. Also, sector-level employment is analyzed. We compute the sectoral elasticity of labor demand, and investigate the employment effect of openness to international trade. Sector-level analysis is critical for the study of aggregate results as stated by Young (2013). On the other hand, this thesis studies the degree of substitutability between labor and capital and how technological improvement affects the relative demand for production factors which is central for unemployment determination in the long-run.

This dissertation is structured in three essays that focus on three main variables: the real wage, capital per worker, and sector-level employment.

The wage level is important for the hiring decisions that determine employment, but also for worker participation in the labor market. This is why the wage level receives much attention: it defines the incentives to hire (labor demand) and to participate providing work effort (labor supply). Wages, therefore, are one of the basic mechanisms behind the observed level of unemployment.

Besides employment, the other main factor in aggregate production is the stock of capital, which is the result of a series of investment decisions. Firms combine both factors in the production of total output, where the state of technology and efficiency play an important role. The time-path of capital intensity, i.e. the ratio of capital per worker, is affected by the degree of substitutability between production factors (with a given production target) and the effect that technological progress may have on the proportion of factors used in aggregate production, a mechanism that directly affects the employment level.

On a step further, it is important to look into the sectoral dynamics behind these aggregate variables. The outcome of aggregate employment is the result of the sum of employment dynamics in each industry or sector. Therefore, for a better understanding of the determination of employment, one must look at sector-level mechanisms. For example, a policy maker interested in raising employment will want to know in which sectors the elasticity of labor demand is higher and where it is lower. Then, sector-targeted policy may be more efficient than general reductions in labor costs because the employment-response of wage variations is heterogeneous across economic sectors.

Thus, the three essays that constitute this PhD thesis have common denominators. In all three essays we apply the analysis to more than one country. International comparison of experiences provides insights on what should be the target of labor market policy. Also in all three chapters we take the degree of openness to trade as an important determinant of our focus variables. It is well-established knowledge in related literature that labor market dynamics are affected by the exposure to international trade and globalization. This is reflected in the fact that trade is present in the three chapters of this thesis. Finally, in all three there is a detailed study of long-run elasticities as a crucial feature of labor market modeling. In the first essay we examine the one-to-one relationship between labor productivity and the real wage, in the second essay we estimate the elasticity of substitution between labor and capital, and in the third essay we discuss the sector elasticity of labor demand (with respect to the sector real wage). These elasticities measure the sensitivity of the focus variables to changes in the exogenous variables and thus provide very useful information for policy design.

The issues studied in this thesis are highly relevant. Proof of this can be found in very recent scholarly publications. The core questions being asked are shared by other authors and they take important part in the present academic debate. Other papers use different approaches or methodological paths. But it is critical that the issues discussed in this dissertation are central to learning more about labor market outcomes.

For example, the first essay studies the wage-effects of deunionization and international trade and their consequent role in the decline of labor's share of income. This same issue is treated in a recent International Labour Organization (UN) discussion paper by Stockhammer (2013). Also, Poilly and Wesselbaum (2013) show that a reform aimed at improving labor market flexibility is not necessarily welfare-enhancing. On the other hand, McAdam and Willman (2013) study the medium-run dynamics of economic growth with emphasis on capital intensity and the factor-biased effect of technical progress, and León-Ledesma *et al.* (2013) analyze the substitutability between labor and capital, the process of productivity growth and how they associate to the modelization of aggregate production. All these are subjects brought to discussion in the second essay. The labor-saving nature of factor-biased technical change in the U.S., also discussed in the second essay, is a result also surveyed in Klump *et al.* (2012). Finally, the third essay computes the sector-level elasticity of labor demand in three countries. Young (2013) recently focused on the elasticity of factor substitution at the industry-level for the U.S. The third essay also studies the employment effect of higher exposure to trade, an issue examined by recent studies like Yanikkaya (2013) and Gozgor (2013). Finally, it argues that technological progress may have a negative effect on employment, at least in the short-run, an issue also discussed by Feldmann (2013).

The research objectives of this thesis required the utilization of essential empirical methodologies. Two paramount econometric methods have been applied, time-series and panel data, and two levels of aggregation have been considered, aggregate national series and sector-level data.

The first essay examines the effect of union density and trade on wage-setting dynamics, which is dependant on the institutional and administrative structure of each economy, and it is reasonable to focus on aggregate data. Furthermore, complete trade data is not available for several sectors. Thus, we choose to run country-level time series estimations for 8 countries, grouped in 3 categories according to their labor market structure, and compare the results.

The second essay looks into the determination of capital intensity, with the inclusion of demand-side pressures to a standard model. Capital intensity, the demand-side approximation and controls like the openness to international trade, are also better analyzed at the aggregate level. We compare two countries with different time-paths of capital intensity and obtain contrasted results.

Finally, it is important to understand that labor market outcomes are also the result of sector-level dynamics. Factor markets associated to each sector are, for example, exposed to different degrees of competition, which affects the employment sensitivity to changes in wages. Hence, a sector-level analysis of labor demand must be included for better understanding of labor market outcomes. In this case, we construct three two-dimensional dynamic panels (i = sector, t = period) for three countries, representative of the 3 aforementioned categories.

To sum up, this dissertation not only works on important research objectives, but to do so, it develops relevant tools that are crucial for applied economic analysis.

3 Overview of main results

3.1 Essay 1: "Productivity, deunionization and trade: Wage effects and labor share implications"

The first essay presents wage-setting analysis applied to 8 countries, according to the labor-market classification in Daveri and Tabellini (2000): Anglo-Saxon (U.S. and U.K.), Continental Europe (France, Italy and Spain), Nordic (Sweden and Finland), and Japan.

The results show that wage determination in recent decades has been conditioned by three structural drivers, irrespective of the differences between these economic models. That is, the results are robust to different institutional structures, e.g., if the labor market is affected by a more or less strict employment protection legislation. The identification of these main drivers of wage determination is crucial for unemployment policy design since they shape labor market outcomes through their pressure on wages.

The first of these drivers is productivity growth, it reflects efficiency gains and is a common factor across all economies. In the absence of productivity growth, real wages in all economies would have displayed a downward trend, relatively flat in the Anglo-Saxon and Nordic countries (the United States, the United Kingdom, Finland and Sweden), and relatively steep in Japan and continental Europe (France, Italy and Spain).

The second structural driver is deunionization, which has a particularly strong effect in Japan, followed by the continental European countries (except Spain). The weakening of union power has had much less of an impact on wages in the Anglo-Saxon countries, which represent the paradigm of deregulated markets, and no significant effect in the Nordic countries. This confirms the well-known result that union power is fundamentally innocuous to the labor market in these economies.

The third structural driver is trade. While the impact of trade on wage setting is found to be simply irrelevant in the closed economies of the United States and Japan, our counterfactual simulations show that trade has prevented wages from increasing in all of the European economies (except Sweden, where its impact has been minor). The strongest wage effects of trade were observed in Italy and Spain, suggesting that labor costs have been critical to these countries' adjustment to the new market conditions brought about by the globalization process.

Lastly, we have also shown that, by preventing real wages from rising further and thereby enhancing the wage–productivity gap, deunionization and trade are significant contributors to the continuous fall in the labor income share.

A version of this essay has been published by the International Labour Organization's academic journal, the International Labour Review (2013, issue 2), coauthored with Hector Sala.

3.2 Essay 2: "The determinants of capital intensity in Japan and the U.S."

Capital intensity (i.e. the capital-per-worker ratio) is usually considered as an input in growth accounting and the empirical assessment of its determinants has been a rather neglected topic. This essay presents an analytical setting that includes demand-side considerations to the single-equation capital intensity model of the type used in Antràs (2004) and McAdam and Willman (2013). By including product demand uncertainty in a monopolistic competition framework we are able to include demand-side forces in the determination of the capital stock per worker. The resulting empirical model of capital intensity includes relative factor cost (which is the key supply-side driver), relative factor utilization (which is the demand-side driver), a time trend (as a proxy for constant-rate technological change), and other relevant controls such as international trade and taxation. It is applied to the cases of Japan and the U.S. with individual time-series analysis.

The estimation results confirm the relative cost of production factors as a key supply-side driver of capital intensity yielding, also, plausible estimates of the elasticity of substitution between capital and labor. The two proxies we consider for the demand-side pressures are also found relevant. This result calls for a wider approach than the usual one when working with production factor demands and, as we have done, when examining the determinants of capital intensity.

This essay also uncovers the possibility of a different nature of technological change in Japan and the US. As argued, this very difference provides an explanation of the different evolution of capital intensity in Japan and the US, and even of their contrasted growth models; Japan having been, traditionally, one of the great world net exporters; and the

US having been, and being, one of the greatest net importing economies.

Policywise, our results warn about a simplistic design of policies exclusively based on supply-side considerations. On the supply-side, our finding also calls for a careful design of policies affecting firms' decisions on investment and hiring. The reason is that these policies crucially affect the procyclical behavior of the ratio between the rates of capacity utilization and (the use of) employment, since in economic expansions the capacity utilization rate tends to increase proportionally more than the employment rate, probably because in the very short run it is less costly to use already installed capacity than to hire new workers.

3.3 Essay 3: "Heterogeneous labor demand: sectoral elasticity and trade effects in the U.S., Germany and Sweden."

This essay analyzes the heterogeneity in labor demand from two empirical perspectives.

On the one hand, we compute the sector-level elasticity of labor demand and find that these values vary significantly across economic activities. They are generally higher in the U.S. and in Sweden than they are in Germany. According to our results, there is no heterogeneous rule regarding whether services sectors or manufactures are more or less flexible. In sum, a one-size-fits-all approach to labor market policy will probably have very dissimilar results depending on economic activities.

On the other hand, we investigate the employment effects of higher exposure to international trade. We do this by augmenting a standard labor demand model with openness to trade in the empirical employment equation, first in its aggregate version, and later disaggregating openness to trade into four variables according to four types of merchandise: manufactures, services, agriculture and fuel. Openness to trade presents a non-negative effect on employment (neutral in Germany and positive in the U.S. and Sweden). But new insights come along with the disaggregation of openness to trade in the aforementioned subcategories. Higher trade in manufactures has a positive effect on employment, as expected, in the U.S. and Sweden. But, a larger degree of openness to trade in services exerts a negative effect on employment in the U.S. and a positive effect in Sweden.

We believe that this last result may be associated to the growing importance of imported services in the U.S. economy, and the important role that service industries already play, in contrast to Sweden, where the services share of the economy is still not as large, and there may be room to increase trade in services and boost domestic employment.

Lastly, this essay also verifies the presence of labor-saving technical change in the three countries studied. This finding is a common result in related literature (Klump *et al.* 2012, Feldmann 2013). In particular, in the U.S. and Sweden there is a similar rate of labor efficiency growth. Since there is a decelerating employment effect of technical change, this smaller rate of efficiency growth in Germany's case may help in explaining its differentiated employment performance over the last decade.

3.4 Policy implications

In all, this dissertation intends to enhance knowledge about labor market dynamics from a macroeconomic perspective. Our intention has been to present strong arguments and corresponding evidence on the determination of labor market dynamics. Consequently, important policy implication arise.

We show how the simultaneous fall in union power and exposure to international trade experienced in recent years has undermined the labor income share, which has important distributive consequences. Hence, policy aimed at improving redistribution should take the phenomenona of deunionization and trade exposure under careful consideration.

We also outline the close connection between economic growth drivers and labor market outcomes. Moreover, we analyze the factors behind the evolution of one of those drivers, capital intensity. In that analysis, we show that demand-side forces must be considered in policy making. Active labor market policy should be undertaken by the public sector, taking into account that demand-side variables can positively shape labor market outcomes. In other words, not only the reduction of labor costs and efficiency growth are the remedies for aching labor markets. Our results show that demand-side variables may provide with robust macroeconomic results.

Additionally, we call for policy strategies designed to address sectoral specificities. Labor market outcomes depend strongly on particular dynamics of each economic sector. These specificities respond to industrial characteristics. For example, a particular sector may produce tradable or non-tradable goods or services, it may use a higher or lower proportion of imported inputs, its production chain may be more or less involved with commodities (such as oil, metals, or grains), it may be more or less exposed to local and foreign competition, among other factors. These sector-level dynamics are also investigated in this dissertation to conclude that policy addressed to improve labor market outcomes must have sectoral-specific ramifications.

The remainder of the dissertation is structured in three main essays with the research paper structure and content summarized above.

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Essay 1

Productivity, deunionization and trade: Wage effects and labor share implications[2]

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Abstract

A key feature of standard macroeconomic and labor market models is the one-to-one relationship between wages and productivity. This taken for granted, empirical studies have extensively focused on the wage and unemployment impacts of 'unfriendly' labor market institutions, and have left aside other considerations. In contrast, in this paper we look at the long-term implications for wages of productivity growth, deunionization, and international trade. Once controlled for this productivity effect, we document an underlying downward trend in wages that is relatively flat in the Anglo-Saxon and Nordic countries (US, UK, Finland, and Sweden), and relatively steep in Japan and Continental Europe (France, Italy, and Spain). This downward trend is mainly associated to changes in the labor relations system of these countries –represented by the evolution of trade union density– and their growing exposure to international trade –measured by the degree of openness–. Our analysis is useful to interpret the fall in the labor income share experienced by these economies in 1980-2010.

[2] A version of this Essay has been published by the International Labour Review (ILO, UN). The full reference is: Judzik, Dario and Hector Sala (2013) "Productivity, deunionization and trade: wage effects and labour share implications", *International Labour Review*, 152(2), 205-236.

1 Introduction

Fully flexible wages ensure labor market clearing and leave unemployment as a voluntary phenomenon. If unemployment is involuntary, as we perceive in the society, conventional wisdom asserts that it is because wages stand above their market-clearing level. Consequently, it has become standard to look at the causes that push wages above their equilibrium level. The conclusion reached points to a set of "unfriendly" labor market institutions (or regulations) that prevent labor demand and labor supply to meet at the full-employment level.

This paper estimates wage equations for a selection of eight OECD economies representative of the Anglo-Saxon, Nordic and Continental European countries (plus Japan). These groups were defined by Daveri and Tabellini (2000) according to the characteristics of their fiscal and welfare state systems. The literature focusing on the effects of institutional wage-push factors has since then tried to disentangle their impact by resorting explicitly or implicitly to this classification. In this paper, in contrast, we take a fresh look at the global forces driving employment compensation. We show that, irrespective of this standard classification, pay determination in last decades has been conditioned by three structural phenomena: (i) productivity growth, which denotes efficiency progress; (ii) deunionization, as reflection of the labor market deregulation process; and (iii) the growing exposure to international trade, resulting from globalization.

In the standard approach,[3] the impact of the labor market institutions is generally assessed by specifying a reduced form equation where unemployment depends on institutions, shocks, and a demand-side control:

- **Institutions**. labor market institutions are normally classified in four sets connected to the employment protection legislation (EPL), the unemployment protection legislation (UPL), trade union power, and the tax system. An important characteristic of the variables related to the first three sets is that they typically consist in indices capturing the relative intensity of EPL, UPL, and the degree of coordination of collective bargaining.[4] Because long annual time-series of these indices are non-available, the estimation is usually conducted on five-year averages. The resulting reduction in the time dimension of the sample period is compensated

[3] See, among others, Nickell (1997), Elmeskov *et al.* (1998), Blanchard and Wolfers (2000), Daveri and Tabellini (2000), Nickell *et al.* (2005), and Bentolila and Jimeno (2006).

[4] In EPL we find indices measuring the strictness of employment protection, and regulations on labour standards (working time, minimum wages, fixed-term contracts). In UPL we find expenditures in the provision of public social security services, the benefit replacement rate, and the generosity and duration of unemployment benefits. Related to trade union power we have trade union density, union coverage, and the degree of centralisation of the bargaining process. The tax system may include direct taxes (payroll taxes), indirect taxes, and the tax wedge.

by the inclusion of enough cross-section units and, consequently, panel data has become the standard estimation method in the field. Moreover, the fact that the analysis is conducted on five-year averages is rationalized on the grounds of focusing on equilibrium relationships making abstraction of business cycle considerations. No attention is thus paid to dynamics and the role of adjustment costs.

- **Shocks**. Shocks in oil prices, the terms of trade, interest rates, and productivity (or total factor productivity) are typically considered to complete the modeling of the supply-side together with the institutions.

- **Demand-side control**. In mainstream analysis, demand-side considerations have been relegated to a minimum. It has become standard to introduce a single demand-side control, which varies with the dependant variable. If unemployment is explained, this control is the change in inflation. When the dependent variable is the real wage, it is the unemployment rate what controls for demand-side pressures.

In contrast to the standard practice of estimating reduced form unemployment equations, Nunziata (2005) focuses on how these three sets of factors affect wage setting. His main statement was that "labor market regulations explain a large part of the labor cost rise in OECD countries in the last few decades once we control for productivity" (ibid, p. 435). In turn, the second main statement was that his results are consistent with the findings in Nickell *et al.* (2005), where most of the unemployment increase is associated to these regulations. No surprise, therefore, on the conclusions reached, even if the explained variable was not the unemployment rate, but directly the average real wage.

The problem with these results are diverse. First of all, the role played by productivity is not appropriately captured because its business cycle component is ignored, only its trend component (obtained by filtering the series using the Hodrik-Prescott filter) is considered in the regressor. Beyond that, the elasticity of wages with respect to productivity is larger than one and thus contrary to theory (Blanchard and Katz, 1999). Empirically this is also problematic because the omission of the business cycle component and the non-validity of the long-run one-to-one relationship between wages and productivity may cause biases in the estimation of the other variables' coefficients. Second, apart from collapsing available information into few time data points (thereby compressing data variability), the use of five-year averages imply the underlying assumption of perfect business cycle synchronization across the 20 economies considered. Existing evidence, however, is not supportive of this perfect synchronization. Third, the use of panel data, and thus the imposition of common slopes, is at odds with the acknowledged heterogeneity of the countries considered (Nunziata, 2005, tests the poolability of his data and is a clear exception in the standard literature). We should recall, on this account, Daveri and Tabellini's (2000) classification of countries which, presumably, should entail a careful analysis by groups of economies (if not by countries themselves) according to their structural characteristics. This is not done in most of the mainstream literature.

In contrast to Nunziata (2005), we endeavor to estimate country-specific dynamic wage equations, where each of them may display a different lagged structure –reflecting different adjustment costs–, and include a particular set of explanatory variables –due to the specific productive and institutional environment in which wage setting takes place. Of course, there are a number of common regressors, but all of them have different short- and long-run effects on wages (except productivity in the long-run). Finally, rather than relying on dummy variables so that "the empirical counterpart of each institutional dimension is represented by an indicator" (ibid, p. 439), we consider variables with long enough time-series availability. Since we conduct country-specific estimations, we are interested in identifying common drivers so as to provide a comparative analysis based on common grounds. Three driving forces, representative of major phenomena in last decades, are found of paramount importance in wage determination:

- **Productivity**. Labor productivity is the central determinant of wages. In this paper we ensure that all estimated equations satisfy the condition of a unit long-run elasticity between wages and productivity. This is crucial to comply with standard theories of wage determination –efficiency wage, insider-outsider, or union models–, and undertake a new empirical exercise where we evaluate the medium-term impact of productivity growth on wages.

- **Deunionization**. We interpret the evolution of trade union density as a global indicator of the changes experienced by the labor market in the advanced economies. Naturally, these changes have taken different forms and intensities, and have followed different routes, but arguably they have also had a common reflection in the weakening of the workers' bargaining power.

  We have searched for the most promising indicator summarizing the evolution of the wage-push factors, and we find trade union density as our best candidate. Statistically, it is the best-performing wage-push variable across country-regressions, but our choice is not purely statistical. It is also judgemental on the grounds that the sequence of never-ending labor market reforms witnessed in the OECD countries in the last decades, and the weakening of the workers' bargaining power, has resulted either in the reduction of wage growth, in the emergence of irregular work, or in both, to some extent.[5] Our claim is thus that the steep decline in trade union density,

[5]Broadly speaking, labor market reforms have followed two routes. The first was the reduction of EPL and UPL with a consequent a rise in labor turnover, and the second consisted of two-tier reforms

which is pictured in Figure A1 in the Appendix, reflects the global deunionization process of the OECD economies, and is an appropriate indicator of the weakening of the workers' bargaining power. It is probably surprising to see that this choice is consistent with Nickell and Layard's (1999) claim that unions and social security systems are more relevant in unemployment dynamics since "by comparison, time spent worrying about strict labor market regulations, employment protection and minimum wages is probably time largely wasted" (ibid, p. 3030). Furthermore, Addison and Teixeira (2003) argue that there is a problem of subjectivity in the construction of these indices, since the choice of weights and other criteria is quite discretional. Note, finally, that in our specifications we also control for a set of tax variables whenever they are found significant.

- **Trade**. A third major phenomenon is the growing exposure of all advanced economies to international trade. This reflects the acceleration, in recent decades, of the inexorable globalization process. We follow the standard procedure (IMF, 2007) and measure this phenomenon through the degree of economic openness (exports plus imports of goods and services over GDP), which is also pictured in Figure A1. It is worth pointing out that, despite their growing exposure to international trade, US and Japan remain as closed economies in contrast to the rest. Interestingly, these are also the two countries where we failed to find openness as a significant determinant of wages.

  Although the impact of trade on unemployment has been widely analyzed (see Felbermayr *et al.*, 2011, for a recent contribution), but not on wages at the aggregate level. This paper shows that this impact is critical in the open economies considered.

Beyond controlling for the influence on wages of these three major phenomena, our estimated equations contain unemployment and different fiscal variables as additional controls. As standard, unemployment accounts for downward wage pressures stemming from labor markets with excessive supply, while tax variables such as direct, indirect, or payroll taxes are among the conventional set of wage-push factors driving employment compensation upwards.

Once provisioned with empirical wage equations for US, UK, Finland, Sweden, France, Italy, Spain, and Japan, we use them to conduct dynamic accounting simulations in which we examine to what extent productivity, deunionization, and international trade

prompting a general use of temporary work that has also caused large increases in job flows. These reforms have centred on marginal flexibilizations of EPL that have generated dual labor markets featuring high and low levels of worker protection. Although the consequences of this two-sided strategy have been widely examined in terms of the resulting labour market volatility (Sala *et al.*, 2012), another salient result has been the fall in union power (Checchi and Lucifora, 2002; Arpaia and Mourre, 2012).

have contributed to shape the trajectories of wages in 1980-2010. Our findings are diverse and challenge conventional accounts of the unemployment problem based on the excessive range of unfriendly labor market institutions. The first salient result is that in the absence of productivity growth real wages display a downward trend with cumulative falls close to 8% in the US, near 12% in the UK, 5.4% in Finland (by 13.6% since 1991), and around 12% in Sweden. This downward trend is steeper in Japan, with an overall fall of 18.4%, and Continental Europe: 14.5% in France, 26.8% in Italy, and 17.0% in Spain if we exclude the end years of the Spanish wild ride just before the crisis. The second main result is that deunionization and trade (or globalization) are key drivers of this falling trend. On one side, the deunionization process has prevented wages from increasing near 10% in the Anglo-Saxon countries, between 10% and 20% in the Continental European ones, and by more than 20% in Japan. No effects, though, are identified in Scandinavia on this respect. On the other side, trade has prevented wages to increase by almost 5% in the UK and France, by close to 10% in Finland, and by more than 25% in Spain and Italy. These, of course, are not all driving forces at work, but they are, quantitatively, the most relevant ones. The third outcome of our analysis is that deunionization and trade, as key contributors to wage control, have played a fundamental role in widening the wage-productivity gap. This result adds to a growing literature now exploring the causes of the continuous fall in the labor income share (see Table 5).

These findings align our work with those skeptical of the conventional wisdom. Indeed, although largely accepted, the mainstream view has also been target of criticism in recent years. Baker *et al.* (2005) conclude that the results are not robust to variations in variable specification, time period and estimation method; rather, "they seem dependent on the particular measures of the institutions used and on the time period covered" (ibid, p. 40). Baccaro and Rei (2007) stress that it is unclear whether there really is robust empirical support for the view that unemployment is caused by labor market rigidities and should be addressed through systematic institutional deregulation. Seemingly, Freeman (2005, 2008) questions the idea that, in the absence of institutions, labor markets would clear and unemployment would be inexistent or very low. He states, moreover, that works like Nickell (1997) and Nickell *et al.* (2005) were well received and widely cited but their empirical results are not accurate (Freeman, 2008, p. 21). Arpaia and Moure (2012) refer to the endogenous nature of the labor market institutions stressing that a one-size-fits-all approach is unrealistic: labor market institutions should respond to each country's needs, structure and idiosyncrasy. Regarding wages specifically, Podrecca (2010) finds that only some of the many usually cited labor market institutions do affect wage setting significantly.

The remaining of the paper is structured as follows. Section 2 deals with the analytical framework. Section 3 presents the estimated wage equations for the eight economies considered. Section 4 shows dynamic simulations where, together with the incidence of labor productivity, we evaluate the impact of the deunionization and globalization processes on the evolution of wages and the labor share of income. Section 5 concludes.

2 Analytical framework

2.1 Theoretical underpinnings

There is a vast literature on microfounded wage setting models. On one side, they have been extensively used to explain staggered nominal wages and prices: Taylor (1979), Rotemberg (1982), and Calvo (1983). On the other side, they have rationalized the existence of a wage setting curve (or positive relationship between real wages and employment) through efficiency wage models (Shapiro and Stiglitz, 1984) or insider-outsider models (Lindbeck and Snower, 1989). Whatever is the modeling strategy, a common feature is the unit long-run elasticity between wages and productivity.

One of the simplest ways to show this result is to assume the following Nash bargaining process. In exchange of their work, employees receive an average compensation W, while firms obtain $Y/N - W$, which is the workers' average product Y/N lessened by the average compensation they receive in exchange (in case of an individual negotiation, the reference would be the marginal product rather than the average one). The solution of this problem involves the maximisation of the following program:

$$\Omega = (W)^{\mu}\left(\frac{Y}{N} - W\right)^{1-\mu},$$

where μ and $(1-\mu)$ are, respectively, the workers' and the firms' bargaining power. Taking the first order condition with respect to wages, we have:

$$\frac{d\Omega}{dW} = \mu\,(W)^{\mu-1}\left(\frac{Y}{N} - W\right)^{1-\mu} + (1-\mu)\,(W)^{\mu}\left(\frac{Y}{N} - W\right)^{-\mu}(-1) = 0,$$

which rearranged implies:

$$W = \mu\left(\frac{Y}{N}\right). \tag{1}$$

This expression implies a one-to-one relationship between wages and productivity[6].

[6]Note that in a more complex bargaining model, unemployment is the outcome of the chosen real wage and equation (1) becomes: $W = \mu\left(\frac{Y}{N}\right) + (1-\mu)u$, where u is the unemployment rate (for textbook cases, see Cahuc and Zylberberg, 2004). See also the model developed in Walsh (2012), which yields this equation with the inclusion of effort (equation 13 on page 641).

In applied work, this relationship is estimated in logarithms so that the estimated coefficients can be interpreted as elasticities. Of course, the crucial estimated parameter is the one capturing the relationship between wages and productivity. For example, in the simple case of equation (1) we would have $w = \alpha_0 + \pi$, where $w = \log(W)$, $\alpha_0 = \log(\beta)$, and $\pi = \log(Y/N)$. This would imply the estimation of

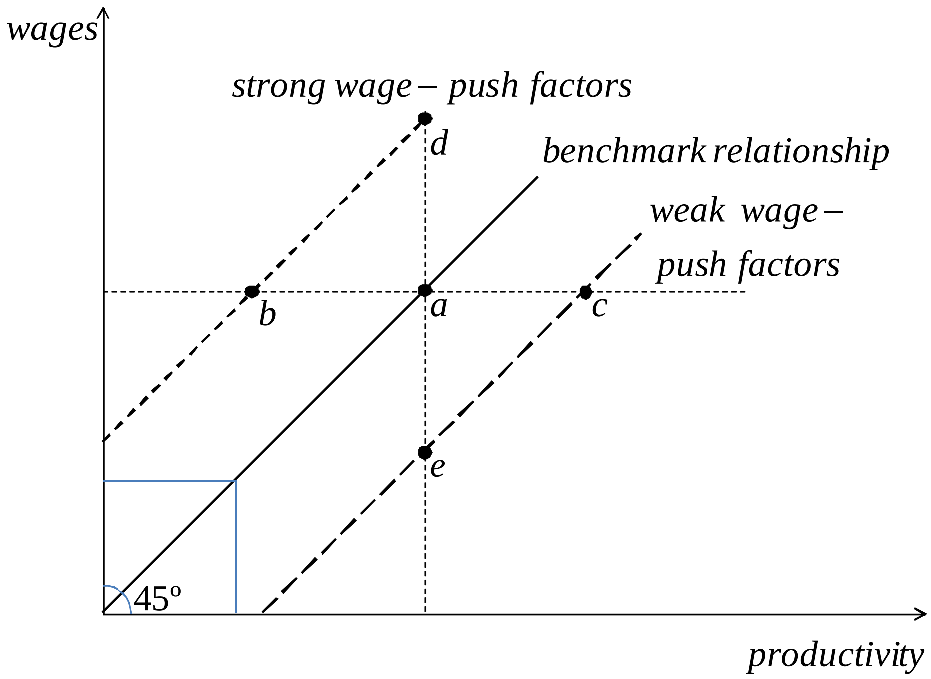
$$w_t = \alpha_0 + \alpha_1 \pi_t + \varepsilon_t, \tag{2}$$

and testing the null hypothesis $H_0 : \hat{\alpha}_1 \simeq 1$. In practice, however, this simple empirical equation is augmented at least in a twofold direction: (i) inclusion of dynamics to account for costly adjustments; and (ii) inclusion of extra control variables such as unemployment and a set of variables representative of wage-push factors.

Figure 1 represents the unit long-run relationship between wages and productivity (if both variables in logs). The benchmark situation is the one depicted by the continuous line departing from zero, while the parallel upper dotted line shows a situation with strong wage-push factors. This parallel shift indicates that, under strong-wage push factors, the same wage is achieved with a lower productivity level (as indicated by point b) or, given the same efficiency, wage compensation is higher (as in point d). In turn, in a scenario of weak wage-push factors, as the one depicted by the lower dotted line, the same wage is achieved with a larger productivity level (point c) or, alternatively, wages are lower (as in point e) at the same level of efficiency. This clarifies Nunziata's (2005) claim that "unemployment has to increase in order to balance the pressure on wages induced by institutions. It follows that the change in unemployment tends to be bigger when the institutional effects are bigger in the opposite direction." (ibid, p. 459).

Consider, for example, a situation in which strong institutions place wages at point d. Nunziata states that the difference between a and d needs to be compensated with high enough unemployment. This line of reasoning creates a direct positive relationship between strong wage-push factors and unemployment without further considerations.

Figure 1. Wages and productivity in the long-run.



The difference between the benchmark situation and the dotted lines is essentially what the standard empirical approximations to wage setting try to measure.

2.2 Related literature

Blanchard and Katz (1999), Nunziata (2005), and Hatton (2007) are among the few studies focusing on wage determination from an aggregate perspective.

Following Blanchard and Katz (1999, p. 70), "most efficiency-wage or bargaining models deliver a wage relation that can be represented (under some simplifying assumptions about functional form and the appropriate indicator of labor-market tightness) as:

$$W_t - P_t^e = \mu b_t + (1 - \mu)\, y_t - \beta u_t + \varepsilon_t,$$

where b is the log reservation wage and y is the log of labor productivity." The value of μ, $0 < \mu < 1$, indicates which of these two variables is more influential in wage setting. The reservation wage depends on the generosity of any form of public income support on the unemployed, notably the unemployment benefits. Given that they are institutionally fixed as depending on past wages, and considering also the aspirations wage literature, Blanchard and Katz (1999) assume a linear dependence between the reservation wage and past wages. They also find reasonable to assume that nonlabour income also grows with productivity. Because the reservation wage depends also on nonlabour income, they postulate the following reservation wage equation:

$$b_t = a + \lambda (W_{t-1} - P_{t-1}) + (1 - \lambda) y_t.$$

The reservation wage is homogeneous of degree 1 in real wages and productivity (in the long-run) because this is consistent with the fact that technological change does not lead to a persistent trend in the unemployment rate. Combination of these two equations yield the benchmark encompassing wage-setting equation here reproduced as (3) which, they claim, summarizes most wage-setting theoretical models:

$$W_t - P_t^e = \mu a + \mu\lambda (W_{t-1} - P_{t-1}) + (1 - \mu\lambda) y_t - \beta u_t + \varepsilon_t, \tag{3}$$

where W is the nominal wage, P^e are expected prices, P actual prices, y is productivity, and u is the unemployment rate; a, μ, and λ are parameters; and ε is a residual (ibid, equation 5, p. 70). Blanchard and Katz (1999) make use of this expression to compare the US and Europe on the basis of the different values taken by μ and λ in the two areas. Our analysis, however, takes a different direction. For our purposes it is just interesting to rewrite equation (3) as

$$w_t = \alpha_0 + \alpha_1 w_{t-1} + (1 - \alpha_1) \pi_t - \alpha_2 u_t + \varepsilon_t. \tag{4}$$

Hence, with respect to the microfounded simple equation (2), we can see that equation (4) incorporates dynamics and the unemployment rate term, and maintains the long-run one-to-one relationship between wages and productivity. Although this expression is much closer to our estimated empirical models, we still lack some additional variables.

To complete the picture, we turn our attention to the works by Nickell (1998) and Nunziata (2005). The latter starts his analysis from the following wage equation (equation 3, p. 438):

$$W - P = -\theta_2 \ln(u) - \theta_3 \Delta \ln(u) + z_w - \theta_1 (P - P^e), \tag{5}$$

which in fact corresponds to equation 4 in Nickell (1998, p. 803), where z_w captures all exogenous factors influencing wages. Nunziata estimates an *ad-hoc* generalization of this model that includes endogenous persistence $(W_{t-1} - P_{t-1})$, trend productivity (π), a vector of exogenous wage pressure factors $(\mathbf{z}_{wt})$, and a vector of nominal and real macro shocks ($\mathbf{s}_t$) that replaces $(P - P^e)$:[7]

$$W_t - P_t = \beta_1 (W_{t-1} - P_{t-1}) - \beta_2 \ln(u_t) - \beta_3 \Delta \ln(u_t) + \beta_4 \pi_t + \boldsymbol{\gamma}' \mathbf{z}_{wt} + \boldsymbol{\upsilon}' \mathbf{s}_t. \tag{6}$$

This equation clearly reflects the conventional wisdom on wage determination. It pays explicit attention to the role potentially played by a set of wage-push variables and a set of macroeconomic nominal and real shocks. In the first set, Nunziata considers indicators for the degree of employment protection, the unemployment benefit replacement rate, union density, bargaining coordination, and the tax wedge. For the second set, he considers terms of trade shocks, the acceleration in TFP, the acceleration in money supply, and a home ownership variable proxying low labor mobility. One of the conclusions reached is that "on average, the major labor cost changes are generated by taxation, the benefit replacement ratio and employment protection" (Nunziata, 2005, p. 459).

In a similar vein, Hatton (2007) postulates the following real wage equation in error correction form:

$$\Delta (W - P)_t = \beta_0 + \beta_1 \Delta \pi + \beta_2 [\pi_{t-1} - (W - P)_{t-1}] - \beta_3 u_{t-1} + x_t, \tag{7}$$

where π is labor productivity (not its trend component as in Nunziata), and "x represents a vector of wage pressure variables including a stochastic term" (ibid, p. 480). For our purposes,[8] it is interesting to write it as:

$$w_t = \alpha_0 + \alpha_1 w_{t-1} + (1 - \alpha_1) \pi_t + \alpha_3 \Delta \pi_t + \alpha_4 u_{t-1} + x_t. \tag{8}$$

This is the equation we estimate next, although we express the rate of unemployment in current terms, and we impose the one-to-one long-run relationship between wages and productivity not ex-ante, but once we ensure that the restriction cannot be rejected by the data.

[7] This generalisation, however, clearly differs from Nickell's (1998) model whose equation 3 (p. 803) is $P - W = z_p - \beta_2 (P - P^e)$, which rearranged gives $(P - P^e) = \frac{1}{\beta_2} (z_p + W - P)$, and substituted in equation (5) yields $W - P = -\theta_2 \ln(u) - \theta_3 \Delta \ln(u) + z_w - \theta_1 \left(\frac{1}{\beta_2} (z_p + w - p) \right)$. Rewritten, this is equivalent to: $W - P = -\beta_2 \ln(u) - \beta_3 \Delta \ln(u) + \beta_4 z_w - \beta_5 z_p$, where $\beta_2 = \frac{\theta_2}{1+\theta_1/\beta_2}$, $\beta_3 = \frac{\theta_3}{1+\theta_1/\beta_2}$, $\beta_4 = \frac{1}{1+\theta_1/\beta_2}$, and $\beta_5 = \frac{\theta_1/\beta_2}{1+\theta_1/\beta_2}$. Therefore, the real wage depends negatively on the log and on the growth rate of the unemployment rate, positively on wage-pressure factors, and negatively on price pressure factors. Note that, in contrast to Nunziata (2005), no role is given to labor productivity.

[8] Hatton presents this wage equation in conjunction with another one explaining unemployment. This system of equations is used to investigate the long-run effects of productivity growth on the non-accelerating inflation rate of unemployment (NAIRU).

3 Econometric analysis

3.1 Data

We use annual data for years 1960-2010 taken from the OECD Labor Market Indicators and from the OECD Economic Outlook no. 90. The sample period starts in the early 1960s in all countries but Spain. This is due to lack of data on trade union density, given that unions were legalized in 1977 and that the corresponding time-series start in 1981. In turn, union time-series are available for all countries up to 2010, but for France (up to 2008) and Spain (2009). Definitions of the variables used are provided in Table 1. It is important to note that standard indicators representative of wage-push factors such as employment or unemployment protection indicators were also considered. Some examples are the EPL indicator from the OECD Employment and Labor Market Statistics, the benefit replacement rate from the CEP-LSE database (see Nickell, 2006), or the unemployment benefit indicator from Allard (2005). However, in view that they sometimes shortened the sample period substantially, did not provide the best fit, and/or they simply were non-significant, these indicators were discarded from the pool of regressors. As a consequence, they are not defined nor considered in the rest of the analysis.

Table 1. Definitions of variables.

w	real compensation per employee		
π	labor productivity $(= y - n)$	τ^h	direct taxes on households, % GDP
y	GDP (at market prices)	τ^i	indirect taxes, % GDP
n	employment	τ^d	total direct taxes, % GDP
ls	labor share $(= w - \pi)$	ss	social security contributions, % GDP
u	unemployment rate	tud	trade union density
tr	trade=(exports + imports)/GDP	d^{75}	dummy (value 1 1975 onwards)
c	constant	Δ	difference operator

Note: All variables are expressed in logs, except the unemployment rate.

Sources: OECD labor Market Indicators and OECD Economic Outlook no. 90.

Although these variables are standard in the literature, two of them deserve a brief comment. The first one is trade union density, which corresponds to the ratio of wage and salary earners that are trade union members, divided by the total number of wage and salary earners. The second one is the labor income share, which in our case (as in Checci and García-Peñalosa, 2010) is not adjusted for self-employment rents.[9] Finally,

[9]Nevertheless, we have compared our series with those provided by the European Commission (in the Ameco database) on adjusted wage shares. Level differences between the series are in the form of parallel

note that in the regressions all variables are treated in logs with the sole exception of the unemployment rate.

3.2 Methodology

As noted in the introduction, following the influential works by Nickell (1997) and Blanchard and Wolfers (2000), among others, the labor market impact of institutions is generally assessed through the estimation of static regressions that use five-year averages of stationary variables (for example, indices proxying the institutional setup). Here, on the contrary, we focus on long-run relationships of non-stationary variables –wages, productivity, and others, as documented in Table 3 below– in the context of the estimation of dynamic models. Therefore, to produce asymptotically unbiased normally distributed estimates of the long-run elasticities, we consider the cointegration analysis within the autoregressive distributed lag (ARDL) framework. As shown by Pesaran and Shin (1999) and Pesaran *et al.* (2001), a central advantage of the ARDL is that it yields consistent estimates of long-run coefficients that are asymptotically normal irrespective of whether the underlying regressors are $I(1)$ or $I(0)$. This approach is also known as a bounds testing procedure for the analysis of level relationships, and is an alternative to the standard cointegration techniques of the Phillips-Hansen semi-parametric fully-modified OLS procedure, and the Johansen maximum likelihood method.

It is important to note that the estimation of a cointegrating vector from an ARDL specification is equivalent to that from an error-correction model. We thus report the ARDL estimates in Table 2, whereas in Table 4 we show the implied cointegrating vectors[10].

In order to ensure that our ARDL estimated vectors conform with those that would be obtained from the Johansen methodology, we also perform the standard Johansen cointegration analysis. We thus use the maximal eigenvalue and trace statistics to confirm that the variables involved in each equation are cointegrated. Then we estimate the corresponding VAR models ensuring that they contain all the variables in our selected wage equations, both the $I(0)$ and $I(1)$ variables, and the same lag ordering. From the estimated VARs we obtain Johansen's cointegrating vectors (which are presented in the third column of Table 4). Then we restrict these vectors to take the values of the cointegrating relationships obtained through the ARDL procedure. Finally, we run

shifts (outwards for some countries, inwards for some other countries). In terms of growth rates, however, there is an almost perfect fit between the two sources in all countries considered. Because our analysis makes use of the variation in these series and is based on estimated elasticities, it follows that our results in terms of the labour income share are not flawed.

[10] The orders of the lags in the ARDL models are selected by the optimal lag-length algorithm of the Schwarz information criterion.

a likelihood ratio test to check whether these restrictions hold, in which case we find evidence that the results from both methodologies do not provide discrepant evidence.

Next subsection presents the estimated equations using the ARDL methodology. In the following one we report the results on the likelihood ratio tests, and validate the implied long-run cointegrating relationships. Once our models are validated econometrically, we use them to assess the driving forces of wages in last decades.

3.3 Estimated equations

In Table 2 we present our selected specifications for each country. Each of them is estimated by Ordinary Least Squares (OLS) and instrumental variables (IV) to discern potential endogeneity biases. Endogeneity could be present, for example, between trade union density and productivity (since labor market regulation may affect productivity as shown by Storm and Naastepad, 2007); between trade union density and unemployment (Checchi and Nunziata, 2011); or even between trade union density and international trade, although Dreher and Gaston (2007) have shown that it is the social dimension of globalization, and not the economic one that we are capturing here through trade, the one that is closely connected with the deunionization process.

Table 2. Estimated equations.

	United States				**United Kingdom**		
	OLS	**IV**	**OLS-R**		**OLS**	**IV**	**OLS-R**
c	0.06 [0.850]	0.04 [0.897]	−0.13 [0.074]	c	−1.39 [0.150]	−1.15 [0.422]	−0.23 [0.072]
w_{t-1}	0.72 [0.000]	0.74 [0.000]	0.75 [0.000]	w_{t-1}	0.58 [0.000]	0.57 [0.000]	0.56 [0.000]
π_t	0.26 [0.002]	0.25 [0.004]	0.26 [*]	π_t	0.52 [0.000]	0.53 [0.010]	0.44 [*]
$\Delta\pi_t$	0.37 [0.011]	0.41 [0.008]	0.41 [0.002]	u_t	−0.28 [0.028]	−0.18 [0.351]	−0.20 [0.065]
tud_t	0.02 [0.121]	0.03 [0.109]	0.03 [0.001]	tud_t	0.06 [0.051]	0.05 [0.160]	0.06 [0.051]
τ_t^d	0.03 [0.020]	0.04 [0.024]	0.04 [0.011]	tr_t	−0.12 [0.104]	−0.11 [0.311]	−0.03 [0.051]
				τ_t^d	0.07 [0.044]	0.11 [0.099]	0.07 [0.032]
				$\Delta\tau_t^d$	0.09 [0.032]	0.14 [0.075]	0.08 [0.052]
Sample		1961-2010		*Sample*		1964-2010	
LL	170.7	—	170.4	LL	143.2	—	142.2
DWH		0.74		DWH		0.89	
ε_{w-pr}^{LR}	0.94	0.95		ε_{w-pr}^{LR}	1.25	1.21	
W	0.53	0.63		W	0.22	0.47	

... Continuation Table 2

	Finland				Sweden		
	OLS	**IV**	**OLS-R**		**OLS**	**IV**	**OLS-R**
c	-0.28 [0.417]	-0.03 [0.948]	-0.21 [0.011]	c	1.32 [0.290]	2.94 [0.583]	-0.53 [0.036]
w_{t-1}	0.80 [0.000]	0.81 [0.000]	0.80 [0.000]	w_{t-1}	0.57 [0.000]	0.60 [0.000]	0.59 [0.000]
π_t	0.21 [0.001]	0.17 [0.069]	0.20 [$*$]	π_t	0.29 [0.003]	0.11 [0.736]	0.41 [$*$]
u_t	-0.16 [0.008]	-0.12 [0.106]	-0.16 [0.007]	u_t	-0.17 [0.122]	-0.23 [0.156]	-0.15 [0.176]
Δssr_t	0.12 [0.000]	0.16 [0.017]	0.12 [0.000]	Δu_t	0.54 [0.012]	0.76 [0.461]	0.37 [0.041]
tud_{t-1}	0.16 [0.048]	0.20 [0.236]	0.16 [0.041]	Δssr_t	0.18 [0.000]	0.22 [0.012]	0.18 [0.000]
tud_{t-2}	-0.13 [0.088]	-0.17 [0.289]	-0.13 [0.086]	τ_t^d	0.12 [0.004]	0.07 [0.578]	0.10 [0.010]
tr_t	-0.03 [0.266]	-0.01 [0.847]	-0.03 [0.040]	τ_t^i	-0.08 [0.002]	-0.08 [0.279]	-0.09 [0.001]
				tud_t	0.11 [0.019]	0.17 [0.106]	0.09 [0.049]
				tr_t	0.13 [0.081]	0.25 [0.480]	0.02 [0.137]
Sample		1963-2010		*Sample*		1964-2010	
LL	143.8	—	143.8	LL	155.0	—	153.6
DWH		0.55		DWH		0.76	
ε_{w-pr}^{LR}	1.04	0.90		ε_{w-pr}^{LR}	0.67	0.28	
W	0.85	0.72		W	0.11	0.42	

... Continuation Table 2

	France				**Italy**		
	OLS	**IV**	**OLS-R**		**OLS**	**IV**	**OLS-R**
c	-0.82	-1.15	-0.21	c	-0.35	-1.46	-0.80
	[0.054]	[0.039]	[0.001]		[0.565]	[0.112]	[0.000]
w_{t-1}	0.78	0.83	0.79	w_{t-1}	0.63	0.61	0.62
	[0.000]	[0.000]	[0.000]		[0.000]	[0.000]	[0.000]
π_t	0.27	0.26	0.21	π_t	0.33	0.45	0.38
	[0.000]	[0.002]	[$*$]		[0.001]	[0.004]	[$*$]
u_t	-0.22	-0.47	-0.07	u_t	-0.25	-0.59	-0.39
	[0.114]	[0.018]	[0.465]		[0.226]	[0.051]	[0.000]
tud_t	0.04	0.02	0.04	tud_t	0.14	0.11	0.13
	[0.001]	[0.098]	[0.000]		[0.000]	[0.021]	[0.000]
tr_t	-0.05	-0.06	-0.01	tr_t	-0.10	-0.18	-0.14
	[0.063]	[0.047]	[0.101]		[0.025]	[0.009]	[0.000]
Sample		1964-2008		*Sample*		1961-2010	
LL	163.5	—	162.3	LL	147.7	—	147.4
DWH		0.08		DWH		0.17	
ε_{w-pr}^{LR}	1.25	1.54		ε_{w-pr}^{LR}	0.89	1.16	
W	0.14	0.07		W	0.45	0.45	

... Continuation Table 2

	Spain				**Japan**		
	OLS	**IV**	**OLS-R**		**OLS**	**IV**	**OLS-R**
c	-0.06	-0.54	-0.23	c	-0.79	-0.63	-0.60
	[0.942]	[0.596]	[0.000]		[0.000]	[0.004]	[0.000]
w_{t-1}	0.79	0.82	0.88	w_{t-1}	0.69	0.72	0.71
	[0.000]	[0.000]	[0.000]		[0.000]	[0.000]	[0.000]
π_t	0.18	0.20	0.12	π_t	0.32	0.28	0.29
	[0.036]	[0.037]	[$*$]		[0.000]	[0.???]	[$*$]
Δss_t	0.25	0.29	0.28	tud_t	0.17	0.15	0.15
	[0.000]	[0.001]	[0.000]		[0.000]	[0.000]	[0.000]
tud_t	0.06	0.04	0.04	d^{75}	0.03	0.03	0.03
	[0.049]	[0.176]	[0.000]		[0.001]	[0.002]	[0.001]
tr_t	-0.05	-0.05	-0.04				
	[0.000]	[0.000]	[0.000]				
Sample		1983-2009		*Sample*		1963-2010	
LL	97.3	—	97.2	LL	161.5	—	160.8
DWH		0.81		DWH		0.09	
ε_{w-pr}^{LR}	0.88	1.14		ε_{w-pr}^{LR}	1.03	1.01	
W	0.53	0.81		W	0.30	0.81	

Notes: Probabilities in brackets; $*$=Restricted coefficient.
LL=Log-likelihood; DWH=Durbin-Wu-Hausman Test;
ε_{w-pr}^{LR}=Long-run elasticity of wages with respect to productivity; W=Wald test.
Instruments on the IV estimation are first lag of each regressor, plus w_{t-2}.

We conduct the Durbin-Wu-Hausman test and we find that none of the estimated equations are suspicious of endogeneity problems (the results of this test for each country are presented in Table 2). The absence of significant endogeneity problems should come as no surprise since two of the key driving forces, deunionization and trade, are independent of one another (Dreher and Gaston, 2007), while the third one, productivity, is restricted so that there is a one-to-one relationship with wages in the long-run. We thus use the OLS estimator which in any case is more efficient than the IV one.

To test the validity of this long-run relationship we conduct a Wald test (whose results are also presented in Table 2). In all cases we conclude that the null hypothesis of a long-run unit elasticity between wages and productivity cannot be rejected. Consequently, we estimate again all equations using OLS (given the results of the Durbin-Wu-Hausman test) and restricting the long-run coefficient of productivity to be unity (given the results of the Wald test). The corresponding estimates are displayed in the third column of country results under the heading OLS-R (where R denotes Restricted). These are our reference estimates and, of course, the ones used in the simulation analysis.

All models capture wage adjustment costs through a single persistence coefficient ranging from around 0.55 to 0.90. UK, Sweden, and Italy have the lowest values around 0.60; the US and Japan come next, with persistence coefficients of at least 0.70; while in Finland, France they get close to 0.80, and in Spain they reach 0.88. Unemployment exerts the expected negative sign in all European countries but in Spain (where wages are not sensitive to a rate of unemployment that has been extremely high and persistent in all years of the sample period), and has no significant influence also in the US and Japan (note that US is the paradigm of a deregulated labor market, whereas Japan is a very specific case, with unemployment rates below 3% until the 1990s' *lost decade*, and with a very particular labor relations system).

Regarding the major driving forces we have, first of all, that the long-run coefficient of productivity is restricted to unity in all equations. It is important to remark that this restriction cannot be rejected in any equation (the results of the corresponding Wald test are reported below the results for each economy). The second noteworthy result is the presence of trade union density, with the expected positive sign, in all specifications. In turn, international trade enters significantly in all open economies, which are the European ones, but not in the closed ones –US and Japan in our sample of countries–, where trade does not play any significant role in wage determination.

Different fiscal variables enter some of the equations as additional controls. These are direct taxes in the US, UK, and Sweden; indirect taxes in Sweden; and social security contributions in Finland, Sweden, and Spain.

3.4 Validity of the long-run cointegration relationships

The first step in testing the validity of the long-run cointegration relationships is to establish the order of integration of the variables involved in the analysis. Table 3 shows the results of the Kwiatkowski–Phillips–Schmidt–Shin (KPSS) unit root test. Although it is conducted on all variables entering the estimated specifications presented in Table 2, we only report the results for variables that are relevant in most of the countries.[11]

Table 3. KPSS unit root tests.

	US	**UK**	**FN**	**SW**	**FR**	**IT**	**SP**	**JP**
w	0.94	0.89	0.93	0.88	0.82	0.72	0.70	0.79
π	0.95	0.95	0.94	0.94	0.92	0.89	0.85	0.89
tud	0.94	0.61	0.71	0.40	0.84	0.22	0.50	0.90
tr	–	0.94	0.90	0.93	0.93	0.94	0.95	–
u	–	0.35	0.61	0.66	0.82	0.66	–	–

Notes: The null hypothesis corresponds to the stationarity of the variable; Critical values are: 0.74 at the 1% level; 0.46 at the 5%; 0.35 at the 10%.

As expected, wages and labor productivity are non-stationary in all economies. The same occurs with trade, thus signifying the structural nature of the globalization process, while trade union density and unemployment behave also as non-stationary variables in most of the countries. The exceptions –at a 5% critical value– can be found in Sweden and Italy in trade union density, and in the UK regarding unemployment. These variables are, therefore, the ones that will be considered in the cointegrating vectors.

Table 4 shows the ARDL cointegrating vectors in the second column. Note that some countries have some empty cells, either because that particular variable did not enter the specification (trade and unemployment in the US and Japan, the latter in Spain) or because the variable is stationary (trade union density in Italy). Two especial cases are unemployment in the UK and trade union density in Sweden, which are considered $I(1)$ on account of its ambiguous properties: at a 10% critical value we cannot reject the null that is a non-stationary variable.[12] In any case, the ARDL values have to be compared with Johansen's cointegrating vectors (in the third column), and it is the Likelihood Ratio test (in the final column) what reveals whether they conform with one another. Note that

[11] We would like to remark, however, that the variables not reported (i.e., those relevant in just some speficic cases) are mainly related to the tax system and, as expected, are stationary in most cases. In other words, they produce the less interesting results which, for the sake of brevity, are not explicitly shown (the only exception is indirect taxes in Sweden, which will be considered in the Johansen cointegration analysis as an $I(1)$ variable).

[12] Moreover, indirect taxes in Sweden behave as a non-stationary variable (the KPSS test yields 0.89) and needs to be included in the cointegrating vector. It takes the extended form $\begin{pmatrix} w & \pi & tud & tr & u & \tau^i \end{pmatrix}$. Table 4 shows its values under the ARDL and Johansen's methodology.

the number of restrictions we impose varies depending on the number of variables entering the cointegrating vector. It ranges from 2, in the US and Japan, to 5 in Sweden.

Table 4. Testing the long-run relationships in the Johansen framework.

	ARDL					Johansen					LR test
	w	π	tud	tr	u	w	π	tud	tr	u	
US	-1	1	0.12	–	–	-1	0.72	0.06	–	–	$\chi^2(2)=4.64[0.098]$
UK	-1	1	0.14	-0.07	-0.45	-1	1.42	0.10	-0.44	-0.85	$\chi^2(4)=4.90[0.298]$
FN	-1	1	0.15	-0.15	-0.80	-1	0.93	0.12	-0.13	-0.17	$\chi^2(4)=4.77[0.312]$
SW			(*)					(**)			$\chi^2(5)=10.9[0.053]$
FR	-1	1	0.19	-0.05	-0.33	-1	1.21	0.15	-0.21	-1.46	$\chi^2(4)=6.16[0.188]$
IT	-1	1	–	-0.37	-1.03	-1	0.97	–	-0.38	-0.85	$\chi^2(3)=2.23[0.525]$
SP	-1	1	0.33	-0.33	–	-1	1.23	0.22	-0.32	–	$\chi^2(3)=0.92[0.819]$
JP	-1	1	0.52	–	–	-1	0.96	0.53	–	–	$\chi^2(2)=3.92[0.141]$

Notes: p-values in square brackets; 5% critical values: $\chi^2(1)=3.84$; $\chi^2(2)=5.99$; $\chi^2(3)=7.82$; $\chi^2(4)=9.49$; $\chi^2(5)=11.07$; (*) ARDL: $\begin{pmatrix} -1 & 1 & 0.22 & 0.05 & -0.37 & -0.22 \end{pmatrix}$; (**) Johansen: $\begin{pmatrix} -1 & 0.78 & 0.21 & 0.09 & 0.14 & -0.02 \end{pmatrix}$.

The results of the LR test allow us to conclude that the specification of the wage equation for all economies yields cointegrated long-run relationships that are indeed robust across methodologies.

4 Simulations

We use our estimated equations to conduct dynamic accounting simulations. We always use the restricted estimates so that the condition of a unit long-run elasticity between wages and productivity holds. We simulate our models in a twofold situation: a first scenario in which all exogenous variables take their actual values, and a second scenario where one of these variables is fixed at its 1980 value. The difference in the fitted values obtained from the two scenarios accounts for the effect on wages of the variable that has been fixed. Hence, our exercise provides answers to questions of the following type: How would the evolution of the real wage have looked like if had, say productivity, not grown since 1980? It is important to stress that we are not claiming that this would have been the true evolution of wages in that case. We are just conducting counterfactual experiments to learn on the driving forces that have shaped the real wage trajectory in the last three decades.

As in Karanassou and Sala (2010), we choose 1980 as the reference year for a variety of reasons: we are looking at structural phenomena and we are interested in a medium- to long-term perspective; we favor country-comparability and the sample period in Spain does not allow us to go further behind; finally, the 1980s are generally considered as an inflection point in many economic dimensions: Keynesian ideas loose their prevalence, most advanced economies undertake extensive and intensive deregulation processes, deunionization becomes apparent, and the globalization course accelerates.

4.1 Productivity

The first simulation exercise corresponds to the example above where productivity is fixed in all countries in 1980 (1982 in Spain due to sample period constraints). The corresponding results are presented in Figure 2 for all countries (we take exponentials to express real wages in absolute values rather than in logs). Solid lines are used for actual values, dotted lines for simulated values. The difference between the two shows the cumulative impact on wages since 1980 of a zero productivity growth rate.

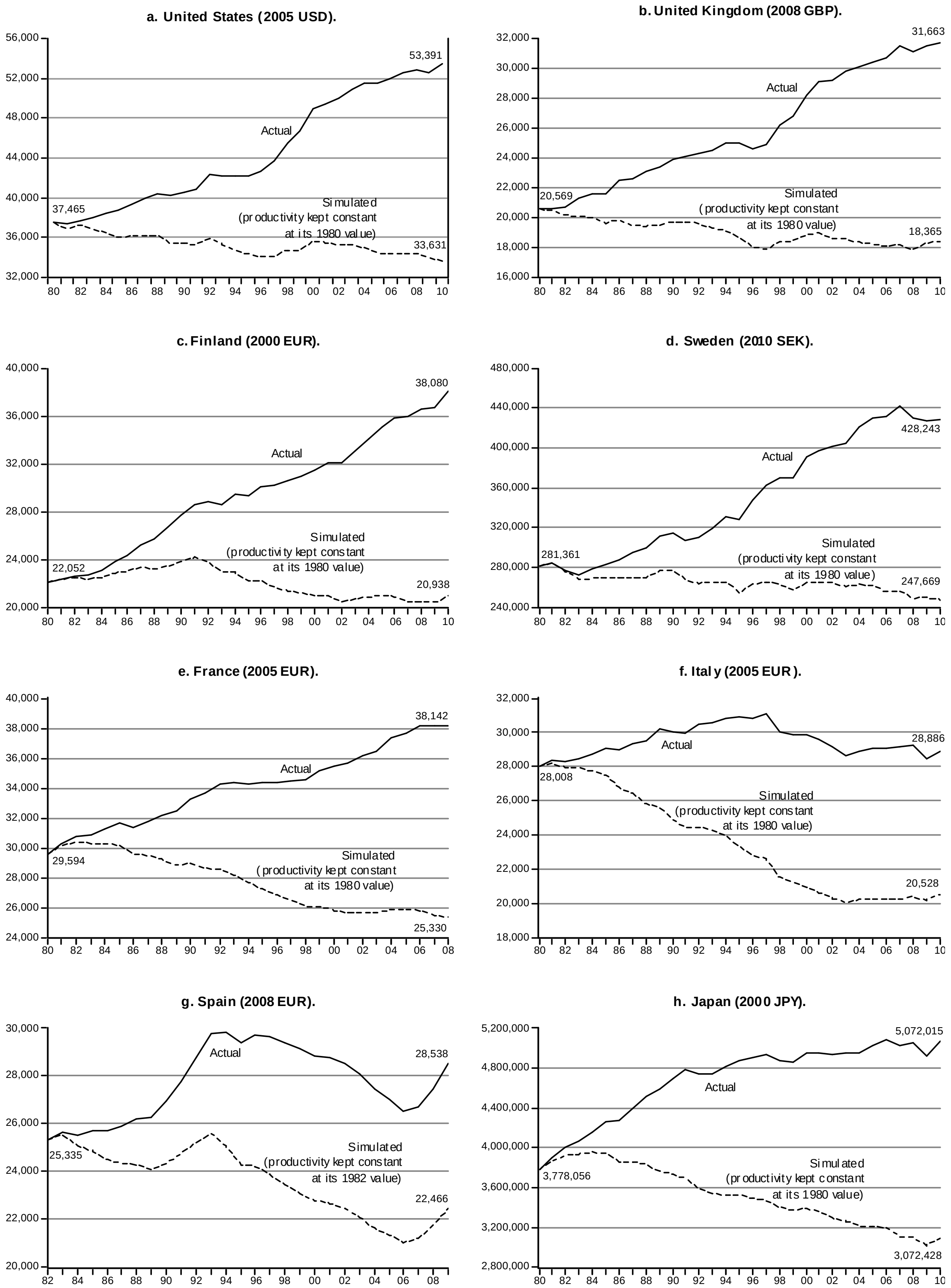
In the Anglo-Saxon and Nordic countries we see a relatively flat downward trend. Real wages would have decreased by close to 8% in the US (from 37,500 in 1980 to 33,600 USD in 2010), and near 12% in the UK (from 20,600 to 18,400 GBP); by 5.4% in Finland (from 22,100 to 20,900 euros), and by around 12% in Sweden (from 281,000 to 248,000 SEK). Finland shows the mildest decrease, overall, but it can be perceived that the fall accelerated since 1991 in parallel with the liberalization process that took place after the deep crisis of the early 1990s (see Honkapohja and Koskela, 1999). Taking the 1991-2010 period, the wage fall in the absence of productivity gains would have been 13.6%.

This downward trend is steeper in Continental Europe. Wages in France go down by 14.5% (from 29,600 to 25,300 euros), while Italy almost doubles this figure with 26.8% (from 28,000 to 20,500 euros). Regarding Spain, if we exclude the end years of the Spanish wild ride just before the crisis, in which wages grew fast, the wage fall in the absence of productivity growth attains 17.0%. If these years are not excluded, this fall would still surpass 11% and resemble that of France (in particular it would have been from 25,300 euros in 1982 to 22,500 euros in 2009). The simulation in Japan yields a similar picture, where wages would have gone down by 18.4% (from 3,800,000 to 3,100,000 JPY).

Overall, although the intensity in the real wage downward trend is diverse, these results provide an unanimous conclusion irrespective of the type of economy considered (Anglo-Saxon, Nordic, Continental European, or Japanese): productivity gains have been the key sustain of real wage progress. Viewed from the opposite side, the conclusion is that, leaving aside productivity, the aggregate influence of the remaining wage determinants has had a negative influence across countries and periods (with only two exceptions in the

1980s in Finland, and the late 1980s/early 1990s in Spain). This influence is quantified next.

Figure 2. Real wage evolution in the absence of productivity growth.



4.2 Deunionization and growing exposure to international trade

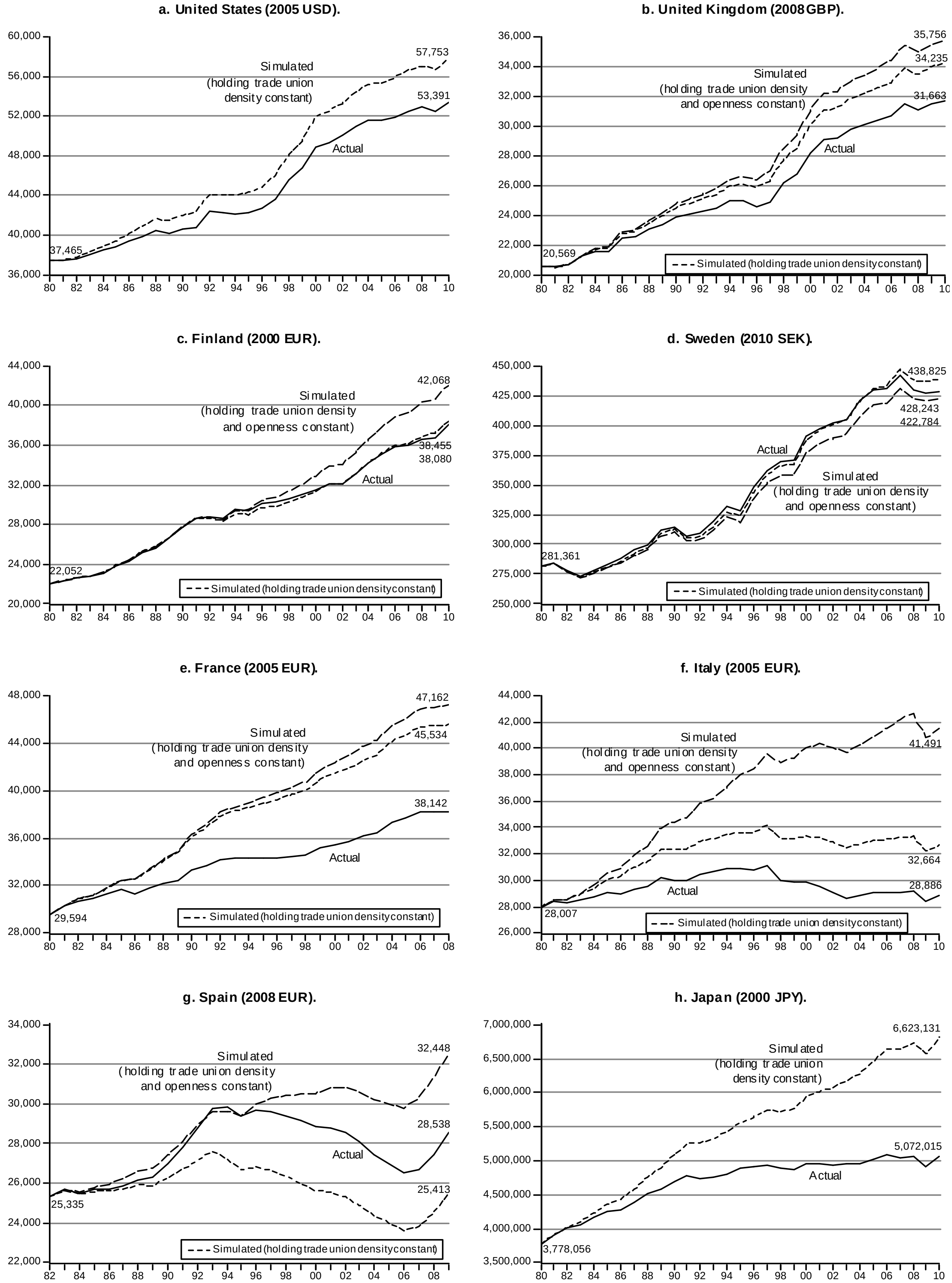
In the next exercise, productivity series are restored to their actual values, whereas trade union density and the degree of openness are kept at their 1980 values in each country. In this way we can assess how deunionization and the growing exposure to international trade have contributed to shape the real wage trajectory in the last decades. The results are presented in Figure 3. The short-dotted line corresponds to the simulation where trade union density is hold constant. The long-dotted line to the simulation where both trade union density and international trade are fixed. The difference between the two accounts for the wage impact of the increasing degree of economic openness (recall that, following the econometric analysis, there is no such impact in the two closed economies US and Japan).

Our estimates show that the deunionization process has contributed to reduce wages in the Anglo-Saxon and Continental European economies (except Spain), and Japan. On the contrary, there is no effect in the Nordic economies. The innocuous wage effect of unions in Scandinavia is generally justified on the grounds of its high degree of centralization of collective bargaining. Here we find this result as the natural consequence of the absence of a secular deunionization process. Trade union density in Finland and Sweden increased in the 1980s and early 1990s, and started going down along with the liberalization process that followed the crisis of the early 1990s. Although the decrease in trade union density has been notable in Sweden, it is still close to 70%, as in Finland.

As pictured in Figure A1, trade union density in the US, UK, France, Italy, and Japan has unambiguously and severely trended down. Figures 3a, 3b, 3e, 3f, and 3h, respectively, show that in the absence of this deunionization process, wages would have gone up to 57,800 USD in the US (instead of 53,400), to 34,200 GBP in the UK (instead of 31,700), to 45,500 euros in France (instead of 38,100), to 32,700 euros in Italy (instead of remaining roughly constant at 28,900), and to 6,600,000 JPY (instead of 5,100,000). This implies extra wage increases of, respectively, 8.2%, 7.3%, 16.3%, 11.6%, and 22.7%. In other words, the deunionization process in 1980-2010 has prevented wages from increasing near 10% in the Anglo-Saxon countries, between 10% and 20% in the Continental European ones (but Spain), and by more than 20% in Japan.

Spain follows a different pattern due to the late legalization of unions in 1977. Trade union density, which was initially low, more than doubled between the early 1980s and 1993, to stabilize afterwards. This union consolidation caused wages to go up by 12.2% –from 25,400 to 28,500 euros. In other words, in the absence of this increase in trade union density, wages in Spain would have been more than 10% lower. This effect is in the range of the union impact found for the other Continental European countries.

Figure 3. Deunionization and openness effects on wages.



The growing exposure to international trade is the third major driving force shaping the trajectory of wages in all European economies, which are all open economies (Figure A1). In the absence of such trend, wages would have been higher in all of them, but Sweden (in Sweden, however, the impact of international trade only explains a 2.4% growth in real wages). Globalization, therefore, appears as a key stopper of wage growth. Interestingly, this is especially the case in the two countries that have suffered the most the consequences of the recent Great Recession.

As it is also shown in Figure 3, in the UK trade prevented wages to increase by further 1,500 GBP (=35,800-34,200), in Finland by 3,600 euros (=42,100-38,500), in Italy by 8,800 euros (=41,500-32,700), and in Spain by 7,000 euros (=32,400-25,400). Although these are all significant magnitudes in levels, there are substantial relative differences. Trade prevented wages to increase by almost 5% in the UK and France, by close to 10% in Finland, and by more than 25% in Spain and Italy. It seems clear that, in these two Southern European economies, labor costs have been a critical adjustment mechanism to the new productive and market conditions brought by the economic integration process.

4.3 Implications for the labor income share

Once examined the wage impact of three major structural phenomena such as productivity, deunionization, and international trade, our final step is to evaluate the resulting consequences for the labor income share.

Regarding the productivity effect, let us go back to equation (1) and rewrite it as $\beta = W/\left(\frac{Y}{N}\right)$. In this simple setting, β represents a labor share that is constant and is thus consistent with the standard Cobb-Douglas framework of analysis (Karanassou and Sala, 2013). Taking a step further, and just for illustrative purposes, we can relax this restriction and consider a time-varying labor share LS_t consistent, for example, with its downward trajectory in last decades (IMF, 2007). Substituting β by LS_t, defining $PR = Y/N$, adding time subscripts, and taking logs (denoted by lower case letters) we have:

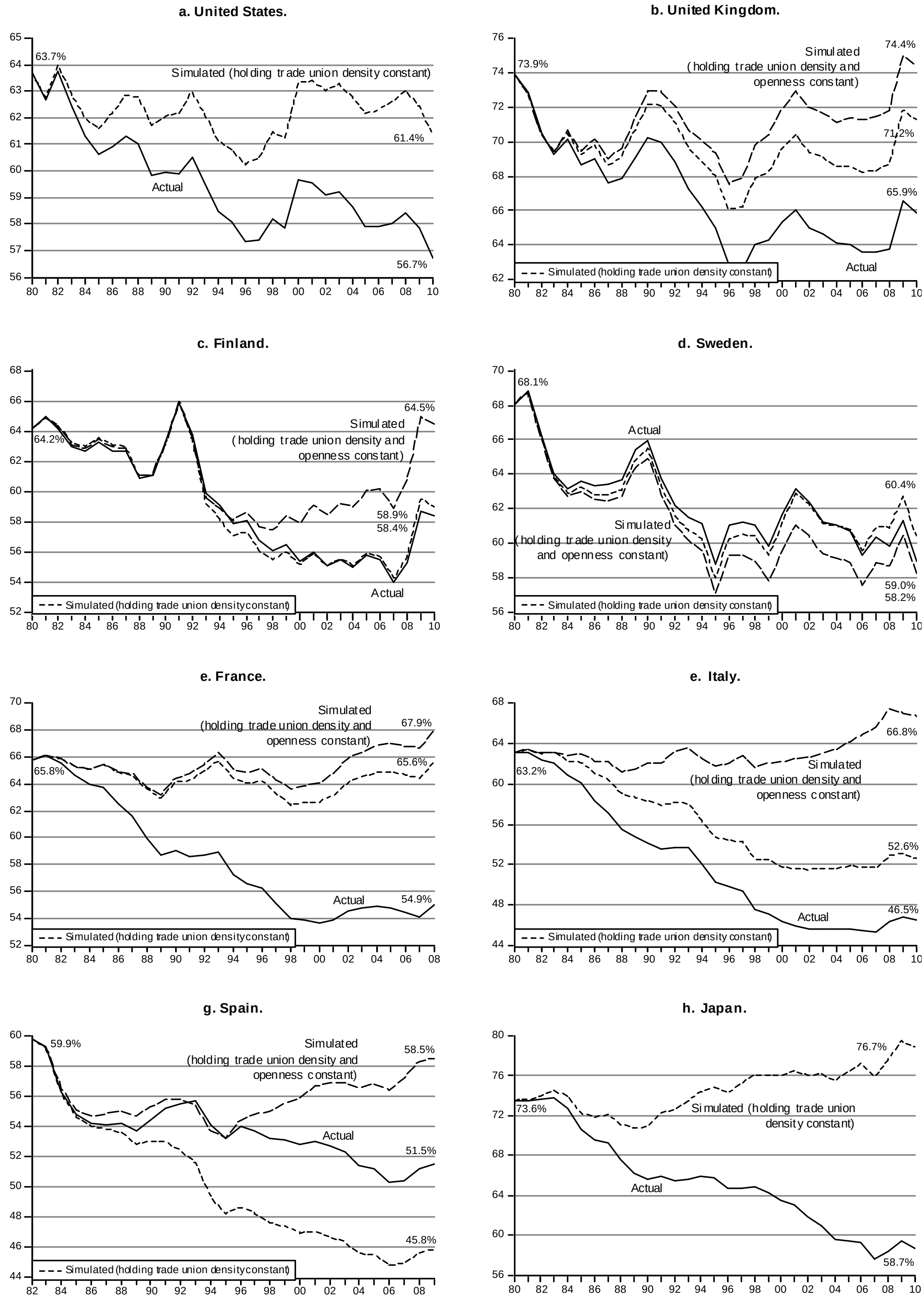
$$ls_t = w_t - pr_t.$$

By taking first-differences, we can see that labor income changes are the outcome of changes in wage and changes in productivity:

$$\Delta ls_t = \Delta w_t - \Delta pr_t.$$

This expression corresponds to the standard conceptualization of the labor share as the wage-productivity gap. It implies that, controlling for productivity (recall that the long-run one-to-one relationship between wages and productivity always holds), we can examine how changes in wages, in response to changes in their driving forces, translate into the labor share.

Figure 4. Deunionization and openness' effects on the labor income share.



We thus proceed as follows. We take our previously simulated wage trajectories (the ones displayed in Figure 3). We compute a simulated labor income share which is the outcome of these simulated wage trajectories minus the actual productivity series. We compare it with the actual evolution of the labor income share, and we examine the implications, for its trajectory, of the deunionization process and the rise in international trade. The outcome of this exercise is presented in Figure 4.

By preventing real wages to rise further, and as a consequence by enhancing the wage-productivity gap, the deunionization process in the US, UK, France, Italy and Japan has contributed significantly to the labor share fall of these economies in last decades. In the Anglo-Saxon countries by around 5 percentage points (pp henceforth), 4.7 pp. in the US and 5.3 in the UK. In Italy by 6.7 pp, in France by near 11 pp, and in Japan by 18 pp.[13] This is the cumulative impact along a time span of three decades. Note, in turn, that in Finland and Sweden the union incidence is negligible since it amounts to less than 1 pp.

Table 5. Literature on union and trade effects on the labor income share.

			LIS effects of:	
Study	**Country**	**Period**	**Unions**	**Trade**
Conyon (1994)	U.K.	1983-1986	+	
Wallace *et al.* (1999)	U.S.	1949-1992	+	
Jayadev (2007)	OECD countries	1972-1996	+	-
Fichtenbaum (2009)	U.S.	1949-2006	+	-
Fichtenbaum (2011)	U.S.	1997-2006	+	-
Schneider (2011)	Western Europe	1980-2005		-
Young and Zuleta (2011)	U.S.	1983-2005	+	
Karanassou and Sala (2013)	U.S.	1962-2009		-
Hogrefe and Kappler (2013)	18 countries	1980-2008		-
Stockhammer (2013)	71 countries	1970-2007	+	-

Note: LIS= Labor Income Share.

Regarding trade, the largest impacts are in Spain and Italy because these are the economies where globalization has had more influence in restricting wage growth. In particular, the growing exposure to international trade in Spain and Italy accounts, respectively, for 12.5 pp and 14.2 fall in their labor income shares in the last three decades. This is, by far, a much larger incidence than in France and the UK (2.3 and 3.2 pp), or Finland (5.6 pp).

[13]This result is consistent with the findings in Agnese and Sala (2011) according to which the fall in the labour share in Japan can be attributed to the changes that took place within the labour relations system, mainly the weakening of unions.

These results add to the growing literature on the effects of unions and international trade on the fall of the labor income share, which is briefly summarized in Table 5 (the positive sign of unions imply, as in our study, a positive connection between union power and the labor income share).

5 Conclusions

Conventional wisdom asserts that unemployment is caused by labor market rigidities and should be addressed through systematic institutional deregulation. This would allow wages to converge to their market-clearing level. In this paper, we divert from the standard approach and take a fresh look at the drivers of employment compensation. We estimate country-specific wage equations for a selection of eight OECD economies and take a medium- to long-term perspective to evaluate the main factors conditioning pay determination. Our selected economies are representative of the Anglo-Saxon, Nordic, and Continental European economic models, as they were classified in Daveri and Tabellini (2000), plus Japan.

We show that pay determination in last decades has been conditioned by three structural phenomena irrespective of the differences in these economic models.

The first one is productivity growth. It reflects efficiency progress and is a common factor in all economies. Our analysis uncovers a relevant fact: in the absence of productivity growth, real wages in all economies display and downward trend. This trend is relatively flat in the Anglo-Saxon and Nordic countries (US, UK, Finland, and Sweden), and relatively steep in Japan and Continental Europe (France, Italy, and Spain).

The second phenomena is the deunionization process, which has had especial incidence in Japan, followed by the Continental European countries (with the exception of Spain). The weakening of union power has had a much lower incidence in the Anglo-Saxon countries –which are the paradigm of deregulated markets–, and no relevant influence in the Nordic economies. This confirms the well-known result that union power is innocuous to the labor market in Scandinavia.

The third phenomena is trade. We find the impact of trade on wage setting simply irrelevant in closed economies such as the US and Japan. On the contrary, our counterfactual simulations show that trade prevented wages to increase in all European economies (with the exception of Sweden, where its impact is of minor magnitude). The effect of trade reaches its maximum in Italy and Spain, and reveals that labor costs have been the critical adjustment mechanism to the new economic conditions brought by the globalization process.

We also show that, by preventing real wages to rise further and, as a consequence, by enhancing the wage-productivity gap, deunionization and trade are significant contributors to the continuous fall in the labor income share

Our results align our work with those studies skeptical of the conventional wisdom, and call for reappraisal of the policy measures so often recommended from standard accounts of the unemployment problem.

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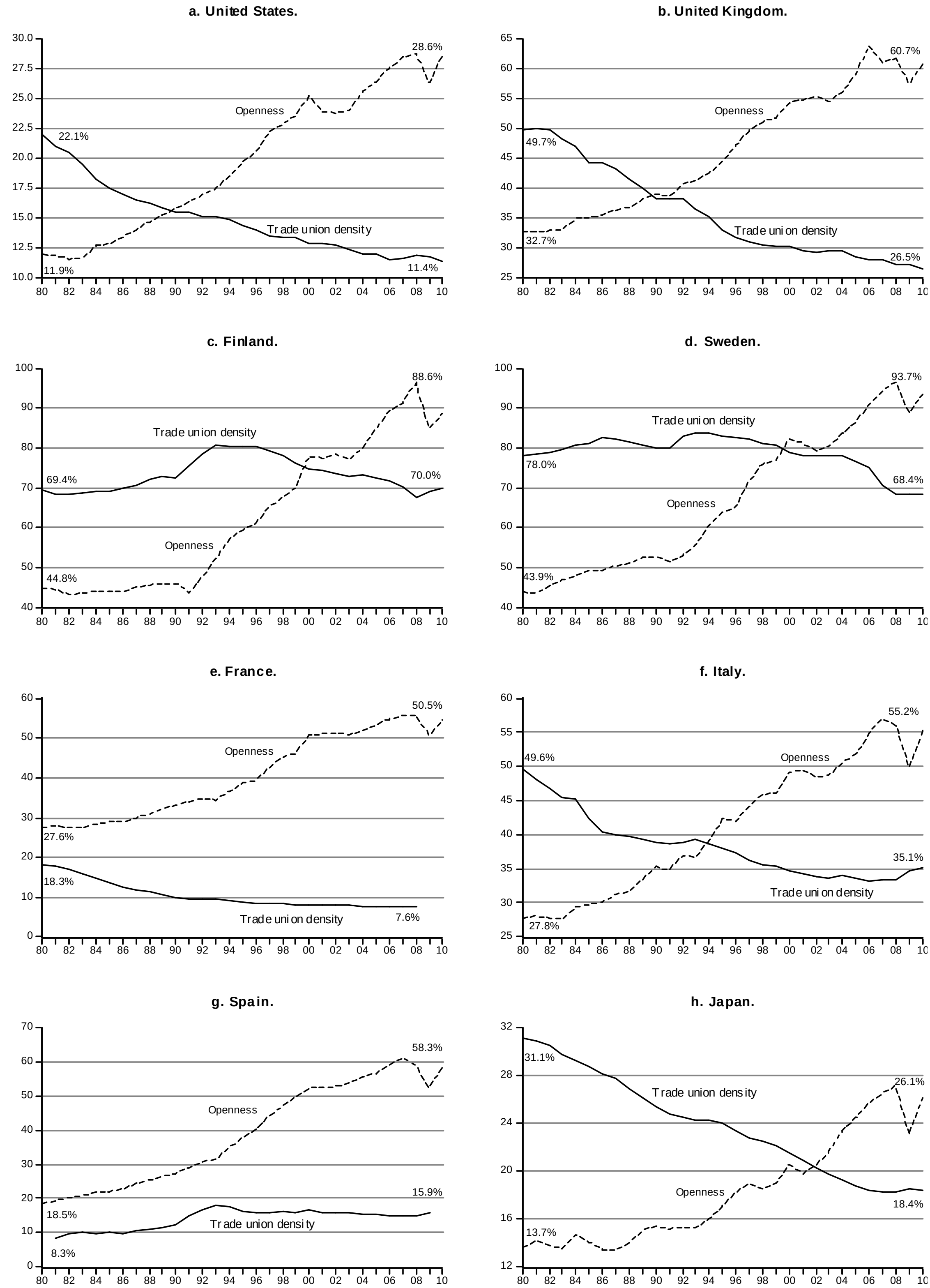
Appendix

Figure A1. Unions and international trade. 1980-2010.



Sources: OECD labor Market Indicators and OECD Economic Outlook no. 90.

Essay 2

The determinants of capital intensity in Japan and the U.S.

Abstract

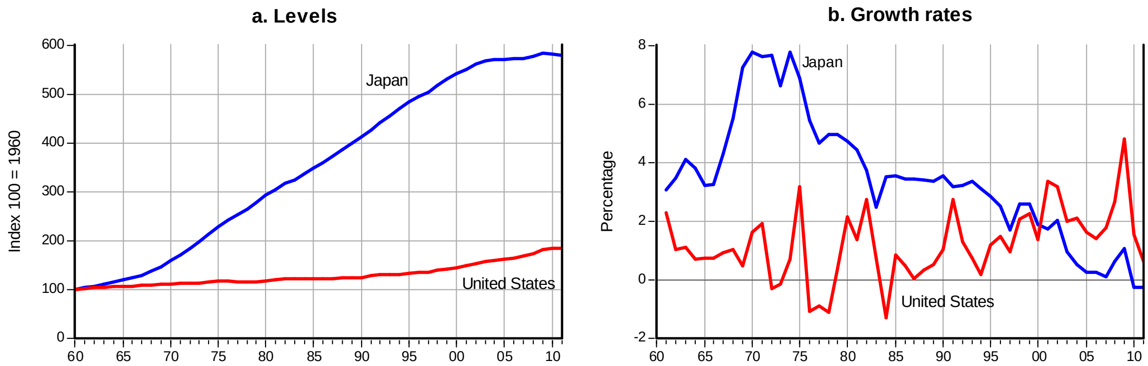
This paper develops a model of capital intensity determination that includes both supply and demand-side drivers. This model is estimated for two economies with different time paths of capital intensity, Japan and the United States, and both sets of determinants are shown to be relevant. We provide new estimates of the elasticity of substitution between production factors (around 0.30 for the U.S. and 0.85 for Japan). We also find a differenced nature of biased technical change in both countries, i.e. capital-saving in Japan and labor-saving in the U.S., which can help in explaining their diverse paths in the capital deepening process. We conduct a counterfactual analysis with dynamic simulations that verifies the role assignable to each determinant of capital intensity in the model. The results indicate that without the fall in the relative factor cost (supply-side driver), capital intensity would not have grown as much, in both countries, over the sample period (around 60% less in Japan, 16% less in the U.S.). In turn, without the fall experienced in relative factor utilization (demand-side driver), capital intensity would have grown much more in the U.S. (a whole 177% instead of the actual 66% growth). This role of demand-side forces is not detected in Japan's case. The simulation of the model without a time-trend to proxy technical change gives a higher trajectory of capital intensity in Japan and a lower one in the U.S., reinforcing the opposite-sign biased effect of technical change in each case. These results may help to enlighten the design of economic growth and labor market policies, which should not be the by-product of economic analysis based exclusively on supply-side considerations. Demand-side forces are, potentially, a crucial part of economic growth, as shown for the U.S. case, and their inclusion in the policy-making frame of thought may bring renewed answers for economic recovery.

1 Introduction

Although capital intensity, i.e. the ratio of capital stock over employment, plays a central role in economic growth models, it is generally considered as an input variable. No effort is devoted to the empirical assessment of its determinants in spite, for example, of the contrasted trajectory of capital intensity across countries, or in spite of the limitation that this imposes in growth accounting analysis.[14]

This paper intends to fill this void by providing evidence on the determinants of capital intensity in two economies with different institutional settings: Japan and the United States. As shown in Figure 1, the different time paths followed by the capital-per-worker ratio is itself calling for an empirical analysis of its causes.

Figure 1. Capital intensity.



Source: Ameco Database.

The process of capital intensity growth was specially intense in Japan, where the amount of capital stock per employee grew almost sixfold between 1960 and 2011. In contrast, it less than doubled in the US (Figure 1a). The origin of these differences lies in the very dynamic process of capital deepening linked to the industrialization process experienced by Japan in the 1960s and 1970s. After peaking in the first half of the 1970s, however, the growth rate of capital intensity has evolved around a steady downward path (Figure 1b). On the contrary, the process of capital deepening in the US accelerated from the mid 1980s until 2009. The Great Recession has then caused a sudden fall similar to those occurred in the aftermath of the oil prices shocks.

Departing from a standard analytical framework for the demand of production factors (Antràs, 2004; McAdam and Willman, 2013), we take a step further and relax the perfect competition assumption. When dealing with product demand uncertainty, firms adjust

[14] Madsen (2010), for example, points out that a problem associated with the traditional growth accounting framework is the lack of information about the factors responsible for the evolution of capital intensity.

their degree of factor utilization *ex post* once investment decisions are already made. In that context, capital intensity is driven by supply-side factors (i.e., factor costs and technology) as well as by demand-side conditions. The result is a model of capital intensity where the capital-per-worker ratio is explained by the relative factor cost, relative factor utilization, a time trend proxying technical change, and further controls, such as taxation and the degree of exposure to international trade (both shown relevant to this analysis by related literature).

Capital intensity, jointly with technological progress, are the main factors of economic growth, and so is acknowledged by the standard macroeconomic literature. In a recent contribution, for example, Takahashi *et al.* (2012) show that capital intensity has been crucial in Japan and other OECD economies' postwar growth. Very few efforts, however, have been assigned to the assessment of the determination of an economy's capital intensity.

In a Hecksher-Ohlin analytical background, Hasan *et al.* (2013) argue that labor and capital market regulations determine the industry-level capital stock per worker. They claim that restrictive labor laws can curb firms' ability to adjust their labor demand to shocks in demand, technology and trade. This may be seen as an indirect way of looking into capital intensity. In our model, we directly study the determinants of capital intensity.

In this way, our paper contributes to the literature in three main dimensions. First of all, in considering a model in which capital accumulation depends both on supply-side and demand-side factors. Second, in providing an empirical account of the determinants of capital intensity from a wider than usual perspective, with controls for demand-side constraints, tax variables and the degree of exposure to international trade. Third, in providing updated estimates of the elasticity of substitution between capital and labor and identifying the nature of factor-biased technical change.

The results show that the estimated coefficients associated to relative factor costs are statistically significant in all specifications in both countries. This is specially important, since those coefficients allow for the estimation of the elasticity of substitution between production factors. In turn, the significance of the proxies of demand-side channels is more scatter but still cannot be dismissed. Further controls, as direct taxation and openness to trade, also show a relevant role in decelerating capital intensity growth. The most striking feature of our results, however, is the different nature of the factor-biased technological change in Japan and the US. This provides a new explanation of the reasons why these economies have followed such different paths in their capital deepening process.

A counterfactual exercise of dynamic simulations disentangles the participation of each determinant in the evolution of capital intensity. The fall in relative factor costs over the sample period impulses capital per worker growth in both countries. In turn, demand-side forces are partly responsible for capital intensity growth only in the U.S., where without the decline in relative factor utilization since the 1970s the growth in capital per worker could have been steeper. When controlling for technical change, the path of capital intensity evolution is higher than the actual one in Japan and lower than the actual one in the U.S. This result points to technological progress as having a capital-saving effect in Japan, and a labor-saving effect in the U.S. Regarding openness to international trade and direct taxation, simulation results prove them responsible of a decelerating effect of capital intensity.

The rest of this paper is structured as follows. Section 2 presents the analytical framework. Section 3 deals with empirical issues related to the data and the estimated models. Section 4 computes the elasticities of substitution between capital and labor, and evaluates technological change in Japan and the US. Section 5 presents counterfactual simulations. Finally, section 6 concludes.

2 Analytical framework

As Antràs (2004) and McAdam and Willman (2013), we depart from a Constant Elasticity of Substitution (CES) production function from which the factor demand equations are first derived, and then combined into a single expression that accounts for the supply-side determinants of capital intensity. Then, along the lines of Andrés *et al.* (1990a, 1990b), Fagnart *et al.* (1999) and Bontempi *et al.* (2010), we consider the possibility of product demand uncertainty. In that context, when expected demand is not met by its actual value, firms are likely to react by adjusting their use of the production factors either by hiring or firing workers, by changing the rate of capacity utilization, or by using both mechanisms. In other words, the uncertainty on the actual level of product demand creates a transmission channel by which the demand-side conditions affect the investment and hiring/firing decisions of the firms. This explains why capital intensity is likely to depend on demand-side factors (on top of the supply-side ones), and justifies the extended model we present at the end of this Section.

2.1 Factor demands and capital intensity

Consider an economy with f identical firms that supply a homogeneous good. These firms acquire inputs in competitive markets and face a cost per unit of labor W, and a cost of capital use CC. Each firm has a CES production technology so that:

$$Y_t = \left[\theta(A_t^N N_t)^{-\beta} + (1-\theta)\left(A_t^K K_t\right)^{-\beta}\right]^{-1/\beta}, \tag{1}$$

where Y is output, N is employment, K is capital stock, A^N is an index of labor-augmenting efficiency (proxying Harrod-neutral technological change), and A^K is an index of capital-augmenting efficiency (proxying Solow-neutral technological change); parameter θ represents the factor share ($0 < \theta < 1$); $\sigma = \frac{1}{1+\beta}$ is the constant elasticity of substitution between capital and labor; and β denotes the degree of substitutability between both factors.

As standard (Antràs, 2004; León-Ledesma *et al.*, 2010), we assume that biased technological progress grows at constant rates denoted, respectively, by λ_N and λ_K. We thus have $A_t^N = A_0^N e^{\lambda_N \cdot t}$ and $A_t^K = A_0^K e^{\lambda_K \cdot t}$, where A_0^N and A_0^K are the initial values of the technological progress parameters, and t is a linear time trend. Note that $\lambda_N = \lambda_K > 0$ would imply Hicks-neutral technical progress, $\lambda_K > 0$ and $\lambda_N = 0$ implies Solow neutrality, $\lambda_N > 0$ and $\lambda_K = 0$ yields Harrod neutrality, and $\lambda_N, \lambda_K > 0$ but $\lambda_N \neq \lambda_K$ indicates factor-biased technical change.

Profit maximization in a perfectly competitive environment yields expressions for the factor demands (as a proportion of total output) that log-linearized can be written as

$$\log(K_t/Y_t) = \alpha_K - \sigma \log(CC_t/P_t) - (1-\sigma)\lambda_K t \qquad (2)$$

$$\log(N_t/Y_t) = \alpha_N - \sigma \log(W_t/P_t) - (1-\sigma)\lambda_N t, \qquad (3)$$

where P is the aggregate product market price; $\alpha_K = \sigma \log(1-\theta) + (\sigma - 1)\log A_0^K$ and $\alpha_N = \sigma \log \theta + (\sigma - 1)\log A_0^N$ are constant; and $1 - \sigma = \frac{\beta}{1+\beta}$. Subtraction of equation (3) from equation (2) yields the following specification for capital intensity:

$$\log(K_t/N_t) = \alpha - \sigma \log(CC_t/W_t) + (1-\sigma)(\lambda_N - \lambda_K)t. \qquad (4)$$

where $\alpha = \alpha_K - \alpha_N$.

Equation (4) is standard; it corresponds, for example, to equation (3') in Antràs (2004, p. 19) and equation (5) in McAdam and Willman (2013, p. 704). It explains that capital intensity depends on two supply-side factors: (i) the relative cost of labor and capital (there is capital intensity growth if the real wage grows relatively to the user cost of capital, making labor relatively more expensive); and (ii) the direction of factor-biased technical change (given that labor and capital are gross complements, i.e. $\sigma < 1$, then if labor-efficiency grows faster than capital-efficiency ($\lambda_N > \lambda_K$), there will be capital intensity growth).

2.2 Product demand uncertainty

Next, we relax the assumption of perfect competition and perfect information in the product market by assuming that firms hold some market power and are subject to random unexpected shocks. This implies that firms will now maximize profits based on an expectation of the stochastic demand (Y^E). Uncertainty about aggregate demand shapes firms' investment decisions (Fagnart *et al.*, 1999; Bond and Jenkinson, 2000; Bontempi *et al.*, 2010) and allows for the inclusion of demand-side considerations.

The sequence of decisions is as follows. Firms maximize profits subject to their expectation of demand in period t. In $t+1$, once the realization of the random (and unexpected) shocks that determine the demand are known, the utilization rate of installed capacity and the corresponding labor demand are adjusted accordingly.

Along the lines of Fagnart *et al.* (1999) firms use a putty-clay technology. Under a fixed productive capacity in the short-run, firms adjust the degree of factor utilization, and capital and labor are substitutes *ex ante*. *Ex post*, with capacity choices made and idiosyncratic shocks known, firms face the actual demand by adjusting the utilization intensity of the production factors, by hiring or firing workers, and by deciding on the capacity utilization rate. At this stage, production factors may be thought as complements to achieve a certain level of production. This model, therefore, allows for *ex post* rationing of factor utilization in contrast to the standard maximization problem.

The realization of the demand faced by firms depends on two factors: the price level (chosen by firms) and random shocks. The expected demand that firms consider in their profit maximization problem is the expected value of this realization Y_t^E:

$$Y_t^E = E_{t-1}\left[Y_t(P_t, \varphi)\right], \tag{5}$$

where E is the rational expectations operator, and φ is represents an idiosyncratic (stochastic) shock with zero mean and a constant standard deviation greater than zero. In other words, firms produce (and decide their factor demands) accordingly to their expectation of product demand, which is a function of the the aggregate product market price and the shocks.

Under these assumptions, the profit-maximization problem of the firm corresponds to a standard monopolistic competition case:

$$\begin{aligned}
\max\ \pi\left(K_t, N_t, P_t\right) &= P_t Y_t^E - W_t N_t - CC_t K_t \\
\text{subject to } &: \\
Y_t^E &= E_{t_{-1}}\left[\theta(A_0^N e^{\lambda_N . t} N_t)^{-\beta} + (1-\theta)\left(A_0^K e^{\lambda_K . t} K_t\right)^{-\beta}\right]^{-1/\beta}
\end{aligned}$$

where π stands for the firms' profit function.

Operating from the first order conditions of this problem, the optimal levels of factor utilization relative to output are obtained as an inverse relation with respect to each factor's cost:

$$\frac{K_t}{Y_t^E} = (1-\theta)^\sigma \left(\frac{CC_t}{P_t}\right)^{-\sigma} (A_0^K e^{\lambda_K . t})^{\frac{-\beta}{1+\beta}} \tag{6}$$

$$\frac{N_t}{Y_t^E} = \theta^\sigma \left(\frac{W_t}{P_t}\right)^{-\sigma} (A_0^N e^{\lambda_N . t})^{\frac{-\beta}{1+\beta}} \tag{7}$$

Note that log-linearization of equations (6) and (7) under perfect competition and perfect information would yield equations (2) and (3).

Through aggregation of the f firms, the overall expected demand can be replaced by the potential aggregate demand level ($\hat{Y}$). Further addition of the ratio $\frac{Y_t}{Y_t}(=1)$ to the left-hand side of both equations then yields:

$$\frac{K_t}{Y_t} = (1-\theta)^\sigma \left(\frac{CC_t}{P_t}\right)^{-\sigma} (A_0^K e^{\lambda_K . t})^{\frac{-\beta}{1+\beta}} \frac{\widehat{Y}_t}{Y_t} \tag{8}$$

$$\frac{N_t}{Y_t} = \theta^\sigma \left(\frac{W_t}{P_t}\right)^{-\sigma} (A_0^N e^{\lambda_N . t})^{\frac{-\beta}{1+\beta}} \frac{\widehat{Y}_t}{Y_t} \tag{9}$$

where the ratio $\frac{\widehat{Y}_t}{Y_t}$ expresses the gap between potential aggregate demand ($\hat{Y}$) and the actual level of aggregate production (Y), once factor demands have been adjusted *ex post*.

2.3 Mind the gap

The $\frac{\widehat{Y}_t}{Y_t}$ ratio is the transmission channel for business cycle effects (Nakajima 2005, Fagnart *et al.* 1999). As such, it is directly related to the gap between total installed production capacity and the rate of capacity utilization of the production factors. Because the later is observable (in contrast to $\widehat{Y}_t$), we assume that the degree of factor utilization can be empirically used as a proxy for the ratio $\frac{\widehat{Y}_t}{Y_t}$. Although this is relatively standard in the literature, we believe that equations (8) and (9) require a specific proxy accounting, as close as possible, for the demand-pressures that capital and labor will experience.

The standard procedure, however, consists in looking at only one of the production factors' side (see, for example, Andrés *et al.* 1990a). The reasoning is the following. Firms invest in capacity with an expectation of potential demand but, because they end up using this capacity only with the intensity required by actual demand (which is called *ex post* rationing), the degree of capacity utilization (or capacity utilization rate) is determined accordingly. Because of the correspondence between the fact that production is responsive to aggregate demand *ex post,* and installed capacity is rigid in the short run, we follow this route regarding equation (8) and proxy the unobservable $\frac{\hat{Y}_t}{Y_t}$ ratio through the standard capacity utilization rate (CUR) variable[15].

With respect to the extent to which the other production factor, labor, is used relative to its potential, we consider the employment rate, which reflects the actual use of the labor factor (employment, N) over its total possible use (working-age population, Z). It follows that, for equation (9), we account for the degree of labor usage through the employment rate ($NR = N/Z$).

In other words, we re-write the factor demand equations as:

$$\frac{K_t}{Y_t} = (1-\theta)^{\sigma} \left(\frac{CC_t}{P_t}\right)^{-\sigma} (A_0^K e^{\lambda_K . t})^{\frac{-\beta}{1+\beta}} . h(CUR_t) \tag{10}$$

$$\frac{N_t}{Y_t} = \theta^{\sigma} \left(\frac{W_t}{P_t}\right)^{-\sigma} (A_0^N e^{\lambda_N . t})^{\frac{-\beta}{1+\beta}} . h(NR_t), \tag{11}$$

where $h(\cdot)$ is a monotonically increasing function as in Andrés *et al.* (1990a, p. 88).

Log-linearization of equations (10) and (11), and substraction of the second one from the first one yields an expression for capital intensity and its determinants:

$$\begin{aligned} \log\left(\frac{K_t}{N_t}\right) = {} & \alpha - \sigma\left[\log\left(\frac{CC_t}{P_t}\right) - \log\left(\frac{W_t}{P_t}\right)\right] + (1-\sigma)(\lambda_N - \lambda_K)\, t \\ & + (\gamma_K - \gamma_N)\left[\log(CUR_t) - \log(NR_t)\right]. \end{aligned} \tag{12}$$

Note that the only difference with respect to equation (4) is the last term, which results from the assumption of an stochastic behavior of aggregate product demand allowing for *ex post* rationing of factor utilization[16].

2.4 Factor-biased technical change

Regarding the second term in the right-hand side of equation (12), we follow McAdam and Willman (2013), León-Ledesma *et al.* (2010, 2013) and Antràs (2004), among others, in applying empirical efforts to identify and measure biased technical change[17]. The evidence

[15] As already said above, this is relatively standard in the literature. For example, Graff and Sturm (2012) associate the degree of capacity utilisation directly to the output gap and Planas *et al.* (2013) to the cycle of total factor productivity.

[16] We assume that the log-linearization of $h(\cdot)$ yields a linear function of the logs of CUR and NR in each case, as presented in (12).

[17] To cite McAdam and Willman (2013, p.698): "... despite renewed interest in models of biased technical change, the corresponding empirical effort to identify (i.e., measure) episodes from macro data has been lacking."

in this strand of the literature shows that making *a priori* assumptions about the form of technical progress (e.g. assuming Hicks neutrality) is likely to misguide the insights on the effect of technical progress on capital intensity.

The dominant assumption of a balanced growth path (BGP) in theoretical growth literature implies a framework where the main macro variables converge to a common growth rate, the underlying ratios (factor income shares and factor-GDP ratios) remain constant [as described by Kaldor (1961)], and technical change is solely labor augmenting (i.e. Harrod neutral). Acemoglu (2003) and McAdam and Willman (2013) suggest that although technical progress is labor-augmenting along the BGP, it can be capital-augmenting in medium-run transitions away from the BGP. With an elasticity of substitution between labor and capital lower than one (gross complements), this pattern allows for long-run asymptotic stability of factor shares, and also non-stationary evolution in the medium-run (which we observe in reality).

Specifically, the model in this Section shows how given that labor and capital are gross complements (i.e., the elasticity of substitution is lower than one[18]), a relatively higher growth of labor-augmenting technical change favors capital intensity, as shown in terms of equation (13)[19].

$$\frac{\partial(K/N)}{\partial(A^N/A^K)} > 0 \quad \text{if} \quad \sigma < 1 \tag{13}$$

In our model, along the lines of Acemoglu (2003), Antràs (2004) and McAdam and Willman (2013), capital accumulation in the long-run implies that labor-efficiency has to grow faster than capital-efficiency. Recall that in (4) and (12), with $\sigma < 1$, higher capital intensity is reached when $\lambda_N > \lambda_K$.

3 Empirical issues

3.1 Estimated models

Consideration of demand-side drivers in factor demand equations is still an open issue. This is the reason why, on top of the relative degree of factor utilization $\log(CUR_t) - \log(NR_t)$, we follow Añón-Higón (2007), and consider the variation in worked hours per employee as an alternative aggregate proxy of demand-side pressures (we take the growth

[18] There a several recent studies that find $\sigma < 1$. For example, Antràs (2004), Chirinko (2008), Chirinko *et al.* (2011), León-Ledesma *et al.* (2010), Klump *et al.* (2012) and McAdam and Willman (2013).

[19] McAdam and Willman (2013, p. 703).

rate because this proxies the business cycle in terms of varying demand-side pressures). The reasoning behind this choice is that worked hours per employee reflect simultaneously an increase in the usage intensity in capital stock and labor. Moreover, the average annual amount of hours worked per employee is likely to avoid the endogeneity problems that would entail considering variations in output (since the dependent variable is indeed made of capital and labor), which is the natural alternative in the literature.

Two additional sets of control variables are related to capital intensity: the degree of exposure to international trade and the fiscal system. Since capital intensity, the degree of substitution between capital and labor, and globalization are deeply intertwined (Hutchinson and Persyn, 2012), inclusion of the degree of trade openness (op) is a must. Regarding the fiscal system, a key variable for firm's decisions is direct taxes on business, which is crucial in defining, for example, investment decisions. This has been studied in Bond and Jenkinson (2000), Edgerton (2010) and Madsen (2010), where the decelerating effect of corporate taxation on capital deepening is explained as a disincentive to firm-level investment. Beyond that, because we have confronted the user cost of capital with the labor cost, and the degree of capital stock utilization with the employment rate, we confront direct taxes on business $\left(\tau^b\right)$ to direct taxes on households $\left(\tau^h\right)$ to capture, if any, the specific impact of taxes on each demand factor. Of course, payroll taxes is another crucial element of the tax system, but its relevance is more related to the wage bargaining process between firms and workers. Since this is implicitly taken into account through the wage variable (total compensation, which includes social security contributions) we do not need any further control.

Overall, we estimate two empirical versions of equation (12). Model 1 will be a straightforward augmented version of equation (12):

$$kn_t = \beta_0 + \beta_1(cc_t - w_t) + \beta_2(cur_t - nr_t) + \beta_3 t + \beta_4 op_t + \beta_5\tau^b + \beta_6\tau^h + u_{1t}, \qquad (14)$$

where $kn_t = \log(K_t/N_t)$, $cc_t = \log(CC_t/P_t)$, $w_t = \log(W_t/P_t)$, $cur_t = \log(CUR_t)$, $nr_t = \log(NR_t)$ and u_{1t} represents a standard error term with zero mean and constant standard deviation. Detailed definitions of the additional controls, op, τ^b,and τ^h (and also of the rest of the variables) are given in Table 1.

Model 2 substitutes relative factor utilization, $cur_t - nr_t$, by the change in worked hours per employee (Δhr):

$$kn_t = \gamma_0 + \gamma_1(cc_t - w_t) + \gamma_2\Delta hr_{t-1} + \gamma_3 t + \gamma_4 op_t + \gamma_5\tau^b + \gamma_6\tau^h + u_{2t}, \qquad (15)$$

where u_{2t} represents a standard error term with zero mean and constant standard deviation. Note that the coefficient on hours is lagged once to help avoiding endogeneity problems. In contrast, the term capturing demand-side pressures in equation (14) is not lagged to maintain coherence with respect to the theoretical model. We have assumed a putty-clay technology and argued that short-run capital stock adjustments take place through changes in the degree of capacity utilization. This implies that demand changes foreseen in $t-1$ are accommodated through changes in investment, not through changes cur which can only respond in period t.

A crucial remark is that these empirical models are estimated as dynamic equations to take into account the adjustment costs potentially surrounding all variables involved in the analysis (endogenous and exogenous). The lagged structure of the estimated relationships is therefore a pure empirical matter.

The coefficients β_1/γ_1 are associated to the relative cost of production factors, and a negative sign is expected. As the wedge between the cost of factors $(cc-w)$ increases, capital becomes relatively more costly than labor, and a deceleration of capital intensity growth is expected.[20] The crucial feature of this coefficient is its correspondence with the constant elasticity of substitution between capital and labor (σ).

The coefficients β_2/γ_2 are associated to the role of demand-side pressures, and a positive sign is expected. A rise in the wedge between the relative intensity in factor utilization $(cur-nr)$ implies that tightness in the capital side is larger than in the labor side. Firms, therefore, are expected to react by investing more intensively than embarking in new hirings. As a consequence, capital intensity is expected to accelerate.

Firm's decisions to expand capacity through investment are based to a large extent on their assessment of their future sales, which we assumed to be uncertain. Managers are naturally cautious about overestimating future sales, as the penalty for doing so tends to be much greater than for losing potential business by failing to expand (Smith, 1996). In our model, the capacity utilization rate is a proxy for the perception of the firm of the economic reality, which reflects on its expectations on aggregate demand. Since the expansion of capacity drives investment, we expect a positive effect on capital intensity when the wedge between a higher degree of capacity utilization rate and a higher employment rate widens.

Given the assumption of constant rates of technical progress, the coefficients $\beta_3/\gamma_3 = (1-\sigma)(\lambda_N - \lambda_K)$ measure an asymmetric progress in the efficiency of each production factor. If $\hat{\beta}_3/\hat{\gamma}_3 > 0$ and $\hat{\sigma} < 1$, there is evidence that labor-augmenting efficiency grows faster than capital-augmenting efficiency (the same holds in case of opposite signs in both estimates). If, on the contrary, the $\hat{\beta}_3/\hat{\gamma}_3 > 0$ are positive and $\hat{\sigma} > 1$, the conclusion is

[20] Decisions to invest in new capacity are influenced by the cost and availability of capital and the target rates of return sought by firms and financial institutions. The dependence on bank loans is an important factor limiting expansion and the user cost of capital is a crucial factor in the expected net return to investment by firms.

that capital-augmenting efficiency grows faster than labor-augmenting efficiency. In both cases, therefore, there is evidence of biased technological change, something that in the standard Cobb-Douglas framework, where $\hat{\sigma} = 1$, cannot be measured.

3.2 Data

We use annual data obtained from various sources. From the European Commission's Ameco database we take net long-time series on net capital stock[21]. Data on the capacity utilitzation rate is obtained from Ministry of Economy, Trade and Industry for Japan, and from the Board of Governors of the Federal Reserve System for the U.S. The rest of the variables is gathered from the OECD Economic Outlook 91 (2012).

Table 1. Definitions of variables.

k	real net capital stock	p	GDP deflator
n	employment	p^i	investment deflator
kn	capital intensity $(= k - n)$	δ	depreciation rate
z	working-age population	i	nominal interest rate
nr	employment rate $(= n - z)$	cc	real user cost of capital
cur	capacity utilization rate		$= \frac{p^i}{p} \left(i + \delta - \Delta p^i \right)$
w	real compensation per employee	T	direct taxes on business
hr	hours of work per employee	τ^b	direct taxes on business (TB) as % GDP
Y	GDP		$= \log \left(TB/Y \right)$
X	exports of goods and services	τ^h	direct taxes on households (TH) as % GDP
M	imports of goods and services		$= \log \left(TH/Y \right)$
op	trade openness	t	linear time trend
	$= \log \left(\left[X + M \right] / Y \right)$	Δ	difference operator
c	constant		

Note: All variables used in the econometric analysis are expressed in logs.

Table 1 provides the concrete definitions of the empirical variables used. All of them are standard and the only clarification refers to the definition of the user cost of capital, which is constructed as in Andrés *et al.* (1990a) assuming a constant depreciation rate equal to 0.1. McAdam and Willman (2013) argue that the observed government bond rate

[21] The net capital stock at constant prices, total economy (OKND), is calculated as $OKNDt = OKND_{t-1} + [OIGTt - (UKCTt : PIGTt) * 100]$, where OIGT = Gross fixed capital formation at constant prices; total economy, UKCT = Consumption of fixed capital at current prices; total economy and PIGT = Price deflator gross fixed capital formation; total economy.

OIGT = Gross fixed capital formation at constant prices; construction + equipment + products of agriculture, forestry, fisheries and aquaculture + other procucts.

does not correctly capture firms' marginal financing costs in an imperfectly functioning financial market. The real observed user cost of capital would have been above the indicated. Along these lines, we follow the deifinition in Andrés et al. (1990a) detailed in Table 1.Note, also, that all variables will be used in logs so as to allow an unambiguous interpretation of the estimated coefficients as elasticities.

3.3 Estimation procedure

We deal with long time-series data and we need to ensure that the long-run estimated relationships between capital intensity and its determinants are non-spurious. Of course, if k, n, cc, w, cur, z, T, and Y behaved certainly as I(1) variables we could argue, since we work with these variables by ratios ($kn, cc - w, cur - nr, (X + M)/Y$ and T/Y), that we end up dealing with I(0) variables and cointegration issues are of no concern. Nonetheless, we have indeed conducted the standard unit root tests and we are able to stress that, yes, we are indeed dealing with I(1) variables (see Table A1 in the Appendix for test results)[22].

Nevertheless, we proceed as if we were unsure of the degree of integration of our variables. This is why our estimation is conducted following the bounds testing approach, or ARDL (AutoRegressive Distributed Lag) approach, which yields consistent short- and long-run estimates irrespective of whether the regressors are I(1) or I(0). This approach, which was developed by Pesaran and Shin (1999) and Pesaran, Shin and Smith (2001), provides an alternative econometric tool to the standard Johansen maximum likelihood, and the Phillips-Hansen semi-parametric fully-modified Ordinary Least Squares (OLS) procedures. The main advantage of the bounds testing approach is the possibility of avoiding the pretesting problem implicit in the standard cointegration techniques. It also yields consistent long-run estimates of the equation parameters even for small size samples and under potential endogeneity of some of the regressors (see Harris and Sollis, 2003).

We first estimate our models by OLS. Then, to make sure that we have indeed obtained non-spurious relationships between potential non-stationary variables, we verified that the residuals resulting from our estimated models are indeed stationary (see Table 4 below).

Regarding the selection of the estimated models, we first select equations that are dynamically stable and satisfy the conditions of linearity, structural stability, no serial correlation, homoscedasticity, and normality of the residuals. Then, among the models that meet these requirements, we select the dynamic specification of each equation by relying on the optimal lag-length algorithm of the Schwartz information criterion (Table A2 in the Appendix shows that these standard diagnostic tests are all passed at conventional significance levels).

[22] In the case of Model 2, Δhrs is, as expected, stationary.

Finally, we estimate the selected specifications by Two Stages Least Squares (TSLS) so as to control for potential endogeneity biases in the estimated effect of the relative factor costs $(cc - w)$, in relative factor utilization $(cur - nr)$ or hours, and in direct taxes on business. The instruments are statistically significant and we find the OLS and the TSLS results to be relatively alike, thus supporting the robustness of the results (only model 2 for Japan rejects by a short margin the Durbin-Wu-Hausman test of exogeneity).

4 Results

4.1 Estimated equations

We present the estimation results for equations (14) and (15) in Table 2, for Japan, and in Table 3, for the US.

Japan's estimation includes three dummy variables d^{9102}, d^{83}, and d^{97}, which take value one in the years indicated by the superscript. They account for the lost decade, and specific events such as the East-Asian crisis. They do not have any special economic relevance, but they help to achieve better results in terms of the misspecification tests (see Table A1).

All estimated equations have a high estimated coefficient associated to the first lag of capital intensity. This high persistence in the dynamic process is to be expected since productive capacity is not easily changed in the short run. The results for Japan show that relative factor costs present a highly statistically significant negative sign coefficient -as expected- in both models. Moreover, the estimated short-run coefficient is quite homogeneous in value. Regarding the proxies for demand-side channels to capital deepening, the cycle component of average hours worked (model 2) has greater statistical significance. Both direct taxes on businesses and households show a decelerating effect on capital intensity in the two models, as also does the degree of openness to international trade. Finally, the estimated coefficient associated to the time trend is negative, which combined with a lower-than-one elasticity of substitution would indicate that capital-associated efficiency grows at a higher rate than labor-associated efficiency (see section 4.2).

As for the U.S., again the coefficients associated to relative factor cost are negative and significant. In this case, now not only Δh_{t-1} presents statistical significance, but also the relative factor utilization. Direct taxes and openness are also detrimental for capital intensity in the U.S., but the latter enters the equations as the variation between two years, indicating that this is more a conjunctural downward pressure than a long-run effect. The estimated coefficient associated to the time trend is positive, and considering that the estimated elasticity of substitution for the U.S. is lower than the unity, that would imply labor-saving biased technical change as labor-efficiency grows faster than capital-efficiency (see section 4.2). Note that, according to these results, the U.S. and Japan would have opposite-direction bias in technological change between both production factors.

Table 2. Japan, 1980-2011.

	Model 1			**Model 2***	
	OLS	TSLS		OLS	TSLS
c	0.005 [0.983]	0.084 [0.759]	c	-0.020 [0.907]	-0.008 [0.969]
kn_{t-1}	0.964 [0.000]	0.954 [0.000]	kn_{t-1}	0.966 [0.000]	0.960 [0.000]
$cc_t - w_t$	-0.033 [0.000]	-0.038 [0.001]	$cc_t - w_t$	-0.034 [0.000]	-0.036 [0.062]
$\Delta\left(cc_t - w_t\right)$	0.020 [0.001]	0.019 [0.003]	$\Delta\left(cc_t - w_t\right)$	0.021 [0.000]	0.031 [0.030]
$\Delta\left(cc_{t-1} - w_{t-1}\right)$	0.015 [0.002]	0.016 [0.005]	$\Delta\left(cc_{t-1} - w_{t-1}\right)$	0.021 [0.001]	0.025 [0.004]
$cur_t - nr_t$	0.005 [0.625]	0.003 [0.813]	Δhr_{t-1}	0.076 [0.106]	0.111 [0.070]
$\Delta\left(cur_t - nr_t\right)$	0.015 [0.188]	0.019 [0.175]			
$\Delta\tau_t^b$	-0.015 [0.001]	-0.014 [0.012]	$\Delta\tau_t^b$	-0.014 [0.001]	-0.007 [0.573]
$\Delta\tau_{t-1}^h$	-0.007 [0.329]	-0.006 [0.426]	$\Delta\tau_{t-1}^h$	-0.009 [0.149]	-0.010 [0.200]
op_t	-0.030 [0.003]	-0.035 [0.008]	op_t	-0.035 [0.001]	-0.056 [0.002]
D^{9102}	0.005 [0.005]	0.005 [0.024]	D^{9102}	0.004 [0.005]	0.003 [0.039]
D^{83}	-0.014 [0.000]	-0.013 [0.000]	D^{83}	-0.014 [0.000]	-0.014 [0.000]
D^{97}	-0.010 [0.000]	-0.011 [0.000]	D^{97}	-0.010 [0.000]	-0.009 [0.034]
t	-0.001 [0.030]	-0.001 [0.145]	t	-0.001 [0.016]	-0.0003 [0.605]
LL	167.6			167.33	
Obs	32	32		32	32

Notes: LL = Log-likelihood; p-values in brackets; Instruments: kn_{t-1} cc_{t-1} w_{t-1} Δcc_{t-1} Δw_{t-1} cur_{t-1} nr_{t-1} op_{t-1} $\Delta\tau_{t-1}^h$ $\Delta\tau_t^b$ $\Delta\tau_{t-1}^b$ D_{9102} D_{83} D_{97} t Δhr_{t-1} Δhr_{t-2}.

Durbin-Wu-Hausman test [prob]: Model 1 [0.95]; Model 2 [0.04].

Table 3. US, 1970-2011.

Model 1	OLS	TSLS	Model 2*	OLS	TSLS
c	0.345 [0.401]	0.327 [0.516]	c	0.136 [0.772]	0.286 [0.582]
kn_{t-1}	0.951 [0.000]	0.941 [0.000]	kn_{t-1}	0.973 [0.000]	0.965 [0.000]
Δkn_{t-1}	0.292 [0.015]	0.308 [0.038]	Δkn_{t-1}	0.215 [0.241]	0.288 [0.176]
$cc_t - w_t$	−0.010 [0.094]	−0.016 [0.132]	$cc_t - w_t$	−0.013 [0.058]	−0.011 [0.443]
$cur_t - nr_t$	0.083 [0.139]	0.126 [0.192]	Δhr_{t-1}	−0.310 [0.317]	−0.204 [0.572]
$\Delta (cur_t - nr_t)$	−0.182 [0.000]	−0.170 [0.001]	Δhr_{t-2}	0.526 [0.027]	0.547 [0.026]
τ_t^b	−0.019 [0.069]	−0.035 [0.114]	τ_t^b	−0.018 [0.115]	−0.008 [0.699]
τ_t^h	−0.002 [0.899]	0.003 [0.862]	τ_t^h	0.014 [0.366]	0.011 [0.535]
Δop_t	−0.099 [0.024]	−0.069 [0.382]	Δop_t	−0.149 [0.002]	−0.206 [0.002]
t	0.001 [0.079]	0.001 [0.191]	t	0.0002 [0.655]	0.0004 [0.492]
LL	161.2			155.5	
$Obs.$	42	42		42	42

Notes: LL = Log-likelihood; p-values in brackets; Instruments: kn_{t-1} Δkn_{t-1} cc_{t-1} w_{t-1} cur_{t-1} nr_{t-1} Δcur_{t-1} Δnr_{t-1} τ_t^b τ_{t-1}^b τ_t^h τ_{t-1}^h Δop_{t-1} t Δhr_{t-1} Δhr_{t-2}

Durbin-Wu-Hausman test [prob]: Model 1 [0.92]; Model 2 [0.73].

Furthermore, regarding the validity of the long-run relationships between variables, the Augmented Dickey-Fuller test (ADF, with null hypothesis of non-stationarity) and the Kwiatkowski-Phillips-Schmidt-Shin test (KPSS, with null hypothesis of stationarity) were performed on the residuals of the estimated equations presented above. The results can be seen in Table 4. They largely reject in all cases the hypothesis of the existence of a unit root in the ADF test, and also do not reject stationarity in the KPSS test, reinforcing the stationarity of the residuals and therefore that we have obtained non-spurious relationships[23].

[23] At the 5% level, the residuals for Japan's model 2 estimation by TSLS may not be strictly stationary according to the KPSS. Nevertheless, the ADF test for those same residuals largely rejects the unit root hypothesis at the 1% level.

Table 4. Unit root tests on the residuals of equations (14) and (15)

	ADF test				KPSS test			
	Model 1 (u_{1t})		Model 2 (u_{2t})		Model 1 (u_{1t})		Model 2 (u_{2t})	
	OLS	TSLS	OLS	TSLS	OLS	TSLS	OLS	TSLS
Japan	-5.84	-5.32	-5.47	-6.61	0.054	0.045	0.447	0.500
U.S.	-4.17	-6.18	-6.41	-6.89	0.054	0.042	0.081	0.083

Note: ADF test critical value is -3.60 at the 1% level.

KPSS test critical values are 0.739 at the 1% level, and 0.463 at the 5% level.

4.2 Elasticities of substitution and technological change

Directed technical change is a consequence of a production factor becoming relatively more scarce, more expensive, or both. Innovation is then directed towards technologies that would save on the relatively more expensive factor. The bias in technical change may have a saving effect on one factor and an augmenting effect on the other one. The degree of substitutability between labor and capital is closely related to this phenomenon. Then, the elasticity of substitution between labor and capital and the rate of factor-biased technical change are key variables in economic growth models, specially in medium-run dynamics (McAdam and Willman, 2013).

Table 5 shows the elasticity of substitution between factors implied by our empirical models, together with the long-run impact on capital intensity of the constant rate of technological progress $\left(\varepsilon^{LR}_{kn-trend}\right)$. Given that the estimated models are dynamic, the elasticity of substitution is computed as the long-run elasticity of kn with respect to $(cc-w)$. Taking the example of Japan using Model 1, we have $0.038/(1-0.954) = 0.83 = \hat{\sigma}$. In turn, $\varepsilon^{LR}_{kn-trend} = (-0.001/(1-0.954)) * 100 = -2.2\%$. These two values are used to compute the implied rate of biased technological change following equations (4) or (12).

More precisely, in case of Model's 1 estimates for Japan, we use $\hat{\sigma} = 0.83$ and $\varepsilon^{LR}_{kn-trend} = -2.2\%$ to compute the value of $(\lambda_K - \lambda_N)$ using:

$$\begin{aligned} -2.2\% &= (1-0.83)(\lambda_N - \lambda_K) \\ &\Longrightarrow \quad (\lambda_K - \lambda_N) = 12.5\% \end{aligned}$$

This result implies that there is factor-biased technical change in Japan ($\lambda_K - \lambda_N > 0$, that is, $\lambda_K \neq \lambda_N$). In this case, the direction is capital saving.

Table 5 shows the calculations for both countries using the instrumental variables estimation of Models 1 and 2.

Table 5. Elasticities of substitution and technological change.

	Model 1				**Model 2**			
			Technical progress				Technical progress	
	$\hat{\sigma}$	$\varepsilon^{LR}_{kn-trend}$	Type	Rate	$\hat{\sigma}$	$\varepsilon^{LR}_{kn-trend}$	Type	Rate
Japan	0.83	-2.2%	Capital saving	12.5%	0.90	-0.7%	Capital saving	7.5%
US	0.27	1.7%	Labor saving	2.3%	0.31	1.1%	Labor saving	1.7%

Notes: $\varepsilon^{LR}_{kd-trend}$ denotes the long-run elasticity of capital deepening with respect to constant technical change; technical change is capital saving whenever $\lambda_K > \lambda_N$ and labor saving whenever $\lambda_N > \lambda_K$.

As noted, we find the elasticity of substitution between capital and labor to be below 1 in Japan. This value is larger than in other studies, which place it between 0.2 and 0.4 (Rowthorn 1999, Klump *et al.* 2012). However, neither the sample period nor the methodology is common to the one followed here.

We find the long-run impact of technological change to be between -0.7% and -2.2%. This implies that a rise in the rate of technological progress is translated, in the long-run and *ceteris paribus*, to a fall in capital intensity. In the context of our model, technological change in Japan is a long-run decelerator of the evolution of capital per worker. This result is critical to understand the deceleration in the process of capital deepening experienced by Japan since the mid 1970s. Together with the estimated elasticity of substitution, it provides evidence of a substantial bias in technological change, which is capital saving, and evolves at a rate between 7.5% and 12.5%. The capital-saving effect comes from the fact that a higher rate of capital related efficiency growth (i.e., $\lambda_K > \lambda_N$) reduces the pace of capital stock growth. This is consistent with the path followed by the process of capital deepening in Japan, with a huge increase in capital accumulation in the expansionary decades of 1960 and 1970, and a steep and continuous decrease in the 1980s, 1990s and 2000s. On this account, let us recall that our sample period for Japan starts in 1980. Not only this prevents us to have noise from the structural break occurred in the Japanese economic growth model, but it also allows us to capture more precisely this extraordinary long period of continuous deterioration in the ratio of capital stock to employment.

Regarding the US, our analysis yields an elasticity of substitution between capital and labor around 0.3 in the US, a value in the lower range of the estimates provided by the literature. In particular, although Chirinko (2008) finds a σ between 0.4 and 0.6, and León-Ledesma *et al.* (2010) and Klump *et al.* (2007) present values in the 0.5-0.7 range, it is important to emphasize Chirinko's *et al.* (2011) indication that the use of time series data at annual frequencies may lower the estimation of σ.[24] Chirinko *et al.* (1999), for example, provide an estimation of the elasticity of substitution rather low for the US of around 0.25.

We also find a long-run impact of technological change on capital intensity between 1% and 2%. This implies that a rise of 1 percentage point in the rate of technological progress is translated, in the long-run, in at least a 1% increase in capital intensity. In terms of biased technological change, we find consistent evidence (since Models 1 and 2 yield similar results) of a labor-saving bias (i.e., labor-related efficiency grows at a faster rate in the U.S., $\lambda_K < \lambda_N$). This may contribute to explain the secular process of industrial firms' delocalization of the US economy including the growing relevance of phenomena such as offshoring and outsourcing. This is however, barely a novelty, since our estimates of biased technological change in the US, between 1.7% and 2.5%, are fully aligned with those supplied by the literature [see the survey of Klump *et al.* (2012), Table 1, where rage of values from many papers is placed between 0.27% and 2.2%, although Antràs (2004) places it slightly above 3%].

5 Simulations

We now use our estimated Model 1 to perform dynamic accounting simulations. We simulate the model under two scenarios: first, a baseline scenario in which all exogenous variables take their actual values, and second, a scenario in which each one of the exogenous variables is kept constant at its value in the beginning of the sample period (1980 for Japan, 1970 for the U.S.). We call this a counterfactual simulation because the difference between the fitted values of capital intensity obtained from each one of those scenarios reveals how much of the actual capital intensity growth can be explained by the factor kept constant in the simulation.

We do not claim that the fitted values from the second scenario are the true values that capital intensity would have taken if one determinant was kept constant, they are

[24] Chirinko *et al.* (2011) argue that time series variations of investment spending largely reflect adjustments to transitory shocks and, in turn, firms usually respond to permanent variation. Hence, an elasticity estimated with time series data will tend to be lower than the "true" long-run elasticity.

only auxiliary to this accounting exercise. Figure 2 plots the results. The scale in all graphs is based on a 100 index for the first year of the sample period. To evaluate these results it is important to take into account the evolution of the exogenous variables (see Figures A1 and A2 in the Appendix).

Figure 2: Simulation results.

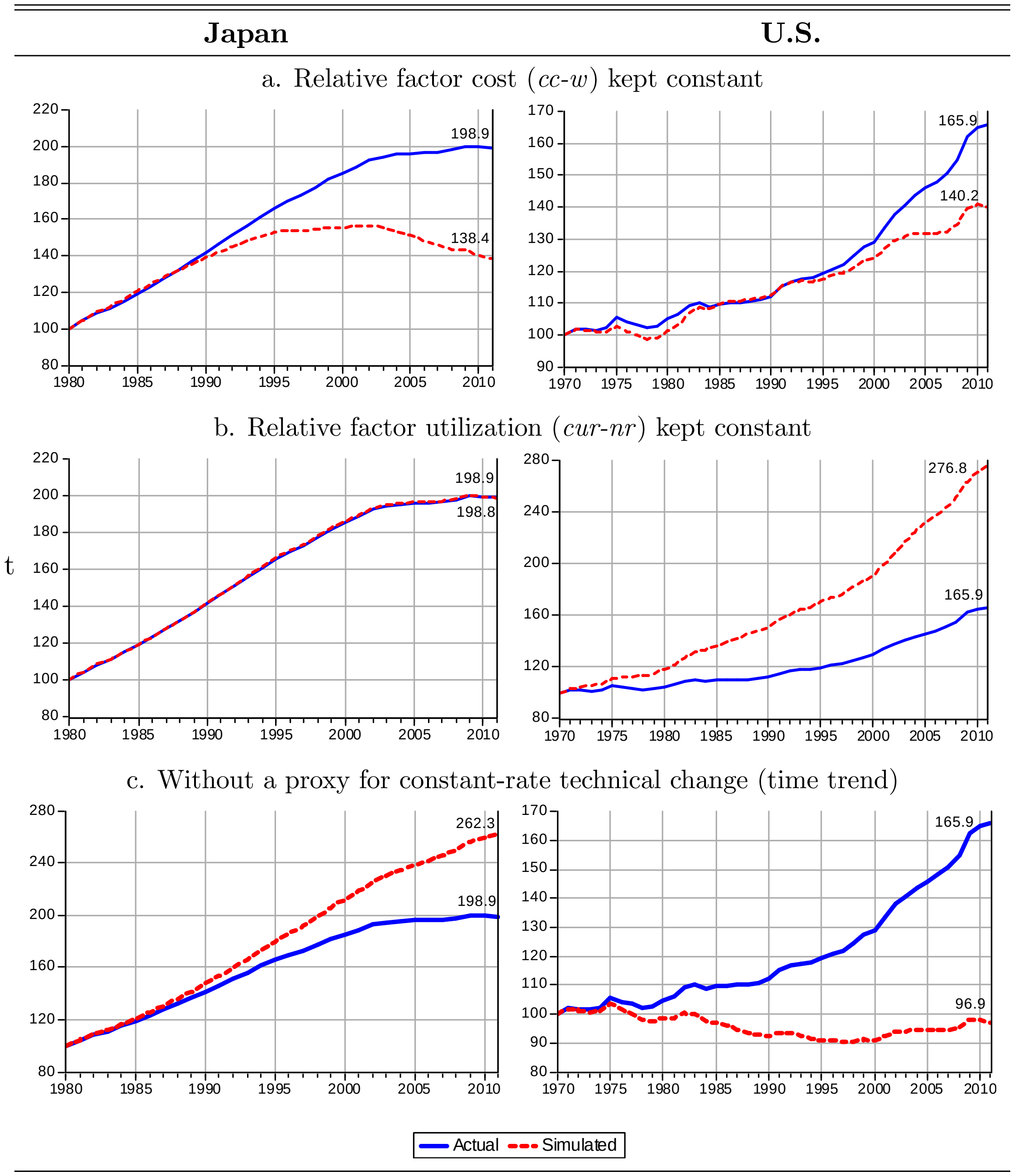
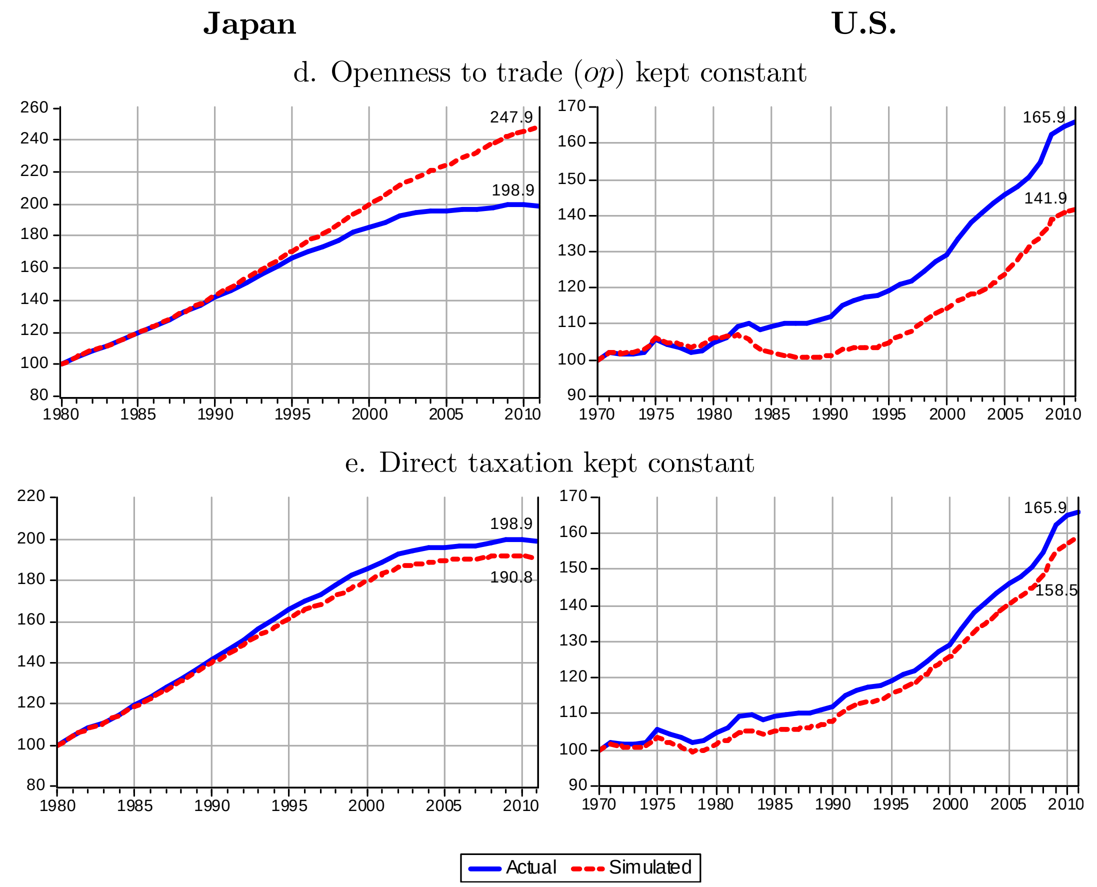


Figure 2 *(cont.)*



The actual path of capital intensity, represented by the baseline scenario (continuous line), almost doubled in Japan between 1980 and 2011 (98.9% growth) and it grew a 65.9% in the U.S. over the 1970 to 2011 period. When the main supply-side determinant of capital intensity ($cc - w$) is kept constant at the beginning of the sample, the fitted values in this second scenario display a growth of only 38.4% in Japan and 40.2% in the U.S. The relative cost of factors declined in both countries over the sample period. That means that without this decline, keeping the value of $cc - w$ constant at the beginning of the sample, capital intensity could have had a lower evolution path. Hence, we verify the negative effect of relative factor cost growth on capital intensity.

When turning to the analysis of the demand-side factor, we find that keeping the relative factor utilization rate at its value in the beginning of the sample has a remarkable effect on the simulated path of capital intensity in the U.S. However, the same cannot be said in Japan's case. The evolution of $cur - nr$ in Japan over the 1980-2011 has different faces: it fell during the 1990s and grew steeply over the 2000s. The value of $cur - nr$ in 2011 is a 10% higher than in 1980. In the U.S. the evolution is more homogeneous: there is a constant fall in $cur - nr$ over the sample period (13% in total).

In the case of Japan, as can be seen in Figure 2 (panel b), the absence of that 10%

growth in the demand-side determinant does not seem to alter the evolution of capital intensity. This is consistent with the fact that the significant variable in the case of Japan is $\Delta(cur - nr)$, so when controlling for this growth the simulated scenario is almost identical to the baseline scenario.

In turn, in the U.S. the fitted values of the second scenario (without the 10% fall in $cur - nr$) grow up to a 176.8% in contrast to the 65.9% growth of the actual path of capital intensity. That is, without the fall in relative factor utilization that the U.S. experienced, the growth of capital per worker could have been steeper, proving the accelerating effect on capital intensity of demand-side pressures, in this case, a higher capacity utilization relative to the employment rate.

Panel c shows the fitted values of simulating Model 1 without the time trend that proxies constant-rate factor-biased technical change. In Japan, when controlling for efficiency growth, the fitted values grow more than the actual capital intensity (162.3% simulated growth vs. a 98.9% of actual growth). In the case of the U.S., the fitted values of the second scenario have an almost flat evolution (a 3.1% decrease over the sample period). This result again reinforces the idea that Japan's factor-biased technical change is capital saving and in the U.S. case it is labor saving. Moreover, in the case of the U.S., labor-saving technical change is so crucial that in absence of that determinant capital intensity would not grow at all.

On the other hand, the degree of openness to trade doubled over the sample period in Japan. The simulated values in the absence of this path represent a considerably steeper growth of capital intensity (panel d). In the U.S., the average growth rate of openness (2.8%) is lower than the initial value in 1970 (6.5%), and the simulated values with this constant growth rate at 2.8% are slightly lower than the actual capital intensity growth. Both results verify the negative effect of a higher exposure to international trade on capital intensity.

Finally, direct taxation also proves to have a negative effect on the evolution of capital per worker. The growth rate of direct taxes (on businesses and households) in Japan is lower in 2011 than at the beginning of the sample (the average growth rate for the sample period is negative). In the U.S., direct taxes on businesses (as a % of GDP) are lower during most of the sample than they were in 1970 and direct taxes on households grow in the second half of the 1990s but they fall again afterwards. In both countries the simulated path of capital intensity is lower than the actual evolution, specially in the U.S. Then, without the fall in direct taxation variables, there would be lower growth of capital intensity, verifying the negative effect of direct taxation on the growth of the capital-over-employment ratio.

6 Concluding remarks

This paper focuses on a generally unattended issue: the determination of capital intensity. The capital-per-worker ratio is usually considered as an input in growth accounting and the empirical assessment of its determinants has been a rather neglected topic.

We develop an analytical setting that includes demand-side considerations to the single-equation capital intensity model of the type used in Antràs (2004) and McAdam and Willman (2013). From this setting, we estimate an empirical model where the determinants of capital intensity include supply- and demand-side determinants, technology, and relevant controls related to international trade exposure and the tax system.

We confirm the relative cost of production factors as a main supply-side driver of capital intensity yielding, also, plausible estimates of the elasticity of substitution between capital and labor. The two proxies we consider for the demand-side pressures experienced by firms are also found relevant in the U.S., and partly so in the case of Japan. In any case, the effect of demand-side pressures is non-neglectable which calls for a wider approach than the usual one when working with production factor demands and, as we have done, when examining the determinants of capital intensity.

By following the factor-specific efficiency growth suggested by Acemoglu (2003), Klump *et al.* (2012), McAdam and Willman (2013), and related papers, we have also uncovered the possibility, at least according to our analytical setting, of a different nature of technological change in Japan and the US. As argued, this divergence between both countries provides an explanation of their contrasted evolution of capital per worker, and even of their diverse growth models; Japan having been, traditionally, one of the great world net exporters and the US having been, and being, one of the greatest net importing economies.

Policywise, our results warn about a simplistic design of policies exclusively based on supply-side considerations. On the supply-side, our finding also calls for a careful design of policies affecting firms' decisions on investment and hiring. The reason is that these policies crucially affect the procyclical behavior of the ratio between the rates of capacity utilization and (the use of) employment, since in economic expansions the capacity utilization rate tends to increase proportionally more than the employment rate, probably because in the very short run it is less costly to use already installed capacity than to hire new workers. From this point of view, the design and implementation of labor market reforms should be closely connected to investment policies, a conclusion already obtained in Sala and Silva (2013) in their analysis of labor productivity.

Usually, demand-side forces are not included in the analysis of economic-growth modeling. This paper joins a strand of the literature that calls for a review of this position. Our results show the incidence of demand-side pressures on the evolution of capital intensity. Furthermore, the simulation exercise indicates that the growth path of capital intensity in the U.S. could have been much steeper without the fall in the relative factor utilization rate. Considering capital intensity is a main growth driver, this result has important policy implications in the fields of economic growth and development.

To conclude, there are three sources of potential improvements in this analysis. The first one is the introduction of imperfect competition in factor markets. There is work done regarding the labor market (Raurich *et al.* 2012), but financial markets, and the associated mark-up over the marginal product of capital, should simultaneously be evaluated. The second one, as explained in León-Ledesma *et al.* (2010), is to block potential identification problems by moving from single-equation estimates of the elasticity of substitution to multi-equation systems in which output and all factor demands are modeled. The third avenue for improvement is to relax the assumptions on technological change and devote further effort in modeling efficiency progress by explicitly considering R&D and innovation. Future research will have to face these compelling challenges.

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Appendix

Table A1. Unit root tests of main variables.

	Japan					US				
	kd	$cc.w$	$cur.nr$	op	Δhrs	kd	$cc.w$	$cur.nr$	op	Δhrs
ADF	0.33 $I(1)$	0.71 $I(1)$	-1.11 $I(1)$	0.03 $I(1)$	-4.80 $I(0)$	3.18 $I(1)$	-0.94 $I(1)$	-2.45 $I(1)$	-0.25 $I(1)$	-4.95 $I(0)$
KPSS	0.72 $I(1)$	0.70 $I(1)$	0.20 $I(0)$	0.61 $I(1)$	0.13 $I(0)$	0.78 $I(1)$	0.54 $I(1)$	0.75 $I(1)$	0.81 $I(1)$	0.23 $I(0)$
Result	I(1)	I(1)	—	I(1)	I(0)	I(1)	I(1)	I(1)	I(1)	I(0)

ADF = Augmented Dickey-Fuller Test. Hypothesis of unit root.

1% and 5% critical values = -3.66 and -2.96 respectively.

KPSS = Kwiatkowski-Phillips-Schmidt-Shin Test. Hypothesis of stationarity.

1% and 5% critical values = 0.739 and 0.463 respectively.

Table A2. Misspecification tests.

	Japan				US			
	Model 1		Model 2		Model 1		Model 2	
	OLS	TSLS	OLS	TSLS	OLS	TSLS	OLS	TSLS
SC [$\chi^2(1)$]	0.01 [0.931]	0.08 [0.772]	1.53 [0.216]	0.64 [0.425]	0.52 [0.470]	0.003 [0.960]	0.01 [0.916]	2.17 [0.141]
HET [$\chi^2(a)$]	9.83 [0.364]	7.55 [0.580]	8.05 [0.529]	8.04 [0.530]	10.9 [0.615]	9.98 [0.696]	10.2 [0.602]	9.03 [0.700]
ARCH [$\chi^2(1)$]	0.52 [0.470]	1.88 [0.170]	0.09 [0.770]	0.04 [0.853]	1.95 [0.163]	1.61 [0.204]	0.16 [0.693]	0.06 [0.801]
NOR [JB]	1.57 [0.457]	1.87 [0.392]	1.46 [0.481]	0.93 [0.627]	0.47 [0.791]	1.62 [0.445]	0.45 [0.799]	1.25 [0.534]

Notes: *p-values* in brackets.

SC = Lagrange multiplier test for serial correlation of residuals;

HET = White test for Heteroscedasticity; NOR = Jarque-Bera test for Normality

ARCH = Autoregressive Conditional Heteroscedasticity; a = number of coefficients in estimated equation (intercept not included).

Figure A1: Evolution of main variables (Japan).

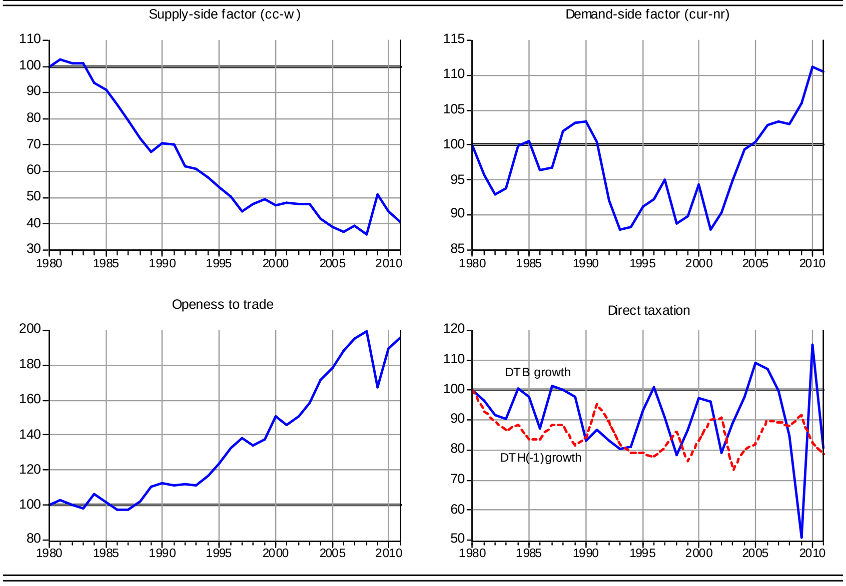
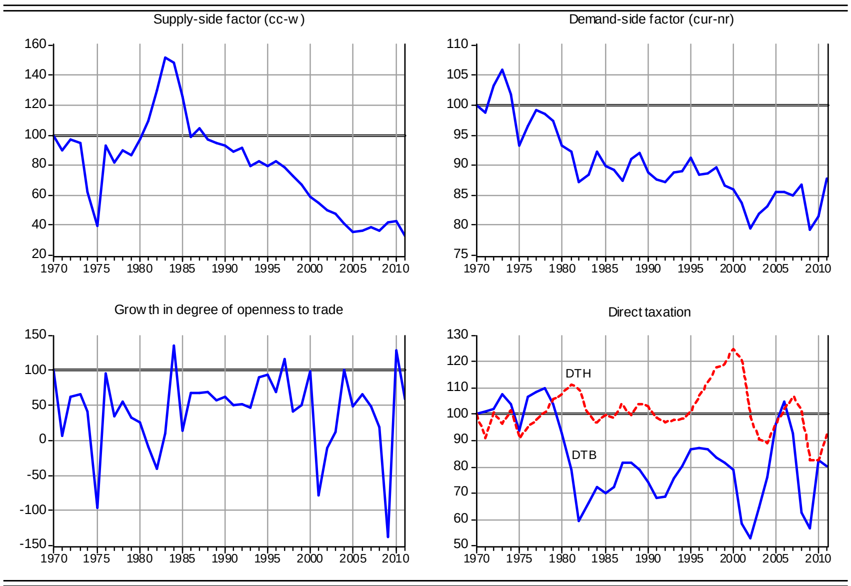


Figure A2: Evolution of main variables (U.S.).



Essay 3

Heterogeneous labor demand: sectoral elasticity and trade effects in the U.S., Germany and Sweden.

Abstract

This paper analyzes labor demand at the sector level in the U.S., Germany and Sweden in two ways: by providing new computations of the sector elasticity of labor demand, and by evaluating the employment effects of trade in manufactures, services, agriculture and fuel. We compute the elasticity through a standard fixed-effects model (i.e., under the assumption of full coefficient homogeneity) and then by taking a semi-pooling sector-level approach (i.e., by flexibilizing the homogeneity assumption). The results reveal that most sector-level elasticities differ largely from the aggregate estimate in all three countries. Also, the sector elasticity values are generally higher in the U.S. and Sweden than they are in Germany. Among the most flexible sectors are manufacturing in Germany and Sweden, the IT sector in the U.S. and Germany, as well as the mining and energy sectors in the U.S. and Sweden. On the other hand, the employment effect of openness to trade is generally positive, although it varies according to country-level differences. We also measure employment effect of technical change and the growth rate of labor efficiency, to find that the latter is similar in the U.S. and Sweden, and small or inexistent in Germany. Since there is a decelerative employment effect of technical change, these results may help in understanding Germany's remarkable employment performance over the last decade.

1 Introduction

To what extent are labor markets flexible (or not)? Should they be further flexibilized? The recent worldwide economic crisis caused high unemployment levels (10.2% in the Euro area and 9.0% in the U.S. in 2011) and aroused the standard economic policy advice of labor market flexibilization. This advice is based on the classical idea that wage rigidity over the market clearing level does not let unemployment to cool down, and has been used to argue, for example, that more flexible labor markets recover faster from financial crises (Bernal-Verdugo *et al.*, 2012). Another strand of the literature, however, dissents from this mainstream view by stressing that recent data shows that the U.S. "flexible jobs machine" may be failing relative to other "less flexible" economies like Germany (Freeman, 2013).

Whatever the case, the achievement of a certain level of unemployment is the result of the aggregation of employment dynamics (jobs creation and destruction) in each economic sector. In this context, in case of sectoral heterogeneity, for a fine tuning of policy design it is crucial to identify these differences.

In a recent contribution, Young (2013) provides new estimates of the elasticity of substitution between labor and capital (σ) in the U.S. at the industry level. He argues that σ differs significantly across industries which creates heterogeneous responses to economic policy. For example, a tax policy that increases the user cost of capital will affect disproportionately the demand for capital where σ is larger. Hence, the focus on sector-level employment is vital for a better understanding of labor market outcomes.

This paper analyzes labor demand at the sector level in the U.S., Germany and Sweden from two perspectives. First, we provide and contrast new computations of sector labor-demand as well as the aggregate labor demand elasticities (ε). Second, we differentiate the effect of trade on employment by four types of merchandises: manufactures, services, agriculture and fuel.

We argue that sector-level mechanisms are essential to labor market outcomes and usually concealed behind aggregate results. The heterogeneity in ε at the sector level is a measure of the unbalanced effects on employment of any potential labor market policy or shock. These diverse effects call for sector-level tailoring of labor market policy, at least as a complement to economy-wide ways of action. The dependence of labor market dynamics on the institutional setting is frequently mentioned in the literature and calls for country-level study and comparability. Accordingly, the analysis in this paper takes a step further than Young (2013) by providing international comparison between economies representative of three different labor market types.

According to Slaughter (2001) the importance of measuring the elasticity of labor demand relies on three main pillars. First, the higher the elasticity of labor demand, then new labor costs (like higher payroll taxes) have a proportionally higher effect on labor than it does on firms. Second, a higher elasticity implies a higher sensitivity of employment to any exogenous shock to wages or labor demand. And third, with a higher elasticity, labor has lower bargaining power over rent distribution, and thus a declining labor income share is expected. Hence, policy addressed to increase the employment elasticity, allegedly intended to lower unemployment, may have backfire effects for workers and households. The decline of the labor income share over labor market deregulation and trade liberalization are issues covered in Judzik and Sala (2013) and Stockhammer (2013).

We also examine the employment effects of higher openness to trade in manufactures, services, agriculture, and fuel. Both aspects, the elasticity of labor demand and trade, relate closely. There is evidence that labor market flexibility has increased in recent decades because of the higher exposure to international trade (e.g. Slaughter 2001, Hijzen and Swaim 2010), although less efforts have been devoted to analyzing the influence of international trade on the number of workers employed using sector level data.

We contribute to this literature by tackling the following question: how does further openness to international trade affect employment? The relevance of this question relies on the fact that the employment consequences of international trade are still an unresolved issue (see, for example, Rueda-Cantuche *et al.,* 2013; and Jansen and Lee, 2007). Jansen and Lee (2007) stress that "the only general conclusion that may be justified is that employment effects depend on a large number of country-specific factors" (ibid, p. 30), which again calls for individual-country analysis. The same authors also argue that most existing studies of trade and employment refer to manufacturing employment, which leaves most of the economy unattended (manufactures represented in 2010 about 12% of total value added in the U.S., 19% in Sweden and 22% in Germany). In this paper, we extend the analysis to the whole economy.

In order to achieve our objective of identifying sector-level elasticities of labor demand, our econometric analysis is performed on a semi-pooling approach (Nunziata, 2005; Heinz and Rusinova, 2011). We estimate a pooled model under the usual assumption of full coefficient homogeneity, and also by applying a semi-pooling approach conceived as an intermediate stage of aggregation between full homogeneity and the other extreme (i.e. individual time-series estimation for each cross-section). This intermediate level of aggregation allows us to find labor-demand elasticities not only for the aggregate economy, but also at the sector-level in each country, while also benefiting from the efficiency gains of pooling control variables.

The analytical framework for our empirical analysis is based on two steps. First, we present a standard formulation of a sectoral labor demand where employment in each sector depends on standard factors such as sectoral average real wage, sectoral value added, openness to trade and a time trend proxying technical change. Second, we compute the output-constant labor-demand elasticity (Hamermesh, 1993) for nine sectors (as defined by the ISIC Revision 4) in the U.S., Germany and Sweden.

The model includes the degree of openness to trade as a determinant of employment following previous research. Its inclusion serves as a control variable aiming at a better estimation of the wage-coefficient in the employment equation and, additionally, it allows the analysis of the effect of trade openness on domestic employment. On a further step, we disaggregate the effects of openness to international trade on sectoral employment in four types of merchandise: manufactures, services, agriculture and fuel. This exercise provides information on which types of trade are more beneficial or detrimental for the evolution of employment in each country. Although we would have preferred to use individual-sector data on international trade, or other variables related to international trade than the degree of openness, data limitations operated as a true constraint.

Our results confirm that the heterogeneity in sector labor-demand elasticity is usually disguised under the common-coefficient assumption imbedded in standard panel data estimations. In other words, the estimated values of sector elasticity of labor demand range in considerably wider intervals than the values found from an aggregate perspective in all three countries. Sector elasticity values are generally higher in the U.S. and in Sweden than they are in Germany.

If we rank sectors according to their estimated labor demand elasticity, some sectors are repeatedly among the highest ranked values. For example, manufacturing in Germany and Sweden, the IT sectors in the U.S. and Germany, and the mining and energy sectors in the U.S. and Sweden. In contrast, the retail trade sector has the lowest elasticities in the U.S. and Germany, together with the finance services sector in Germany and Sweden. Notably, in our results we do not observe general criteria in terms of manufacturing having lower or higher elasticity than services sectors at this level of disaggregation. In sum, a one-size-fits-all approach to labor market policy will probably be inefficient since it will have very dissimilar results depending on economic activities and country (or institutional setting).

Regarding openness, a larger exposure to trade is associated with an impulse on employment in the U.S. and Sweden, but not in Germany. This is consistent with our finding of higher sector labor-demand flexibility in the U.S. and Sweden than in Germany. Exposure to international trade tends to increase labor market flexibility and, according to this result, trade has a stronger effect on labor market dynamics in the U.S. and Sweden.

When looking into different types of merchandise, openness to trade in manufactures has also an accelerating effect on employment in both the U.S. and Sweden, as expected. However, a higher level of trade in services has a positive effect on employment in Sweden and a negative impact in the U.S. We believe that the different role of services industries in each of these countries, plus service offshoring and its skill-biased effect on domestic employment may provide possible explanations.

In line with recent research, we identify and measure technological change. Our estimations assert that there is a relevant negative employment effect of technical change. At first glance, this effect is similar in Sweden and Germany, and bigger in the U.S. But when turning to the growth rate of labor efficiency, it results to be similar in Sweden and the U.S., and small in Germany. As we explain below, this result contributes to the understanding of the better employment performance in Germany over the last decade.

The rest of the paper is structured as follows. Section 2 presents a bird-eye view of stylized facts regarding employment structure and openness to trade. Section 3 provides the analytical framework. Section 4 stresses the econometric analysis, and section 5 presents and discusses the results. Finally, section 6 concludes.

2 Stylized characterization of sectoral employment and trade exposure

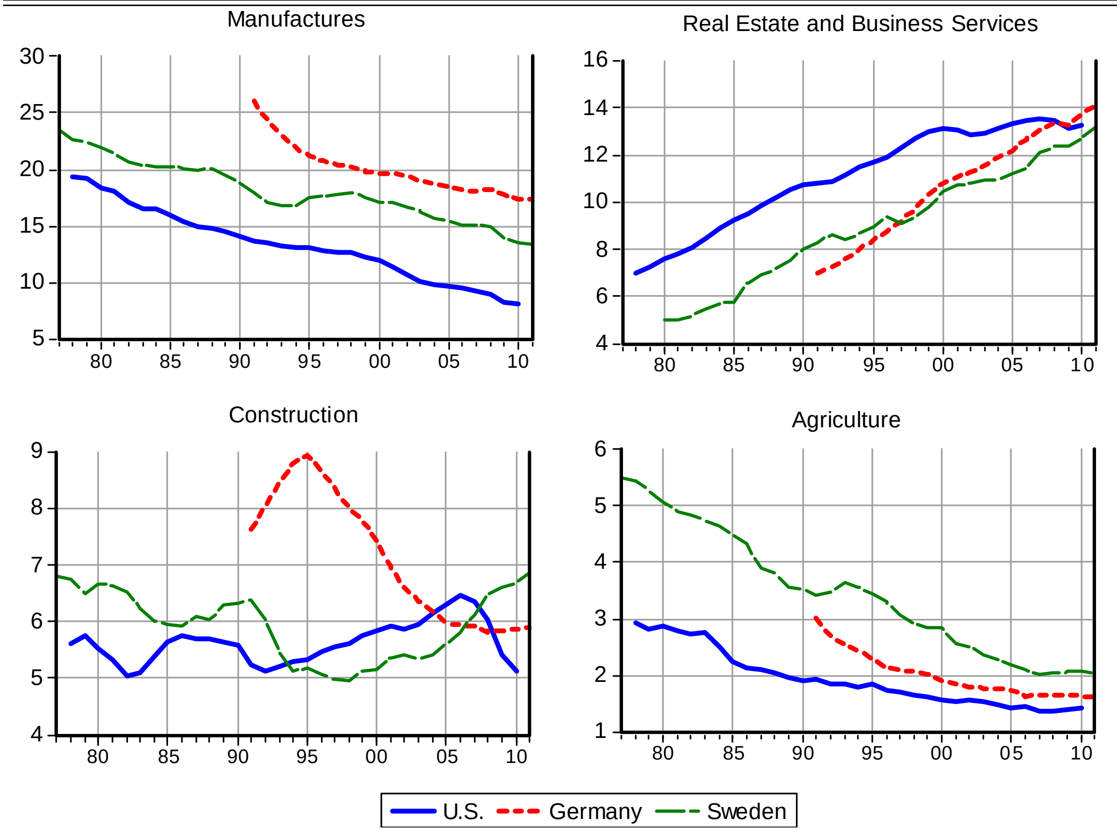
The U.S., Germany and Sweden are industrialized economies with diverse labor market structures and exposures to international trade. Regarding the institutional setting of the labor market, these three countries represent examples of three frequently cited categories of labor market structure according to their tax and welfare systems (e.g. Daveri and Tabellini, 2000): the Anglo-Saxon (U.S.), the Continental Europe (Germany) and the Nordic (Sweden) setting.

At the aggregate level, these three economies had different labor market performances over the last few decades, specially after 2008. Germany introduced major labor market reforms between 2003 and 2005 (so-called Hartz reforms) that included new strong employment policy and services, a reduction in long-term unemployment with new incentives for job searching, and deregulation of fixed-term contracts to stimulate labor demand. These reforms contributed to Germany's resilience to the Great Recession (Rinne and Zimmermann, 2013) and went further than mere flexibilization.

As put by Freeman *et al.* (2010), "the Swedish economic model is perhaps the most ambitious and publicized effort by a capitalist market economy to develop a large and active welfare state" (ibid, p. 1). Sweden suffered a strong economic crisis in the first part of the 1990s from which recovered by curbing this ambitious welfare state; with strong policy reforms concerning flexible exchange rates and inflation targeting for stronger currency and export-led growth, contraction of the public sector, reduced generosity in social insurance systems, and deregulation in product markets (Freeman *et al.*, 2010). The recession that started in 2008 in the U.S. had similar causes than the 1990s crisis in Sweden (deregulated financial markets and bubble burst in asset pricing transmitted from banks to the whole economy). This time around, perhaps, Sweden was better prepared.

But when looking at sector-level behavior, sectors have evolved in different ways. The U.S. and Sweden have become more service-oriented economies, whereas in Germany manufactures and construction represent more important parts of the economy. Figure 1 presents the evolution of employment of selected sectors in the U.S., Germany and Sweden.

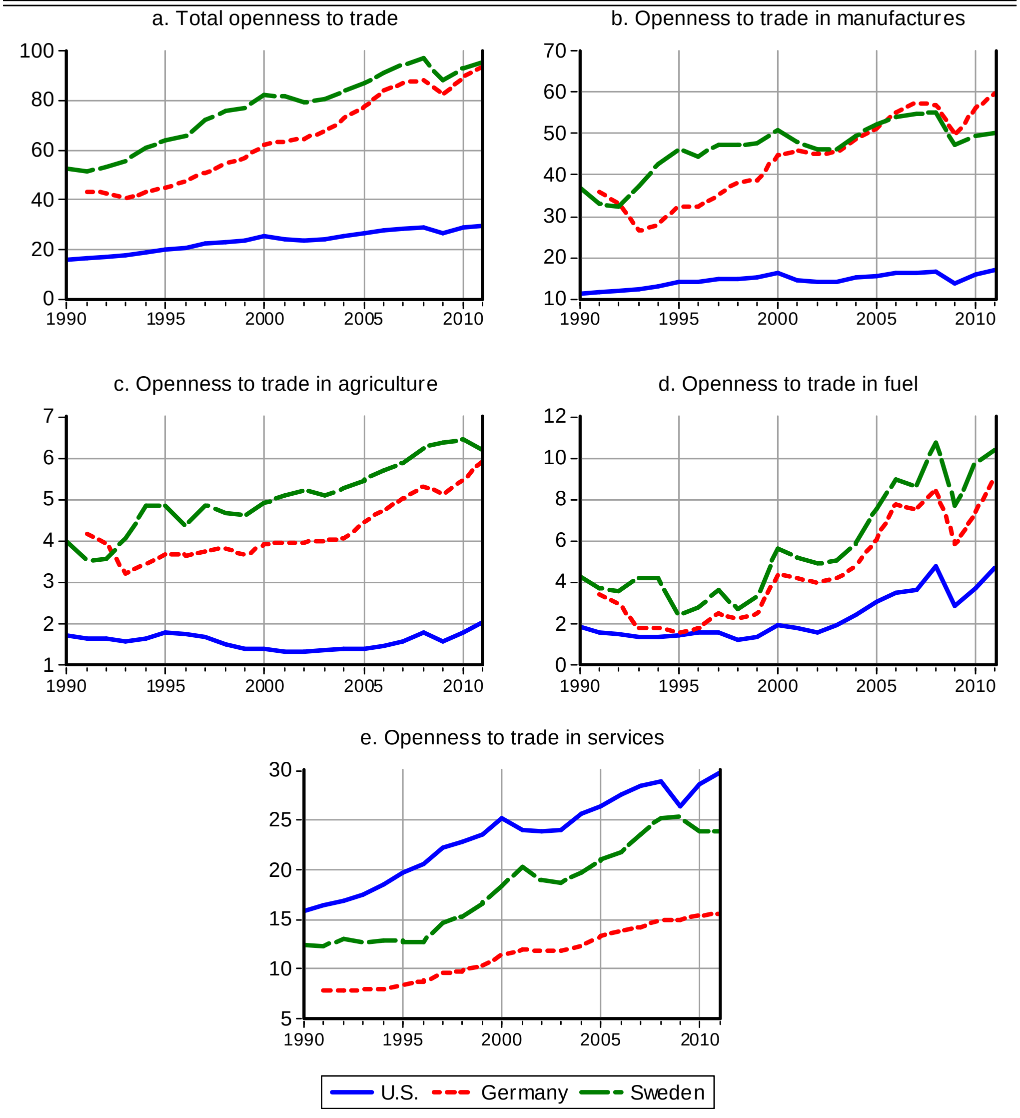
Figure 1. Sectoral employment (% of total employment).



The percentage of employment allocated in manufactures is higher in Germany than in Sweden and the U.S., but it has declined in all three since the 1980s. In turn, the real estate and business services sector employs an increasing proportion of workers. Note that at the last available observation, in Germany there is still a higher percentage of employment in manufactures than in real estate and business services (17.4% and 14.1% respectively), in Sweden it is almost the same (13.4% and 13.2%), while in the U.S. the proportion of employment in manufactures is now lower than that of the real estate and business services sector (8.2% and 13.5%). This structural change in sector-level employment has been the object of study in several works (e.g. Schettkat and Yocarini, 2006).

Structural change has not arrived everywhere. In all three countries the retail trade and financial services sectors have not increased significantly the proportion of employment over the last decades. The U.S. has the highest proportion of sectoral employment in both sectors, Sweden has the lowest, and Germany is an intermediate case. Retail trade represents more than 20% of employment in all three countries, while finance and insurance services still represent less than 5% of total employment.

Figure 2. Degree of openness to international trade (%).



On the other hand, these three economies have different degrees of exposure to international trade. The rate of total trade (exports plus imports) over GDP is a frequently used proxy of the degree of openness to international trade. It is of around 30% in the U.S., 94% in Germany and 95% in Sweden (data of 2011). In this sense the former is a lesser open economy and the two latter are much more open ones (Figure 2, plot a). The degree of openness in the U.S. had a flat evolution since 1990, while it doubled (or nearly doubled) in Germany and Sweden.

These aggregate values, however, do not tell the whole story. For example, Germany and Sweden display a high degree of trade openness in manufactured goods, while trade of manufactures over GDP is less than 20% in the U.S. (plot b). In contrast, the U.S. has the highest level of trade in service industries, closely followed by Sweden, while Germany has a lower third place (plot e).

We believe that these differentiated labor market structures and performance, combined with diverse experiences in employment across sectors (Figure 1), plus also differentiated trade exposures (Figure 2) call for a sector-level computation of the elasticity of labor demand. Different industries have diverse hiring and firing dynamics and, hence, sector labor demand elasticity computations may provide new information than the usual aggregate labor demand elasticity. Moreover, individual-country analysis should be judged appropriate considering that employment responsiveness is conditional on the institutional structure of each economy and unique multifunctional policy cannot be properly tailored.

3 Analytical framework

3.1 A sector-level labor demand model

We follow Young (2013), who adopts industry subscripts to the CES production function with factor-augmenting technological change *à la* Antràs (2004) and McAdam and Willman (2013), in order to incorporate sectors. This scheme represents the behavior of the representative firm for each industry instead of the representative firm for the aggregate economy.

Accordingly, consider a CES production function where the representative firm in sector i in period t produces real output Q following:

$$Q_{it} = \left[\theta_i (A_t^N N_{it})^{-\beta_i} + (1-\theta_i)\left(A_t^K K_{it}\right)^{-\beta_i}\right]^{-1/\beta_i}, \tag{1}$$

where K = capital stock and N = employment; A_t^N and A_t^K are time-varying coefficients of technological change; A_t^N proxies labor-augmenting (Harrod-neutral) technical change and A_t^K proxies capital-augmenting (Solow-neutral) technical change; $\sigma = \frac{1}{1+\beta}$ is sector

i constant elasticity of substitution between capital and labor and θ_i is sector i constant coefficient of factor share $(0 < \theta < 1)$.

Note that $A_t^N = A_t^K$ implies Hicks-neutral technical change. We apply no *a priori* restrictions in this sense (see Section 3.2) and apply a flexible set-up with factor-specific efficiency growth.

- **The sectoral demand for labor**

A profit-maximizing firm in a competitive environment will employ labor so that the marginal productivity equals the real wage rate:

$$\frac{\partial Q_{it}}{\partial N_{it}} = MPL_{it} = W_{it} \tag{2}$$

where W = real wage rate and MPL = marginal productivity of labor. According to (2), deriving from (1) we find that:

$$W_{it} = \theta_i (A_t^N)^{-\beta_i} (Q_{it})^{1+\beta_i} (N_{it})^{-(1+\beta_i)} \tag{3}$$

Solving for N:

$$N_{it} = (\theta_i)^{\frac{1}{1+\beta_i}} (W_{it})^{\frac{-1}{1+\beta_i}} (A_t^N)^{\frac{-\beta_i}{1+\beta_i}} Q_{it} \tag{4}$$

and log-linearizing we find an employment equation representation of a marginal productivity condition:

$$n_{it} = \sigma_i \log \theta_i - \sigma_i w_{it} + q_{it} - (1 - \sigma_i) \log A_t^N \tag{5}$$

where $n = \log(N)$, $w = \log(W)$ and $q = \log(Q)$.

Following the hypothesis in Antràs (2004) we assume that labor efficiency grows at a constant rate and A_t^N is determined as follows:

$$A_t^N = A_0^N e^{\lambda_N . t} \tag{6}$$

where t is a time trend, λ_N is the constant rate of labor-augmenting efficiency growth and A_0^N is the initial value of the efficiency coefficient.

Moreover, we include openness to trade for two reasons: as a control variable (since there is evidence that trade liberalization affects the elasticity of labor demand) and to analyze its effect on employment. Then further disaggregation in four types of merchandise provides information on what sort of trade is more or less favorable to domestic employment in the three economies studied.

Hence, (5) can be re-expressed as:

$$n_{it} = \alpha_i - \sigma_i w_{it} + q_{it} - (1-\sigma_i)\lambda_N t + \lambda_{op} op_t \tag{7}$$

where $\alpha_i = \sigma_i \log \theta_i - (1-\sigma_i) A_0^N$ is a cross-section specific intercept. Equation (7) is the baseline equation. It presents the time-evolution of employment in each sector as determined by: a cross-section intercept, the average real wage in that sector, the sectoral output or value added, a time trend as a proxy for technical change, and the degree of openness to international trade. Note that the coefficient associated to the real wage is the sector-level constant elasticity of substitution between labor and capital.

- **The output-constant elasticity of labor demand**

Following Hammermesh (1993) we compute the output-constant elasticity of labor demand at the sector level using the estimated elasticity of substitution between labor and capital (σ_i) from the model described above. The Hicks-Allen elasticity of substitution was defined as changes in relative factor price on relative inputs of the two factors, holding output constant. That is:

$$\sigma = \frac{d\ln(K/N)}{d\ln(w/r)} = \frac{d\ln(K/N)}{d\ln(F_K/F_N)} = \frac{F_N.F_K}{Y.F_{NK}} \tag{8}$$

where $F(K,N)$ is a generic production function, r is the user cost of capital and $F_N = w$ and $F_K = r$ under the assumption of a competitive environment.

Then Hamermesh (1993) defined the own-wage elasticity of labor demand (with output and cost of capital constant) as:

$$\varepsilon_i = -(1-s_i)\sigma_i \tag{9}$$

where s_i is labor's share in sectoral value added and subscript i represents each sector. Note that in (9) output is kept constant but the capital-labor ratio is allowed to vary as the relative price of production factors changes. Each ε_i is computed with the estimated σ_i from our empirical model and the average s_i from the data . Thus, the computed sector elasticity depends on the relative availability of capital in that sector and the elasticity of substitution. The sectors where labor represents a lower share of income are associated with a higher elasticity of labor demand. Likewise, a higher elasticity of substitution makes labor more easily substitutable by capital, and this also implies a higher elasticity of labor demand.

3.2 Discussion

Our functional form for aggregate production [equation (1)] is equivalent to equation (1) in Young (2013). In this way, we follow a broad strand of the literature that deals with the modeling of the aggregate production assuming a CES functional form along the lines of Arrow *et al.* (1961). The employment equation obtained is a productivity condition derived from the optimization of aggregate production.

First, it is important to stress that it is a mistake to interpret the coefficient of real wage as an output-constant elasticity of demand, since by equation (7), it is actually σ. According to Hamermesh (1993), the output-constant elasticity of labor demand is the elasticity of substitution between labor and capital adjusted by the capital share of total income. If a Cobb-Douglas technology of production is assumed, the elasticity of substitution between labor and capital is one, the labor share of income around 0.66, and the elasticity of labor demand around -0.33. But when flexibilizing the aggregate production to take a CES form, the long-run coefficient associated to the real wage is the elasticity of substitution between labor and capital. The crucial point is that this elasticity can be estimated instead of assumed to be unity.

Hence, the substitutability between capital and labor is at the core of the elasticity of labor demand with respect to the real wage. As stressed by Rowthorn (1999), economics based on Cobb-Douglas production functions (with $\sigma = 1$) implies that an increase in real wages generated by investment in new capital leads to a loss of employment on existing equipment, which is enough to offset entirely the extra jobs created on new equipment, and therefore capital investment cannot increase employment in the long run. There is large evidence that the elasticity of substitution between capital and labor is significantly lower than one, specially in the U.S. (e.g., McAdam and Willman, 2013; Klump *et al.,* 2012; Chirinko *et al.,* 2011; León-Ledesma *et al.,* 2010; and Chirinko, 2008).

Second, once an aggregate production is modeled as a CES production function, there is a choice between Hicks-neutral or factor-augmenting technical change. On this account, we follow Acemoglu (2003), Antràs (2004), León-Ledesma *et al.* (2010), and McAdam and Willman (2013), among others, in adopting a factor-augmenting approach. This allows for the identification of factor-biased technical change and the measure of its incidence, instead of undertaking, for example, the *a priori* assumption of Hicks neutrality. This literature is relatively new, and presents estimates that identify a significant labor-saving effect of technological change in the U.S. (Klump *et al.*, 2012).

Lastly, the functional form of the sector employment equation must reflect the fact that exposure to trade affects labor market outcomes. Recent evidence points in the direction that higher trade intensity affects employment (e.g., Felbermayr *et al.*, 2011; Gozgor, 2013; and Yanikkaya, 2013).

4 Econometric analysis

This section discusses the methodological aspects of our endeavor. The choices that determine those aspects are made in correspondence to the type of data and empirical objectives of this study. Also, we follow recent literature in dealing with the critical issues faced by related research. The first part is standard: we select estimation methods appropriate for our database and empirical model. Second, regarding the issue of cross-section heterogeneity, we argue in favor of a semi-pooling approach as our empirical strategy. We understand as a semi-pooling approach as an intermediate stage between full parameter homogeneity (that is, one constant coefficient for all cross-sections, the most common approach to panel data) and the individual cross-section estimations for all variables in time-series models. In this paper, a semi-pooled regression refers to the estimation of individual cross-section coefficients for key variables and homogenous coefficients associated to control variables.

4.1 Estimation methodology

The choice of estimation methodology in panel-data macroeconomic models is not trivial. Usually, panel data estimations are designed for a large cross-section dimension ($i = 1...N$) and a few time periods ($t = 1...T$). Moreover, some underlying assumptions are based on the fact that many N homogeneous cross-sections are randomly selected out of a much bigger population (e.g., individuals, households or firms). In this scenario, to model with common coefficients for all cross sections is efficient and advisable.

In our case, we have three panels with $N = 9$ sectors that cover the whole economy, and the maximum availability of time-periods (T differs from 19 to 40 annual observations, depending the case). These are panels with $T > N$ where the homogeneity assumption does not hold. When the database is a pool of short time-series, where each one constitutes a cross-section unit with a strong personality like countries or sectors, the standard panel data models may not be the best fit.

A common practice is the inclusion of fixed-effects (FE, i.e. cross-section specific intercepts) to control for some degree of baseline heterogeneity (that is, constant heterogeneity through time). Not only this control for heterogeneity is not enough in our case, but also the OLS with FE model presents a bias in dynamic specifications as shown in Nickell (1981) and henceforth known as Nickell bias. This bias may be reduced when T is high, which it is so in our panels.

Another issue comes along the inclusion of the lagged endogenous variable for the explicit modeling of dynamics in sector employment: it introduces the impossibility to hold the OLS assumption of strict exogeneity of the regressors. Regarding this issue, an instrumental variable method should be considered to avoid the menace of endogeneity bias.

Pooling time-series together introduces new problems related to the spherical errors assumption. While cross-sectional errors may be homoscedastic and non auto-correlated, the pool has new issues, because homoscedasticity is required across both dimension. When having cross-section units with strong personality as in our case (economic sectors) it is likely that cross-section residuals will have different variances and thus the panel will be heteroscedatic across N. Also, since they are sectors of the same economy (and country) they have common unobservable variables, so that the disturbances are presumably correlated.

The related literature deals with these issues by using the panel-corrected standard errors (PCSE) suggested by Beck and Katz (1995) and Beck (2001), the feasible generalized least squares estimator (FGLS) and instrumental variables (Gnagnon 2013; Zhu 2013).

The OLS estimation with PCSE, while still assuming same-unit homoscedasticity as the usual time-series models, corrects for contemporaneous correlation of common unobservables and inter-unit heteroscedasticity (the so-called "panel heteroscedasticity") caused by the pooling of several time-series (Beck and Katz, 1995). Therefore the PCSE is a robust standard error approach for cross-unit dependence (Zhu, 2013).

Moreover, the standard FGLS is a highly used method among the studies with $T > N$ panels (e.g. Heinz and Rusinova, 2011). Instrumental variables are included to control for the potential endogeneity of the dynamic modeling as well as for the fact that real wage may not be exogenous to employment (Lewis and McDonald, 2002). Then, the second method used is a two-stage FGLS with instrumental variables (TS-FGLS). Recent contributions like Young (2013) also add instrumental variables to the GLS framework for the same reason. In this FGLS context, cross-section weights provide residuals robust to cross-sectional heteroscedasticity.

4.2 To pool or not to pool?

A standard modelization under full cross-section homogeneity would give biased estimations. Many argue that this assumption rarely holds in non-randomized observational studies (Zhu, 2013). The heterogeneity bias that arises from estimating constant coefficients for all cross-sections in a heterogeneous dynamic panel model persists regardless the number of cross-section dimensions, time periods and choice of instrumental variables (Pesaran and Smith, 1995). Moreover, cross-section units that respond to sectors or countries rather than individuals or firms are likely to be heterogeneous. It follows that an effective control for heterogeneity must be examined.

The fixed-effects (FE) model controls for baseline unobserved heterogeneity with a cross-section specific intercept. In a dynamic heterogeneous model the FE approximation, which imposes coefficient homogeneity (i.e., identical slopes for all cross-section units), may give inconsistent estimations (Steiner, 2011). A dynamic heterogeneous panel model needs to take into account the different responses of sector employment to changes in the main variables. But even if the FE model was not biased we would be estimating an "average" slope. So we need to ask ourselves: is this useful to our empirical objective? One can easily find, for example, a "not significant" slope (i.e. statistically zero) when actually every cross-sectional slope is non-zero, but as they are "summed up" they cancel out each other (Juhl and Lugovskyy, 2013).

At the opposite end, there is the random coefficient model where both intercept and all estimated coefficients vary across economic sectors (i). This model entails the estimation of numerous coefficients thus requires large panel dimensions (degrees of freedom). For that reason this model may not be adequate for our database.

It is a main concern in panel data analysis how much to pool. For the reasons described we must consider an intermediate degree of pooling between the full-homogeneity assumption and the individual coefficient estimation for all intercepts and variables included in the model. Juhl and Lugovskyy (2013) argue that the specification of a "partially heterogeneous" model where some variables share a common slope and others are allowed to be heterogeneous is a viable solution for the pooling issue.

In our model of sector employment, cross-section units consist on nine sectors that clearly present an heterogeneous behavior (see Figure 1), as studied by the structural change literature. Nunziata (2005) faces a similar challenge in a wage-setting study where cross-sections are countries with institutional heterogeneity. He argues that the pooled model yields more efficient estimates than the country by country regression, but the poolability test results are not robust enough to justify a pure coefficient homogeneity framework. In his view, this situation calls for an intermediate degree of poolability that allows for some degree of heterogeneity (at least in key variables), in a pooled data framework that gains efficiency from a common estimation of control variables. This procedure reduces the potential bias from assuming full homogeneity in actually heterogeneous models.

Ultimately, our objective is to find reliable estimates for the elasticity of substitution between production factors at the sector level (σ_i). It would be our preference to perform sector-level time-series estimations, but we come across the lack of large annual time series in several sectors as an inexorable shortcoming. As in Nunziata's case, we need to explore an intermediate degree of pooling that improves the degrees of freedom from the lack of large sector-level time series and at the same time, that allows for cross-section specific estimations of the main coefficients. Zhu (2013) stresses that pooling different time series together while accounting for cross-section heterogeneity can compensate the lack of extended annual data.

As argued by Beck and Katz (2007), there are relatively few attempts like Nunziata (2005) to go beyond the limited heterogeneity provided by the fixed-effects model. They argue that the degree of pooling should be a scientific decision, and then intermediate situations should be explored. Heinz and Rusinova (2011) also decide to pool together the observations for all countries using panel estimation but allowing for differential slopes. They argue that if there are reasons for expecting heterogeneous behavior, this technique could substantially reduce the potential bias introduced by the homogeneity restriction.

This paper uses both methodologies and contrasts the full aggregation of the data with a semi-pooling approach where individual cross-section coefficients are estimated for the key variables (in particular, for those required for the estimation of the elasticity of substitution between capital and labor). The other control variables included in the model share common coefficients to all cross-sections. This system "borrows strength" by estimating only one homogeneous coefficient for control variables and keeping cross-section heterogeneity in the main interest variables: real wage and persistence coefficient (lagged employment). Then we can compute the elasticity of labor demand (with respect to the real wage) for each sector while also gaining efficiency by estimating common coefficients associated to the control variables (value added, openness to trade, and time trend).

4.3 Data

Regarding the data, this paper employs OECD STAN sector-level data including nine sectors following the two-digit ISIC Revision 4 classification: (1) agriculture, hunting, forestry and fishing, (2) mining, energy and waste management, (3) manufacturing, (4) construction, (5) wholesale and retail trade, transportation and storage, accommodation and food service activities, (6) information and communication, (7) finance and insurance activities, (8) real estate and business activities, and finally (9) community, social and personal services. Table 1 defines the variables used in the empirical analysis.

The sample availability for the United States is 1978-2010 for five industries. For the rest, 1988-2010 for community services, 1989-2010 for retail trade, 1998-2011 for mining and energy, and 2000-2010 for information and communication. For Germany the availability is a balanced sample for the 1993-2011 period. For Sweden, the availability of data is 1970-2011 for agriculture, manufactures, mining and energy, and construction, and 1993-2011 for all other sectors.

Table 1. Variable definitions and sources of data.

n	Total employment (number engaged)[25].	OECD Stan
w	Labor compensation of employees.	OECD Stan
va	Value added, volume.	OECD Stan
op	Openness to trade (Exports + Imports) / GDP.	OECD Economic Outlook 91
opm	Openness to trade, Manufactures.	WTO and OECD
ops	Openness to trade, Services.	WTO and OECD
opa	Openness to trade, Agriculture.	WTO and OECD
opf	Openness to trade, Fuel.	WTO and OECD
s	Labor income share ($= \frac{W.N}{VA}$).	
t	Time trend.	

Note: all variables are in logs (except s and t).

Aggregate data of international trade (exports and imports) and GDP are national series from the OECD Economic Outlook 91 (December 2012). Disaggregated data on trade of manufactures, agriculture, fuel and services were extracted from the WTO official database. The labor income share (s) for each sector is computed as the ratio of labor compensation over value added.

4.4 Empirical strategy

We estimate equation (7) from different perspectives on the degree of pooling, alternating both estimation methods discussed in the previous section: the panel-corrected standard errors least squares (PCSE) and two-stage feasible generalized least squares with instrumental variables (TS-FGLS). It is crucial to stress that the empirical models are estimated as dynamic equations to take into account the adjustment costs potentially surrounding all variables involved in the analysis (endogenous and exogenous). Also, the signs of the estimated coefficients will be determined empirically: *ex ante* all coefficients are presented with a + sign behind them.

First we assume full homogeneity of the coefficients, only with fixed effects for each cross-section in order to mitigate estimator bias. In this case, the estimated equation takes the form represented by (10).

$$n_{it} = \beta_{0i} + \beta_1 n_{it-1} + \beta_2 w_{it} + \beta_3 q_{it} + \beta_4 t + \beta_5 op_t + \upsilon_{it} \tag{10}$$

where $\beta_{0i} = \alpha_i$ includes a cross-section fixed effect, β_1 is the persistence coefficient, $\beta_2 = \sigma$

[25] Includes full-time, part-time and self-employed.

is the aggregate elasticity of substitution, $\beta_3 = (1-\sigma)\lambda_N$, $\beta_4 = \lambda_{op}$ and υ_{it} is a well-behaved error term.

On a second step, we estimate an augmented equation with a disaggregation in nine sectors, as detailed above.

$$n_{it} = \gamma_{0i} + \gamma_{1i}n_{it-1} + \gamma_{2i}w_{it} + \gamma_3 q_{it} + \gamma_4 t + \gamma_5 op_t + \nu_{it} \tag{11}$$

The difference with equation (10) is that, in equation (11), the coefficients γ_{1i} and $\gamma_{2i} = \sigma_i$, associated to the effect of the real wage on employment, are estimated individually for each sector (for all i). The rest of estimated coefficients, γ_3, γ_4 and γ_5, remain as homogeneous coefficients (under the borrowing strength concept explained previously).

A third and last step includes the disaggregation of total openness to trade in four variables according to the type of merchandise: openness to trade in manufactures (opm), services (ops), agriculture (opa) and fuel (opf). This gives rise to our third empirical model represented by equation (12).

$$n_{it} = \gamma_{0i} + \gamma_{1i}n_{it-1} + \gamma_{2i}w_{it} + \gamma_3 q_{it} + \gamma_4 t + \gamma_5 opm_t + \gamma_6 ops_t + \gamma_7 opa_t + \gamma_8 opf_t + \nu_{it} \tag{12}$$

Combining the empirical models of sector-level employment represented in equations (10), (11) and (12), and the estimation methods explained in the previous section (PCSE and TS-FGLS), we compute the sectoral elasticity of labor demand in the nine industries included in the sample and evaluate the effect of openness to trade on employment.

5 Results

This section presents the empirical results of our study in three subsections. First, we present and discuss the estimated values of σ_i and computed values of ε_i for each one of the three countries studied. Second, we discuss the employment effect of a higher exposure to international trade. Third, we disclose the employment effect of technological change.

Note that in all tables in section 5.1. we abbreviate the sectors as follows: AG for agriculture, hunting, forestry and fishing; ME for mining, energy and waste management; MA for manufacturing; CO for construction; RT for wholesale and retail trade, transportation and storage, accommodation and food service activities; IT for information and communication; FI for finance and insurance activities; RE for real estate and business activities; and SE for community, social and personal services. Additionally, all the estimated equations are available in the Appendix.

5.1 Sector elasticity of labor demand

The results for the U.S. are generally consistent with the values found by Young (2013). He estimates 35 industry-level elasticities of substitution between capital and labor (σ), with three different specifications and three estimation methods. Table 2 compares our results to those of Young (2013), in an adaptation of his industry-level classification to the 9 sectors used in this paper, this is the reason why it is designed with 3 columns presenting, each, a range of values for: the results in our study, Young's preferred method (GMM), and his alternative method that is similar to one of the used in this paper, that he calls three-stage generalized instrumental variables (GIV).

The ranges of values of our estimated elasticities overlap to those of Young (2013). Only the estimated elasticity for the IT sector is outside the range of values found by Young (2013), although the adaptation from his disaggregation in 35 industries to our sectors is not perfect. For example, finance, real estate and insurance services are combined into one industry, whereas the ISIC Revision 4 classification considers two separate sectors; finance and insurance services on the one hand, and real estate services and on the other one.

Table 2. Estimated U.S. sectoral elasticity of substitution (σ_i).

	This study	Young (2013)	
		GMM	GIV
Agriculture (AG)	[0.35 0.52]	[-0.39 0.68]	[-0.09 0.84]
Energy (ME)	[0.62 0.85]	[0.62 0.87]	[0.57 1.64]
Manufactures (MA)	[0.91 1.30]	[0.02 1.41]*	[-0.34 1.26]
Construction (CO)	[0.83 1.12]	[0.32 0.50]	[0.29 1.01]
Retail (RT)	0.49	[0.42 0.60]	[0.11 1.12]
IT Services (IT)	[1.06 1.22]	[0.42 0.48]	[0.57 1.11]
Finance (FI)	[0.54 1.21]	[0.99 1.00]	[0.66 0.92]
Real Estate (RE)	[1.14 3.68]	**	
Community (SER)	< 0	0.39	[-0.02 1.32]

* [0.21 1.10] without leather industry.

** included in finance and insurance.

The estimation of the sectoral elasticity of substitution between capital and labor (σ_i) is an input in the overall analysis. It is used in the subsequent calculation of the sector elasticity of labor demand (ε_i) which is the central variable of interest. Tables 3, 4 and 5 present the main results for the U.S., Germany and Sweden. In all tables the first column presents the sector labor income share (s_i) computed with the OECD Stan data and used in the calculation of ε_i. In turn, HC denotes homogeneous coefficients, corresponding to the results under the assumption of full coefficient homogeneity [equation (10)].

Table 3. **U.S.** sectoral labor shares, elasticity of substitution and labor demand elasticity.

		1		**2**		**3**		**4**		**5**	
		PCSE		TS-FGLS		PCSE		TS-FGLS		PCSE	
	s	σ	ε	σ	ε	σ	ε	σ	ε	σ	ε
AG	0.25	0.37	-0.28	0.49	-0.37	0.35	-0.26	0.52	-0.39	0.40	-0.30
ME	0.18	0.85	-0.70	0.62	-0.51						
MA	0.64	0.95	-0.34	1.30	-0.47	0.87	-0.32	1.28	-0.46	0.91	-0.33
CO	0.68	0.98	-0.32	0.83	-0.27	1.12	-0.36	0.80^b	-0.26^b	0.84	-0.27
RT	0.72	0.19^b	-0.05^b	0.49	-0.14						
IT	0.57	1.06	-0.45	1.22	-0.52						
FI	0.56	0.95	-0.41	1.21	-0.53	0.95	-0.42	1.07	-0.47	0.54	-0.24
RE	0.35	1.81	-1.17	2.96	-1.91	1.59	-1.03	3.68	-2.38	1.14	-0.73
SE	0.81	-4.70	*	0.33^b	-0.06^b						
HC	0.61	0.50^a	-0.19^a	0.59	-0.23	0.44^a	-0.17^a	0.33^b	-0.13^b	0.52	-0.20
Sample		1978 2010		1978 2010		1978 2010		1978 2010		1980 2010	
Obs		230		229		165		165		155	

Note: PCSE = Panel-corrected standard errors. TSFGLS = two-stage feasible generalized least squares. No superscript = wage-coefficient significance at 10% level.

$^a = 0.10 <$ p-value < 0.15 $^b =$ p-value > 0.15 $^* = \varepsilon_i > 0$

In the case of the U.S., specifications 1 and 2 in Table 3 present the unbalanced estimation with all available observations by PCSE and TS-FGLS respectively. Specifications 3 and 4 are performed with a balanced sample of the sectors for which a complete 1978-2010 sample is available. Specification 5 includes the disaggregation of openness to trade, it is also estimated with a balanced sample, and by PCSE. One can see that the values of the estimated sector elasticity of labor demand is broadly robust to a change in estimation methodology, sample (sector selection), and control variables.

The aggregate elasticity of labor demand for the U.S. lies in the -0.23 to -0.17 interval according to our results. Hence, it is likely that the actual value is significantly below the standard Cobb-Douglas assumption of -0.33.

Furthermore, when relaxing this assumption and allowing for sector specific elasticities of substitution, we find that the elasticity of labor demand varies heterogeneously depending on the economic activity. Table 4 shows that 29 out of a total of 33 estimated elasticities are statistically different than zero (at a 15% level).

Take for instance specification 2, which is a two-stage FGLS unbalanced estimation for all sectors. The labor-demand elasticity we find for the real estate and business services sector (RE) is -1.91, far away from the aggregate estimation. But this is the highest value. If we consider manufacturing (MA) or the finance services sector (FI), the elasticities are -0.47 and -0.53 respectively, which is more than twofold the upper-bound aggregate value (-0.23). On the other hand, the estimated value for the retail trade, transportation and accommodation services sector (RT) is -0.14, lower than the aggregate value. The wage coefficient associated to community and social services sector (SE) is non-significant and hence statistically zero.

If we would gather only the homogeneous coefficients (HC) result, we would conclude that the elasticity of labor demand in the U.S. lies in the -0.23 to -0.17 range, and elaborate labor market policy accordingly. This paper shows that this procedure could be a seriously mistaken, since we would be missing out on the fact that the level of flexibility varies significantly across sectors. Then, labor market policy meant to increase employment could have very dissimilar outcomes. The bottom line is that sector-level analysis has to be taken into account in order to design effective policies.

Table 4 presents the results for Germany. All specifications have balanced samples (1993-2011). Specifications 3, 4 and 5 have a reduced sample of sectors based only on the statistical performance in specifications 1 and 2.

Specifications 1 and 2 for Sweden (Table 5) are unbalanced samples for all sectors (1972-2011). Then, specifications 3, 4 and 5 are balanced samples (1993-2011), with again a restriction of two sectors based on statistical performance in specifications 1 and 2.

The estimation results for Germany and Sweden present similar patterns than those of the U.S. The elasticity of labor demand at the sector level is in fact heterogeneous. Moreover, the values obtained are generally robust to sample period, estimation method and control variables. The same may be said about the ordinal ranking of sectors from the highest to the lowest estimated elasticity. Hence, the findings associated to the results for the U.S. are also robust to applying our empirical model to three different countries, with diverse labor market structures, size and degree of exposure to international trade.

In Germany's case, the HC estimated elasticity of labor demand ranges in the -0.72 to -0.23 interval. In turn, when adopting sector-level computations of the elasticity of labor demand, we find estimated values between -1.07 and -0.04. Taking again specification 2 as an example, the estimated elasticity under HC is -0.72, while the estimated sector labor demand elasticity for the finance services sector (FI) is -0.08 and the one for the retail trade sector (RT) is -0.09, both rather low. This low elasticity of labor demand in the retail sector is also found in the U.S. The most sensitive sectoral labor demand in Germany are agricultural activities (AG), where a 10% increase in the real wage may have a 8% reduction in sectoral labor demand. In general, Germany clearly presents lower ε_i than the U.S. (in absolute value) and in that sense it is, broadly, a less flexible labor market.

Table 4. **Germany** sectoral labor shares, elasticity of substitution and labor demand elasticity.

		1		**2**		**3**		**4**		**5**	
		PCSE		TS-FGLS		PCSE		TS-FGLS		PCSE	
	s	σ	ε	σ	ε	σ	ε	σ	ε	σ	ε
AG	0.29	0.88	-0.62	1.12	-0.80	0.96	-0.68	1.51	-1.07	1.03	-0.73
ME	0.45	1.05	-0.57	-0.10							
MA	0.70	0.66	-0.20	0.89	-0.27	0.68	-0.20	1.06	-0.32	0.69	-0.21
CO	0.75	1.62	-0.40	0.73	-0.18	1.47	-0.36	0.51	-0.12	1.71	-0.42
RT	0.66	0.13^b	-0.04^b	0.27	-0.09	0.06^b	-0.02^b	0.24	-0.08	0.02^b	-0.01^b
IT	0.57	0.25^b	-0.11^b	0.55	-0.24	0.21^b	-0.09^b	0.56	-0.24	0.17^b	-0.07^b
FI	0.66	0.18	-0.06	0.23	-0.08	0.17	-0.06	0.31	-0.11	0.12	-0.04
RE	0.22	0.79^b	-0.62^b	0.39^b	-0.30^b						
SE	0.74	0.68^b	-0.17^b	0.21^b	-0.05^b						
HC	0.58	0.78	-0.32	1.73	-0.72	0.58	-0.24	0.66	-0.27	0.56	-0.23
Sample		1993 2011		1993 2011		1993 2011		1993 2011		1993 2011	
Obs		171		171		114		114		114	

Note: PCSE = Panel-corrected standard errors. TSFGLS = two-stage feasible generalized least squares. No superscript = significance at 10% level. ^{b} = p-value > 0.10

In Table 5 we display the results for Sweden, where again we find heterogeneity of sector elasticity values when dropping the HC assumption. The HC estimated values range in the -0.31 to -0.19 interval, while sector-level values range from -1.06 to -0.04. As in Germany's case, the agriculture (AG) sector presents high labor-market flexibility. Taking the attention towards specification 2, as in the previous cases, the elasticity of labor demand in the manufacturing sector and the retail trade sector would be -0.49 and -0.55 respectively. We also find positive values of labor demand elasticity in the personal and community services sector (SE), as we did for one specification in the U.S. case, and the same possible explanation applies.

Table 5. **Sweden** sectoral labor shares, elasticity of substitution and labor demand elasticity.

	1			**2**		**3**		**4**		**5**	
	PCSE			TS-FGLS		PCSE		TS-FGLS		PCSE	
	s	σ	ε	σ	ε	σ	ε	σ	ε	σ	ε
AG	0.29	1.50	-1.06	2.06^b	-1.46^b	0.45	-0.32	0.45^b	-0.32^b	0.45	-0.32
ME	0.28	0.50^a	-0.36^a	0.55^b	-0.40^b	0.59	-0.42	0.56^b	-0.41^b	0.53^a	-0.38^a
MA	0.63	1.74	-0.65	1.30	-0.49	0.48	-0.18	0.77	-0.29	0.52	-0.19
CO	0.81	0.88	-0.17	0.82^b	-0.16^b	1.49	-0.29	1.19	-0.23	1.41	-0.27
RT	0.69	1.18	-0.37	1.77	-0.55	4.67	-1.45	2.10	-0.65	1.13	-0.35
IT	0.64	0.81	-0.29	0.67	-0.24	0.42^a	-0.15^a	1.00	-0.36	0.47^a	-0.17^a
FI	0.46	0.08^a	-0.04^a	0.20^b	-0.11^b	0.04^b	-0.02^b	0.47	-0.26	0.05^b	-0.03^b
RE	0.38	0.37^b	-0.23^b	-2.63	*						
SE	0.89	-1.24	*	-2.46	*						
HC	0.64	0.84	-0.31	0.73	-0.27	0.53	-0.19	0.52	-0.19	0.52	-0.19
Sample	1972 2011			1972 2011		1995 2010		1995 2010		1995 2010	
Obs	229			225		112		112		112	

Note: PCSE = Panel-corrected standard errors. TSFGLS = two-stage feasible generalized least squares. No superscript = wage-coefficient significance at 10% level.
^{a} = 0.10 < p-value < 0.15 ^{b} = p-value > 0.15 * = $\varepsilon_i > 0$

5.2 Exposure to international trade

We now turn the attention towards the effect of international trade on employment. As argued by related research, it would be expected that the higher the openness to trade, the higher the labor market flexibility. To control for this phenomenon, our specifications include different controls for the degree of openness to trade. In specifications 1 to 4 we include aggregate openness to trade (calculated as the ratio of total trade over GDP). Specification 5 includes a disaggregation of openness to trade in four types of merchandise: manufactures, services, agriculture and fuel.

It is reassuring to find that the computations of ε_i are quite robust to changes in the control for international trade since in specification 5 the estimated values lie around the same values found in the previous specifications (1 to 4) that include only total openness to trade as control.

But the net employment effects of higher openness are still under debate. Trade liberalization has been associated both with job destruction and job creation. It is a rule of thumb that exporting sectors would expand production and their demand for labor, while sectors exposed to competition with imports would reduce production and hence reduce the employment of labor (Jansen and Lee, 2007).

Associated to trade openness is international outsourcing, since the balance of payments includes the trade in services. As put by Amiti and Wei (2005), in the past, service sectors were considered virtually unaffected by trade. For example, "accountants did not fear that someone abroad would take their high-paying jobs", but this scenario has changed.

Tables 6 and 7 present the elasticity of the openness to trade variables in our specifications with respect to sector employment. The values in those tables are computed with the homogeneous coefficients estimations (HC). The reason for the utilization of the HC results is that the construction of the elasticity requires the openness coefficient (β_5), plus a global coefficient of persistence (β_1). Table 6 shows in each column the elasticity computed from specifications 1 to 4, and Table 7 refers to the results from specification 5 that disaggregates openness to trade.

Table 6. Long-run employment impact of international trade.

	1	**2**	**3**	**4**
U.S.	1.86***	-0.02	2.88***	-0.26
Germany	0.15	-0.72*	0.06	0.08
Sweden	1.70***	1.40***	1.17***	0.90**

Note: ***,** and * = significance at 1%, 5% and 10% level.

The degree of openness to international trade has a quite strong positive effect on sectoral employment in Sweden in all specifications (1 to 4; Table 6). The value of the long-run elasticity of this effect lies in the 0.90-1.70 range, thus the exposure to trade in Sweden is likely to be elastic with respect to employment. This positive effect also appears in the U.S., with even higher elasticities (1.86 and 2.88). In turn, the positive employment effect of trade cannot be detected in Germany's case. Not only that, but one specification for the case of Germany (number 2) suggests a negative effect of further openness to trade on employment.

The results for the U.S. and Sweden are consistent with recent evidence. Gozgor (2013), for example, includes four different measures of trade liberalization and globalization in a reduced-form unemployment equation and estimates the parameters for a panel of G7 countries, and all four proxies present a negative and significant effect on equilibrium unemployment. Also, Felbermayr *et al.* (2011), in panel and cross-sectional data specifications for several OECD countries, find a long-run reduction of unemployment associated to a higher exposure to international trade. In this context, Germany is an exception, where the exposure to trade has a non-positive effect on employment (that is, a low negative effect, or altogether inexistent).

The disaggregation of the openness to trade variable on to four sectors of merchandise brings further insights (Table 7). Germany still presents no significant effects of openness to trade on employment. The U.S. and Sweden present a robust positive effect on employment of further openness to trade in manufactures, with a similar elasticity than the aggregate case. The case of the degree of exposure to trade in services deserves particular discussion: it has a negative effect on employment in the U.S. and a positive effect in the case of Sweden.

Table 7. Disaggregated employment effect of openness to trade.

	Manufactures	Agriculture	Fuel	Services
U.S.	2.02***	-0.51	0.02	-0.83**
Germany	0.14	0.17	0.05	-0.44
Sweden	1.34***	0.09	-0.17	0.43*

Note: ***,** and * = significance at 1%, 5% and 10% level.

In a recent paper, Yanikkaya (2013) finds that a higher total openness to international trade has a negative effect on the growth rate of industrial employment and a positive effect on the growth rate of service employment. It follows that higher trade intensity may have diverse effects in different sectors.

In order to understand the opposite effect of higher trade intensity in services industries on employment in the U.S. and in Sweden we must take a look at what is different between these two countries. Figure 2 (panel e) shows that openness to trade in service industries grew in both countries over the last decades. Nevertheless, it has been always higher in the U.S., especially during the 1990s. Later, this difference has slightly declined (in 2011, openness to trade in services was 30% in the U.S. and 24% in Sweden). Also, most service sectors represented a higher proportion of employment in the U.S. than in Sweden over the sample period, with the exception of the information and communications sector[26].

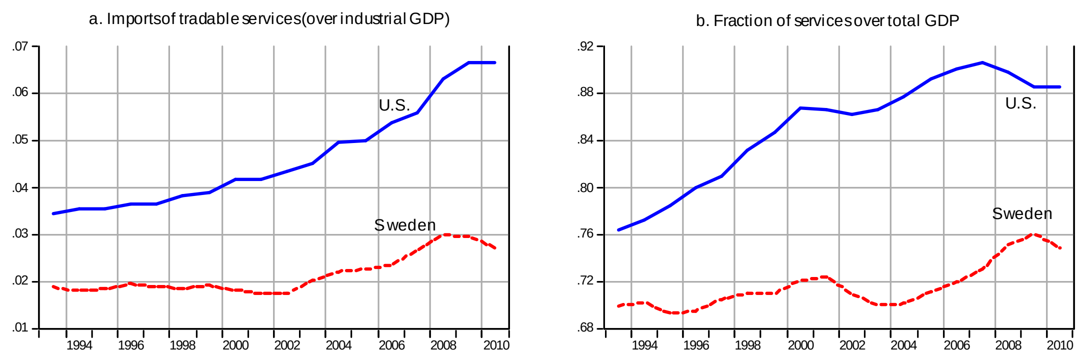
As aforementioned, in recent years there has been a strong debate over the effect of offshoring (and international outsourcing) on domestic employment. Crinò (2009) presents a complete review of empirical results: Amiti and Wei (2005) find a negative and significant effect of offshoring on employment in an industry-level study for the UK (1992-2000 period), OECD (2007) finds a positive but non-significant effect on employment in 24 industries across 17 OECD countries, and Crinò (2010) estimates the elasticity of service offshoring on domestic employment in 135 occupations in the U.S. over the 1997-2006

[26] Excluding the public sector (community services).

period and finds mostly negative effects on low and medium-skilled workers and a slim positive effect on high-skilled worker. The results of Crinò (2010) are consistent with our result for the U.S. The results in OECD's report of 2007 for a panel of 17 countries agree that there may be a positive effect in a given country.

So the answer must be in the role of trade in services in the U.S. and Sweden. Figure 3 presents the ratio of service imports over the industrial GDP (panel a) and the share that services represent on total GDP (panel b). The ratio of U.S. imports of business services over the industrial GDP has grown almost twofold over the 1993-2010 period. In turn, in Sweden this ratio had a flatter evolution, with moderate growth starting only after 2002. These contrasted evolutions combine with the fact that the dimension of the so-called structural change is higher in the U.S. Over the 1990s there was a steep growth of services fraction of GDP in the U.S. while it had a broadly constant evolution in Sweden.

Figure 3: Services in the U.S. and Sweden.



In all, further openness to trade in services in the U.S. can threaten domestic employment, mainly because it could translate in strong growth of imports, continuing the trend depicted over the last decade (Figure 3a). In Sweden, since domestic service industries do not represent as much of total income, and imports of services do not display a strong positive trend, higher levels of trade in services may favor employment.

It is important to recall that the effect of service offshoring is skill-biased and has different effects on high-skilled white-collar, low-skilled white-collar and blue-collar workers (Crinò, 2010). In that sense, certain sectors are more sensitive to openness in services industries than others, bringing different results for different economies depending on their economic structure.

5.3 Technical Change

Finally, we discuss the employment effect of technological progress. Recall that the time trend included in the model is a standard proxy of a constant-rate technical change (e.g. Antràs, 2004). The estimated coefficient associated to the time trend is $(1-\sigma)\lambda_N$ (by equations 10, 11 and 12). In almost all specifications $\sigma < 1$. Hence, if the estimated coefficient is negative, then technical change, in the context of the employment model in this paper, would have a negative effect on employment.

Table 8 presents the calculation of the long-run elasticity of the time trend with respect to employment (ε_{n-t}^{LR}), which quantifies the employment effect of constant-rate technical change, along with the annual growth rate of labor efficiency implied by our employment equations (λ_N). It is based on the homogeneous coefficient (HC) estimations for long-run computations as in the previous cases of trade effects. Cells left blank represent that the trend coefficient is not statistically different from zero.

Take for example specification 1 in the U.S. The estimated elasticity of substitution between factors is $\hat{\sigma} = 0.59$, the long-run elasticity of the time trend is -6.8% [in the context of equation (10), $\varepsilon_{n-t}^{LR} = \beta_4/(1-\beta_1)$] and then $\lambda_N = 17\%$ [since $\lambda_N = \varepsilon_{n-t}^{LR}/(1-\sigma)$].

Table 8. Employment effect of technical change.

	1		**2**		**3**		**4**		**5**	
	ε_{n-t}^{LR}	λ_N	ε_{n-t}^{LR}	λ_N	ε_{n-t}^{LR}	λ_N	ε_{n-t}^{LR}	λ_N	ε_{n-t}^{LR}	λ_N
U.S.	-6.8%	17%			-12%	22%			-3.2%	7%
Germany			2.8%	4%						
Sweden	-3.6%	22%	-2.7%	10%	-1.7%	4%				

Notes: ε_{n-t}^{LR} = long-run elasticity of time trend with respect to employment.

λ_N = growth rate of labor-related efficiency.

According to specification 5, a one percentage point increase in the rate of labor-related technical change would imply a 3.2% fall in employment. This labor-saving technical change grows at an annual constant rate of 7%. It follows that technological progress is a detrimental force of employment.

In Germany's case, specification 2 has a positive-sign long-run elasticity for the time trend, but in that same specification $\hat{\sigma} = 1.7 > 1$ (Table 4), and then the effect is also labor-saving (McAdam and Willman, 2013). The employment effect of higher technical change is 2.8%, and the growth rate of labor efficiency around 4%.

The long-run employment effect of technical change in Sweden is similar to that of Germany, both considerably smaller than in the U.S. case. In Sweden, an additional percentage point in the rate of technical change implies a fall in employment of between a 1% and a 4%. The growth rate of labor efficiency, in turn, is similar to those of the U.S.[27]

In this paper we estimate the growth rate of labor-related efficiency (λ_N), and not the full bias in technical change (i.e., the gap in the growth rates of labor- and capital-related efficiency). Nevertheless, labor-saving technical change was already identified for the U.S. by several studies as surveyed by Klump *et al.* (2012), so the direction of the employment effect of technological change may come as no surprise (see also Feldmann, 2013).

Note that technical change that presents a labor-saving effect in all three countries, is either smaller in Germany than in the U.S. and Sweden, or almost inexistent. This result may help in explaining the better performance of employment in Germany over the last decade.

6 Concluding remarks

This paper analyzes the heterogeneity in labor demand from two empirical perspectives. On the one hand, we provide calculations of the sector-level elasticities of labor demand and find that these values vary significantly across economic activities. If we rank sectors according to their estimated labor demand elasticity, some sectors are repeatedly among the most sensitive sector-level labor markets. For example, the IT sector in the U.S. and Germany, manufacturing in Germany and Sweden, and the mining and energy sectors in the U.S. and Sweden have the most elastic employment effects to changes in labor costs.

In contrast, the retail trade sector has the lowest elasticities in the U.S. and Germany, together with the finance services sector in Germany and Sweden. Notably, in our results we do not observe general criteria in terms of manufacturing having lower or higher elasticity than services sectors at this level of disaggregation. The standard economic knowledge is that more competitive sectors (or with a more competitive labor market, with less barriers like employment protection and benefits) are supposed to be more flexible and thus exert a higher elasticity of labor demand.

Policywise, the main implication of these results is that a one–size-fits-all approach to labor market policy will probably be inefficient. The reaction of employment to policy will be quite different depending on economic activities. According to our results, different economic sectors have different sensitivities in their demand for labor. Then, for a better outcome, labor market policy should be properly conceived taking into account sectoral

[27] In Germany's case, only one specification provides an statistically signicant time trend in the employment equation (number 2). That same specification gives a high elasticity of substitution between labor and capital, at least higher than those estimated in Sweden's case. That is why although having similar long-run elasticities of techinical change with respect to employment, Sweden has much higher efficieny growth rates than Germany.

particularities, in contrast to multifunctional recipes.

Also, Germany clearly presents lower sector-level elasticities of labor demand than the U.S. (in absolute value) and in that sense it is, in general, a less flexible labor market. Thus, looking at the performance that both labor markets had during the Great Recession, the following policy question arises: is flexibilization of European labor markets the answer? We join those that call for a rethinking of labor market policy, trying to go beyond labor market flexibilization and place more emphasis on investment, technology, productivity, and related issues. The expansion of productive capacity, via targeted investment and labor-enhancing innovation, is probably a better long-run solution than social welfare deconstruction.

On the other hand, we investigate the employment effects of higher exposure to international trade. We do this by including the degree of openness to trade in the empirical employment equation, first in its aggregate version, and later disaggregating openness to trade into four variables according to four types of merchandise: manufactures, services, agriculture and fuel. Openness to trade presents a non-negative effect on employment (neutral in Germany and positive in the U.S. and Sweden). But new insights come along disaggregating aggregate openness to trade. Higher trade in manufactures has a positive effect on employment, as expected, in the U.S. and Sweden. Interestingly, a larger degree of openness to trade in services has a negative effect on employment in the U.S. and a positive effect in Sweden.

We believe that this result may be associated to the growing importance of imported services in the U.S. economy and the important role that service industries already play, in contrast to Sweden, where the services share of the economy is still not as large and there may be room to increase trade in services and boost domestic employment. The skill-biased effect of offshoring and international outsourcing is a phenomenon that should be considered.

Lastly, this paper also verifies the presence of a negative employment effect of technical change in the three countries studied. The long-run elasticity of constant-rate technical change with respect to employment is between 3% and 12% in the U.S., less than 4% in Sweden, and around a 3% in Germany. The annual growth rate of labor efficiency implied by the model is 7% to 22% in the U.S., 4% to 22% in Sweden, and 0% to 4% in Germany. The fact that this growth rate is small(er) in Germany may help in explaining its better employment performance over the last decade. The main drivers of growth in economic modeling are total factor productivity (or efficiency growth) and capital intensity. In each economy, one of these drivers prevails over the other one, or they combine together. This fact may contribute to explain the difference in the effect of efficiency growth on employment that we observe in our results between the three countries, and should be further investigated.

Future research should explore ways to estimate the elasticity of labor demand from the empirical model directly instead of indirectly computing it via the estimated elasticity of substitution. Also, disaggregated effects of openness to trade and technical change for each one of the nine sectors in ISIC Revision 4 should be undertaken as a methodological challenge. This paper uses a semi-pooling approach because it is the methodology that suits best our sample. With higher T [or even higher N via disaggregation in industries like, for example, Young (2013)], empirical methodology should be explored in order to have n_{t-1i} and w_{ti} for all i sectors (or industries). This should also allow to estimate the employment effects of technical change and trade exposure at the sector level.

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Appendix

Table A1. United States. Semi-pooled model.

[1] PCSE				[2] TS-FGLS			
c	0.636 [0.290]			c	0.171 [0.863]		
Δn_{t-1}	0.227 [0.003]			Δn_{t-1}	0.333 [0.000]		
Δn_{t-2}	-0.184 [0.027]			Δn_{t-2}	-0.178 [0.005]		
va_t	0.211 [0.000]			va_t	0.301 [0.001]		
t	-0.007 [0.000]			t	-0.001 [0.738]		
op_t	0.179 [0.000]			op_t	0.017 [0.688]		
n_{t-1}^{AG}	0.535 [0.000]	w_t^{AG}	-0.171 [0.000]	n_{t-1}^{AG}	0.538 [0.191]	w_t^{AG}	-0.226 [0.006]
n_{t-1}^{ME}	0.631 [0.000]	w_t^{ME}	-0.314 [0.000]	n_{t-1}^{ME}	0.623 [0.000]	w_t^{ME}	-0.232 [0.012]
n_{t-1}^{MA}	0.697 [0.000]	w_t^{MA}	-0.287 [0.000]	n_{t-1}^{MA}	0.706 [0.000]	w_t^{MA}	-0.384 [0.000]
n_{t-1}^{CO}	0.720 [0.000]	w_t^{CO}	-0.273 [0.011]	n_{t-1}^{CO}	0.616 [0.000]	w_t^{CO}	-0.320 [0.030]
n_{t-1}^{RT}	0.536 [0.000]	w_t^{RT}	-0.086 [0.201]	n_{t-1}^{RT}	0.502 [0.003]	w_t^{RT}	-0.246 [0.053]
n_{t-1}^{IT}	0.860 [0.000]	w_t^{IT}	-0.148 [0.102]	n_{t-1}^{IT}	0.730 [0.031]	w_t^{IT}	-0.329 [0.015]
n_{t-1}^{FI}	0.772 [0.000]	w_t^{FI}	-0.217 [0.000]	n_{t-1}^{FI}	0.706 [0.000]	w_t^{FI}	-0.355 [0.000]
n_{t-1}^{RE}	0.857 [0.000]	w_t^{RE}	-0.260 [0.010]	n_{t-1}^{RE}	0.830 [0.000]	w_t^{RE}	-0.502 [0.000]
n_{t-1}^{SE}	0.845 [0.000]	w_t^{SE}	0.732 [0.001]	n_{t-1}^{SE}	0.714 [0.000]	w_t^{SE}	-0.095 [0.697]
Unbalanced Sample: 1978-2010				Unbalanced Sample: 1978-2010			
Total obs: 230				Total obs: 229			

Notes: p-values in brackets; Instruments: Δn_{t-1} Δn_{t-2} va_{t-1} op_{t-1} t and n_{it-1} w_{it-1} $\forall i$. $c =$ intercept.

Table A2. United States. Semi-pooled model.

[3] PCSE				[4] TS-FGLS			
c	1.017 [0.085]			c	0.040 [0.974]		
Δn_{t-1}	0.268 [0.002]			Δn_{t-1}	0.327 [0.006]		
Δn_{t-2}	-0.218 [0.016]			Δn_{t-2}	-0.190 [0.003]		
va_t	0.209 [0.000]			va_t	0.283 [0.020]		
t	-0.008 [0.001]			t	0.001 [0.841]		
op_t	0.192 [0.000]			op_t	-0.021 [0.705]		
n_{t-1}^{AG}	0.533 [0.005]	w_t^{AG}	-0.163 [0.000]	n_{t-1}^{AG}	0.604 [0.141]	w_t^{AG}	-0.205 [0.013]
n_{t-1}^{MA}	0.685 [0.000]	w_t^{MA}	-0.275 [0.000]	n_{t-1}^{MA}	0.713 [0.000]	w_t^{MA}	-0.367 [0.000]
n_{t-1}^{CO}	0.737 [0.000]	w_t^{CO}	-0.294 [0.008]	n_{t-1}^{CO}	0.652 [0.000]	w_t^{CO}	-0.277 [0.168]
n_{t-1}^{FI}	0.788 [0.000]	w_t^{FI}	-0.202 [0.001]	n_{t-1}^{FI}	0.699 [0.000]	w_t^{FI}	-0.322 [0.001]
n_{t-1}^{RE}	0.862 [0.000]	w_t^{RE}	-0.220 [0.041]	n_{t-1}^{RE}	0.854 [0.000]	w_t^{RE}	-0.538 [0.000]
Balanced Sample: 1978-2010				Balanced Sample: 1978-2010			
Total obs: 165				Total obs: 165			

Notes: p-values in brackets; Instruments: Δn_{t-1} Δn_{t-2} va_{t-1} op_{t-1} t and n_{it-1} w_{it-1} $\forall i$. $c =$ intercept.

Table A3. United States. FE model (HC).

[1] PCSE		[2] TS-FGLS		[3] PCSE		[4] TS-FGLS	
c	0.454 [0.114]	c	0.416 [0.083]	c	0.425 [0.202]	c	0.130 [0.731]
n_{t-1}	0.912 [0.000]	n_{t-1}	0.939 [0.000]	n_{t-1}	0.935 [0.000]	n_{t-1}	0.957 [0.000]
Δn_{t-1}	0.330 [0.000]	Δn_{t-1}	0.520 [0.000]	Δn_{t-1}	0.353 [0.000]	Δn_{t-1}	0.522 [0.000]
Δn_{t-2}	-0.233 [0.005]	Δn_{t-2}	-0.275 [0.000]	Δn_{t-2}	-0.285 [0.003]	Δn_{t-2}	-0.258 [0.004]
va_t	0.072 [0.007]	va_t	0.019 [0.469]	va_t	0.057 [0.060]	va_t	0.023 [0.538]
t	-0.006 [0.000]	t	0.0001 [0.933]	t	-0.008 [0.000]	t	-0.001 [0.815]
op_t	0.163 [0.000]	op_t	-0.001 [0.968]	op_t	0.186 [0.000]	op_t	-0.011 [0.823]
w_t	-0.052 [0.006]	w_t	-0.030 [0.131]	w_t	-0.028 [0.150]	w_t	-0.014 [0.517]
Unbalanced Sample: 1978-2010				Balanced Sample: 1978-2010			
Total obs: 230		Total obs: 229		Total obs: 165			

Notes: p-values in brackets; Instruments in [2] and [4]: n_{it-1} Δn_{t-1} Δn_{t-2} va_{t-1} op_{t-1} t w_{it-1}
$c =$ intercept.

Table A4. United States. Specification [5]. PCSE.

[5] Semi-pooled.				[5] HC	
c	1.366 [0.028]			c	0.191 [0.618]
va_t	0.209 [0.000]			n_{t-1}	0.905 [0.000]
t	-0.002 [0.167]			va_t	0.065 [0.054]
opm_t	0.217 [0.000]			t	-0.003 [0.042]
ops_t	-0.094 [0.008]			opm_t	0.192 [0.000]
opa_t	-0.095 [0.006]			ops_t	-0.079 [0.046]
opf_t	0.008 [0.557]			opa_t	-0.048 [0.147]
n_{t-1}^{AG}	0.471 [0.022]	w_t^{AG}	-0.212 [0.000]	opf_t	-0.002 [0.898]
n_{t-1}^{MA}	0.595 [0.000]	w_t^{MA}	-0.368 [0.000]	w_t	-0.049 [0.031]
n_{t-1}^{CO}	0.624 [0.000]	w_t^{CO}	-0.315 [0.006]		
n_{t-1}^{FI}	0.621 [0.000]	w_t^{FI}	-0.205 [0.002]		
n_{t-1}^{RE}	0.793 [0.000]	w_t^{RE}	-0.236 [0.089]		
Balanced Sample: 1978-2010				Balanced Sample: 1978-2010	
Total obs: 165				Total obs: 165	

Notes: p-values in brackets. c = intercept.

Table A5. Germany. Semi-pooled model.

[1] PCSE				[2] TS-FGLS			
c	0.299 [0.217]			c	0.170 [0.541]		
Δn_{t-1}	0.328 [0.000]			Δn_{t-1}	0.352 [0.000]		
va_t	0.145 [0.000]			va_t	0.206 [0.000]		
t	−0.003 [0.075]			t	0.010 [0.581]		
op_t	0.036 [0.191]			op_t	0.010 [0.743]		
n_{t-1}^{AG}	0.909 [0.000]	w_t^{AG}	−0.080 [0.017]	n_{t-1}^{AG}	0.838 [0.000]	w_t^{AG}	−0.182 [0.000]
n_{t-1}^{ME}	0.918 [0.000]	w_t^{ME}	−0.086 [0.027]	n_{t-1}^{ME}	0.908 [0.000]	w_t^{ME}	0.009 [0.850]
n_{t-1}^{MA}	0.677 [0.000]	w_t^{MA}	−0.213 [0.008]	n_{t-1}^{MA}	0.593 [0.000]	w_t^{MA}	−0.364 [0.000]
n_{t-1}^{CO}	0.863 [0.000]	w_t^{CO}	−0.222 [0.013]	n_{t-1}^{CO}	0.801 [0.000]	w_t^{CO}	−0.146 [0.020]
n_{t-1}^{RT}	0.646 [0.000]	w_t^{RT}	−0.044 [0.522]	n_{t-1}^{RT}	0.483 [0.000]	w_t^{RT}	−0.140 [0.020]
n_{t-1}^{IT}	0.743 [0.000]	w_t^{IT}	−0.064 [0.248]	n_{t-1}^{IT}	0.737 [0.000]	w_t^{IT}	−0.146 [0.020]
n_{t-1}^{FI}	0.658 [0.000]	w_t^{FI}	−0.063 [0.018]	n_{t-1}^{FI}	0.577 [0.000]	w_t^{FI}	−0.098 [0.026]
n_{t-1}^{RE}	0.913 [0.000]	w_t^{RE}	−0.069 [0.679]	n_{t-1}^{RE}	0.876 [0.000]	w_t^{RE}	−0.048 [0.803]
n_{t-1}^{SE}	0.868 [0.000]	w_t^{SE}	−0.091 [0.287]	n_{t-1}^{SE}	0.777 [0.000]	w_t^{SE}	−0.048 [0.523]
Balanced Sample: 1993-2011				Balanced Sample: 1993-2011			
Total obs: 171				Total obs: 171			

Notes: p-values in brackets; Instruments: Δn_{t-1} va_{t-1} op_{t-1} t and n_{it-1} w_{it-1} $\forall i$. c = intercept.

Table A6. Germany. Semi-pooled model.

[3] PCSE				[4] TS-FGLS			
c	0.813 [0.016]			c	0.810 [0.023]		
Δn_{t-1}	0.373 [0.000]			Δn_{t-1}	0.419 [0.000]		
va_t	0.145 [0.000]			va_t	0.297 [0.000]		
t	-0.003 [0.087]			t	-0.002 [0.466]		
op_t	0.053 [0.165]			op_t	0.024 [0.589]		
n_{t-1}^{AG}	0.918 [0.005]	w_t^{AG}	-0.078 [0.037]	n_{t-1}^{AG}	0.839 [0.141]	w_t^{AG}	-0.241 [0.000]
n_{t-1}^{MA}	0.690 [0.000]	w_t^{MA}	-0.210 [0.014]	n_{t-1}^{MA}	0.518 [0.000]	w_t^{MA}	-0.514 [0.000]
n_{t-1}^{CO}	0.854 [0.000]	w_t^{CO}	-0.215 [0.020]	n_{t-1}^{CO}	0.701 [0.000]	w_t^{CO}	-0.151 [0.025]
n_{t-1}^{RT}	0.629 [0.000]	w_t^{RT}	-0.023 [0.772]	n_{t-1}^{RT}	0.220 [0.285]	w_t^{RT}	-0.184 [0.012]
n_{t-1}^{IT}	0.731 [0.000]	w_t^{IT}	-0.058 [0.336]	n_{t-1}^{IT}	0.628 [0.000]	w_t^{IT}	-0.210 [0.004]
n_{t-1}^{FI}	0.618 [0.000]	w_t^{FI}	-0.065 [0.022]	n_{t-1}^{FI}	0.562 [0.001]	w_t^{FI}	-0.137 [0.052]
Balanced Sample: 1993-2011				Balanced Sample: 1993-2011			
Total obs: 114				Total obs: 114			

Notes: p-values in brackets; Instruments: Δn_{t-1} va_{t-1} op_{t-1} t and n_{it-1} w_{it-1} $\forall i$. c = intercept.

Table A7. Germany. FE model (HC).

[1] PCSE		[2] TS-FGLS		[3] PCSE		[4] TS-FGLS	
c	-0.236 [0.130]	c	-1.125 [0.026]	c	-0.022 [0.903]	c	0.042 [0.767]
n_{t-1}	0.906 [0.000]	n_{t-1}	0.824 [0.000]	n_{t-1}	0.853 [0.000]	n_{t-1}	0.832 [0.000]
Δn_{t-1}	0.366 [0.000]	Δn_{t-1}	0.013 [0.945]	Δn_{t-1}	0.367 [0.000]	Δn_{t-1}	0.399 [0.000]
va_t	0.104 [0.000]	va_t	0.289 [0.005]	va_t	0.126 [0.000]	va_t	0.142 [0.001]
t	-0.001 [0.410]	t	0.005 [0.107]	t	-0.001 [0.422]	t	-0.001 [0.518]
op_t	0.014 [0.603]	op_t	-0.128 [0.068]	op_t	0.009 [0.793]	op_t	0.014 [0.771]
w_t	-0.073 [0.000]	w_t	-0.305 [0.009]	w_t	-0.085 [0.001]	w_t	-0.110 [0.001]
Δw_t	0.020 [0.298]	Δw_t	-1.032 [0.040]				
Balanced Sample: 1993-2011				Balanced Sample: 1993-2011			
Total obs: 171				Total obs: 114			

Notes: p-values in brackets; Instruments in [2] and [4]: n_{it-1} Δn_{t-1} va_{t-1} op_{t-1} t w_{it-1} w_{it-2}
$c =$ intercept.

Table A8. Germany. Specification [5]. PCSE.

[5] Semi-pooled.				[5] HC	
c	0.631 [0.112]			c	−0.034 [0.903]
Δn_{t-1}	0.397 [0.000]			n_{t-1}	0.857 [0.000]
va_t	0.145 [0.000]			Δn_{t-1}	0.379 [0.000]
t	0.001 [0.636]			va_t	0.120 [0.000]
opm_t	0.048 [0.176]			t	−0.001 [0.814]
ops_t	−0.121 [0.129]			opm_t	0.020 [0.542]
opa_t	−0.036 [0.333]			ops_t	−0.062 [0.417]
opf_t	0.017 [0.178]			opa_t	0.025 [0.437]
n_{t-1}^{AG}	0.927 [0.000]	w_t^{AG}	−0.075 [0.057]	opf_t	0.007 [0.559]
n_{t-1}^{MA}	0.658 [0.000]	w_t^{MA}	−0.238 [0.015]	w_t	−0.080 [0.002]
n_{t-1}^{CO}	0.862 [0.000]	w_t^{CO}	−0.235 [0.016]		
n_{t-1}^{RT}	0.602 [0.001]	w_t^{RT}	−0.009 [0.925]		
n_{t-1}^{IT}	0.706 [0.000]	w_t^{IT}	−0.050 [0.400]		
n_{t-1}^{FI}	0.559 [0.001]	w_t^{FI}	−0.055 [0.069]		
Balanced Sample: 1993-2011				Balanced Sample: 1993-2011	
Total obs: 114				Total obs: 114	

Notes: p-values in brackets. c = intercept.

Table A9. Sweden. Semi-pooled model.

[1] PCSE				[2] TS-FGLS			
c	-0.398 [0.270]			c	-0.398 [0.483]		
Δn_{t-1}	0.247 [0.000]			Δn_{t-1}	0.217 [0.001]		
va_t	0.248 [0.000]			va_t	0.161 [0.007]		
t	-0.003 [0.002]			t	-0.003 [0.007]		
op_t	0.151 [0.000]			op_t	0.152 [0.005]		
n_{t-1}^{AG}	0.942 [0.000]	w_t^{AG}	-0.086 [0.004]	n_{t-1}^{AG}	0.976 [0.000]	w_t^{AG}	-0.086 [0.004]
n_{t-1}^{ME}	0.678 [0.001]	w_t^{ME}	-0.162 [0.140]	n_{t-1}^{ME}	0.681 [0.010]	w_t^{ME}	-0.162 [0.140]
n_{t-1}^{MA}	0.829 [0.000]	w_t^{MA}	-0.297 [0.000]	n_{t-1}^{MA}	0.831 [0.000]	w_t^{MA}	-0.297 [0.000]
n_{t-1}^{CO}	0.720 [0.000]	w_t^{CO}	-0.245 [0.002]	n_{t-1}^{CO}	0.852 [0.000]	w_t^{CO}	-0.245 [0.002]
n_{t-1}^{RT}	0.706 [0.000]	w_t^{RT}	-0.347 [0.000]	n_{t-1}^{RT}	0.856 [0.000]	w_t^{RT}	-0.347 [0.000]
n_{t-1}^{IT}	0.671 [0.000]	w_t^{IT}	-0.148 [0.102]	n_{t-1}^{IT}	0.731 [0.000]	w_t^{IT}	-0.148 [0.102]
n_{t-1}^{FI}	0.772 [0.000]	w_t^{FI}	-0.217 [0.000]	n_{t-1}^{FI}	0.582 [0.010]	w_t^{FI}	-0.217 [0.000]
n_{t-1}^{RE}	0.857 [0.000]	w_t^{RE}	-0.260 [0.010]	n_{t-1}^{RE}	0.552 [0.125]	w_t^{RE}	-0.260 [0.010]
n_{t-1}^{SE}	0.743 [0.000]	w_t^{SE}	0.732 [0.001]	n_{t-1}^{SE}	0.743 [0.000]	w_t^{SE}	0.732 [0.001]
Unbalanced Sample: 1972-2011				Unbalanced Sample: 1972-2011			
Total obs: 229				Total obs: 225			

Notes: p-values in brackets; Instruments: Δn_{t-1} va_{t-1} op_{t-1} t and n_{it-1} w_{it-1} $\forall i$. $c =$ intercept.

Table A10. Sweden. Semi-pooled model.

[3] PCSE				[4] TS-FGLS			
c	1.639 [0.021]			c	1.635 [0.428]		
Δn_{t-1}	0.242 [0.009]			Δn_{t-1}	0.232 [0.018]		
va_t	0.148 [0.014]			va_t	0.302 [0.010]		
t	-0.007 [0.005]			t	-0.005 [0.080]		
op_t	0.299 [0.000]			op_t	0.283 [0.004]		
n_{t-1}^{AG}	0.699 [0.000]	w_t^{AG}	-0.134 [0.037]	n_{t-1}^{AG}	-0.136 [0.938]	w_t^{AG}	-0.517 [0.458]
n_{t-1}^{ME}	0.633 [0.004]	w_t^{ME}	-0.215 [0.052]	n_{t-1}^{ME}	0.153 [0.805]	w_t^{ME}	-0.478 [0.224]
n_{t-1}^{MA}	0.505 [0.000]	w_t^{MA}	-0.236 [0.000]	n_{t-1}^{MA}	0.505 [0.000]	w_t^{MA}	-0.381 [0.000]
n_{t-1}^{CO}	0.660 [0.000]	w_t^{CO}	-0.506 [0.002]	n_{t-1}^{CO}	0.499 [0.004]	w_t^{CO}	-0.594 [0.007]
n_{t-1}^{RT}	0.938 [0.000]	w_t^{RT}	-0.289 [0.002]	n_{t-1}^{RT}	0.740 [0.010]	w_t^{RT}	-0.545 [0.001]
n_{t-1}^{IT}	0.692 [0.000]	w_t^{IT}	-0.130 [0.148]	n_{t-1}^{IT}	0.641 [0.000]	w_t^{IT}	-0.357 [0.011]
n_{t-1}^{FI}	0.342 [0.094]	w_t^{FI}	-0.024 [0.676]	n_{t-1}^{FI}	0.558 [0.082]	w_t^{FI}	-0.209 [0.059]
Balanced Sample: 1995-2010				Balanced Sample: 1995-2010			
Total obs: 112				Total obs: 112			

Notes: p-values in brackets; Instruments: Δn_{t-1} va_{t-1} op_{t-1} t and n_{it-1} w_{it-1} $\forall i$. $c =$ intercept.

Table A11. Sweden. FE model (HC).

[1] PCSE		[2] TS-FGLS		[3] PCSE		[4] TS-FGLS	
c	0.568 [0.002]	c	0.608 [0.000]	c	1.359 [0.001]	c	1.772 [0.000]
n_{t-1}	0.888 [0.000]	n_{t-1}	0.925 [0.000]	n_{t-1}	0.761 [0.000]	n_{t-1}	0.782 [0.000]
Δn_{t-1}	0.300 [0.000]	Δn_{t-1}	0.263 [0.000]	Δn_{t-1}	0.187 [0.036]	Δn_{t-1}	0.123 [0.180]
va_t	0.062 [0.001]	va_t	0.018 [0.430]	va_t	0.075 [0.080]	va_t	0.014 [0.831]
t	−0.004 [0.000]	t	−0.002 [0.039]	t	−0.004 [0.011]	t	−0.001 [0.621]
op_t	0.190 [0.000]	op_t	0.104 [0.006]	op_t	0.279 [0.000]	op_t	0.196 [0.017]
w_t	−0.094 [0.000]	w_t	−0.055 [0.009]	w_t	−0.127 [0.000]	w_t	−0.113 [0.014]
Balanced Sample: 1995-2011				Balanced Sample: 1995-2011			
Total obs: 112				Total obs: 112			

Notes: p-values in brackets; Instruments in [2] and [4]: n_{it-1} Δn_{t-1} va_{t-1} op_{t-1} t w_{it-1}
c = intercept.

Table A12. Sweden. Specification [5]. PCSE.

[5] Semi-pooled.				[5] HC	
c	1.758 [0.029]			c	1.518 [0.002]
Δn_{t-1}	0.188 [0.062]			n_{t-1}	0.793 [0.000]
va_t	0.150 [0.036]			Δn_{t-1}	0.179 [0.064]
t	−0.004 [0.310]			va_t	0.050 [0.268]
opm_t	0.162 [0.076]			t	−0.001 [0.801]
ops_t	0.142 [0.011]			opm_t	0.278 [0.001]
opa_t	−0.054 [0.456]			ops_t	0.089 [0.089]
opf_t	−0.020 [0.424]			opa_t	0.018 [0.790]
n_{t-1}^{AG}	0.668 [0.000]	w_t^{AG}	−0.150 [0.021]	opf_t	−0.035 [0.158]
n_{t-1}^{ME}	0.680 [0.004]	w_t^{MA}	−0.170 [0.152]	w_t	−0.109 [0.000]
n_{t-1}^{MA}	0.527 [0.000]	w_t^{MA}	−0.245 [0.000]		
n_{t-1}^{CO}	0.677 [0.000]	w_t^{CO}	−0.457 [0.007]		
n_{t-1}^{RT}	0.785 [0.002]	w_t^{RT}	−0.243 [0.018]		
n_{t-1}^{IT}	0.689 [0.000]	w_t^{IT}	−0.146 [0.130]		
n_{t-1}^{FI}	0.334 [0.167]	w_t^{FI}	−0.033 [0.596]		
Balanced Sample: 1995-2010				Balanced Sample: 1995-2010	
Total obs: 112				Total obs: 112	

Notes: p-values in brackets. c = intercept.