Trade and Unemployment: What do the data say?*

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Abstract

This paper documents a robust empirical regularity: in the long-run, higher trade openness is associated to a lower structural rate of unemployment. We establish this fact using: (i) panel data from 20 OECD countries, (ii) cross-sectional data on a larger set of countries. The time structure of the panel data allows us to control for unobserved heterogeneity, whereas cross-sectional data make it possible to instrument openness by its geographical component. In both setups, we purge the data from business cycle effects, include a host of institutional and geographical variables, and control for within-country trade. Our main finding is robust to various definitions of unemployment rates and openness measures. Our preferred specification suggests that a 10 percent increase in total trade openness reduces aggregate unemployment by about three quarters of one percentage point.

Keywords: international trade, real openness, unemployment, GMM models, IV estimation.
JEL codes: F16, E24, J6

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1. Introduction

Does exposure to international trade create or destroy jobs? In the short run, trade liberalization increases job turnover as workers are reallocated from shrinking to expanding sectors.[1] Empirical evidence suggests that those adjustments temporarily raise frictional unemployment at the aggregate level, as documented by Trefler (2004) for the case of NAFTA. On the other hand, the long run effect of trade liberalization on the equilibrium rate of unemployment is less clear.[2]

A burgeoning literature introduces labor market imperfections into workhorse models of international trade. Most papers conclude that trade openness matters for the equilibrium rate of unemployment; however, the sign of the relationship differs across papers. Blanchard (2006) talks about an "*overabundance of theories*" of wage setting and unemployment. Interacted with different explanations for international trade (comparative advantage versus product differentiation models), the number of possible theoretical frameworks is large. Brecher (1974) and Davis (1998) incorporate minimum wages into Heckscher-Ohlin models and find that trade liberalization can exacerbate unemployment. Davidson and Matusz (1988, 1999) introduce frictional unemployment in models of comparative advantage and find that the sign of the relationship depends on a comparison of capital-labor endowments across countries. Egger and Kreickemeier (2009) introduce fair wages into a model with increasing returns to scale and find that trade liberalization can increase unemployment. Felbermayr, Prat, and Schmerer (2009) introduce search frictions into a similar trade model and find that unemployment is likely to be decreasing in the degree of openness. Helpman and Itshoki (2008) also use the search-matching approach, but combine comparative advantage motives and increasing returns to scale. They find that globalization can increase unemployment.[3]

The state of the theoretical literature therefore suggests turning towards an empirical assessment. As stated by Davidson and Matusz (2004), whether trade affects the level of equilibrium unemployment is "*primarily an empirical issue* ". Yet, "*there is very little empirical work on the aggregate employment effects of trade policies* ". This paper attempts to shed some light on this question. Rather than testing a specific theoretical model, it presents some robust facts about the relationship between the rate of unemployment and openness in cross-sections of countries. There are two important challenges on the way. First, published data on unemployment rates are notoriously unreliable, with *measurement bias* systematically related to determinants

[1]See Bernard, Redding and Schott (2007) for recent evidence.

[2]Paul Krugman (1993) famously argues that "... *the level of employment is a macroeconomic issue, depending in the short run on aggregate demand and depending in the long run on the natural rate of unemployment, with microeconomic policies like tariffs having little net effect.*" However, theoretical considerations, as well as empirical evidence suggest that at least *some* microeconomic policies–such as product market regulation–do affect the structural rate of unemployment; see Blanchard and Giavazzi (2003) for the theoretical argument and Bassanini and Duval (2006) for a survey of the empirics.

[3]The theoretical literature is large and quickly growing; our short summary cannot be but a very incomplete list of papers.

of unemployment. Moreover, “good data ”on labor market regulation is available only for a few countries. Second, the incentive for politicians to erect trade barriers as a response to unemployment shocks, may introduce a negative *spurious correlation* between unemployment and openness. If the timing of trade liberalization and labor market reform coincide, domestic demand shocks will concurrently reduce unemployment and increase imports.

We tackle the data quality problem by focusing on two different samples. We start with a high-quality data set of 20 rich OECD countries, provided by Bassanini and Duval (2006, 2009). Great efforts have been made at the OECD to construct unemployment rates and indicators of various labor market institutions with meaningful time and cross-sectional variance. In a second step, we use a lower-quality cross-section of countries, for which we average yearly unemployment rates from various data sets such as provided by the World Bank, the International Labor Organization, the International Monetary Fund, or the CIA and draw on labor market variables provided by Botero et al. (2004). To avoid spurious results, we do our best to purge the data from business cycle effects and we use a comprehensive set of variables to control for labor market institutions. To address simultaneity bias in the OECD panel, we use various GMM-based techniques and exploit the time dimension of the data to construct instruments. In the cross-section, we use the geographical component of trade openness as an instrument.

Across different econometric models, different specifications, and different data sources, we are able to flesh out an important and robust result: the structural rate of unemployment is a non-increasing function of openness to trade. In the largest share of our regressions, higher trade openness actually decreases unemployment. In some exercises, it is irrelevant. It never turns out to be positively correlated with unemployment. We find the following additional results. (i) There is no evidence that the effect of openness on unemployment is biased upwards due to endogeneity. Quite to the contrary, we find that OLS yields a negative bias, which signals that attenuation bias due to non-systematic measurement error in the openness measure (which biases results to zero) dwarves the endogeneity bias. (ii) It is important to adjust the openness measures for differences in the relative prices of non-traded goods, as suggested by Alcalá and Ciccone (2004) in the context of cross-country growth regressions. In particular, the unadjusted openness measure tends to exaggerate the effect of openness on unemployment.[4] (iii)It appears that the reduction in aggregate unemployment is primarily due to lower unemployment of high-skilled workers.

Related literature. Apart from the theoretical literature discussed above, our exercise is closely related to two important strands of empirical research. First, labor economists have long estimated cross-country unemployment regressions, usually based on panel data for a restricted

[4]Note that this issue is of much less concern in our panel analysis, where we can effectively control for the time-invariant component of cross-country variation in relative prices.

sample of rich OECD countries. Following Blanchard and Wolfers' (2000) seminal paper, the literature is mainly concerned with the explanatory power of labor market institutions and macroeconomic shocks. Nickell et al. (2005) provide a recent example of this approach, whereas Bassanini and Duval (2006) present a comprehensive survey. The terms "international trade", "openness" or "globalization" do not appear in their comprehensive 130 pages study. Hence, it appears to us that the role of international trade in cross-country regressions has not yet been thoroughly addressed.[5] To connect our results with previous research, we closely follow the received methodology since we use similar data, econometric techniques and specifications. To the best of our knowledge, this paper is the first to systematically assess the role of trade openness for unemployment within the context of standard cross-country unemployment regressions for OECD countries.[6] Surprisingly enough, the influence of trade turns out to be much more robust than that of many labor market institutions.

We also incorporate insights from the large empirical literature about the effect of trade openness on per capita income. Frankel and Romer (1999) have proposed an instrumentation strategy based on geography which is, as a matter of fact, applicable only in cross-sections. The consensus is that the positive effect of openness on per capita income is not robust to seemingly unrelated geographical controls, such as the distance to equator.[7] Their paper has triggered a debate on the relative importance of trade, institutions, and the common underlying exogenous driver, geography. Prolonging this line of investigation, a recent paper by Dutt et al. (2009) test specific implications of the Davidson and Matusz (1999) model using cross-country regressions and a geography-based instrument. Although their sample, data sources and methodology are different, their results are qualitatively in line with ours. Interestingly, our own IV estimates, much inspired by the approach of Alcalá and Cicone (2004), suggest a negative relationship between openness and unemployment that is robust to inclusion of variables such as distance to equator or general institutional controls.

Structure of the paper. In section 2. we provide a brief first glance at the data. We identify two key concerns about data quality and endogeneity bias. This motivates section 3., where we sketch the empirical strategy for our different data sets. Section 4. contains our core results on the trade-unemployment relation. Section 5. presents robustness checks and discusses the role of TFP as the channel through which openness affects unemployment. Finally, section 6. concludes.

[5] Scarpetta (1996) uses an index measuring the pervasiveness of trade restrictions to proxy the intensity of competition. One also should add that many papers interact terms-of-trade shocks with labor market variables. However, they do not use the level of openness as an independent covariate. Boulhol (2008) interacts trade openness with labor market institutions, but does not address the endogeneity problem.

[6] The report of the European Economic Advisory Group at CESifo (2008) also includes some cross-country regressions of unemployment rates on openness, but does not attempt to sort out correlation from causality.

[7] See, for example, Rodriguez and Rodrik, 2000.

2. A descriptive look at the data

As a first step, this section discusses the data that we use in our empirical exercise: unemployment rates and different measures of openness to international trade. It also provides a first heuristic look at the unemployment-openness relationship. A detailed discussion of the data is contained in the Appendix.

2.1. Data sources and variables

2.1.1. Unemployment rates

International institutions such as the OECD, the World Bank or the International Labor Organization (ILO) provide harmonized aggregate unemployment rates that are calculated following the same conventions. Across different international institutions, these rules can differ. For example, the rates published by the OECD or the World Bank rely on national administrative sources, while the ILO data is based on labour market surveys. The former strategy presupposes the cooperation of national statistical agencies; the latter is probably better suited to developing countries. Country coverage is always an issue: While the World Bank has 185 members, in the year 2000 it reports unemployment rates only for 93 of them. The ILO data exhibits an even lower degree of country coverage (86 countries). Skill-specific unemployment rates are from the World Bank (WDI data base), but time and country coverage is fairly poor.

In all cases the accuracy of the published rates depends on the quality of the data delivered by the institutions' member states. Data quality is only a minor issue for the 20 rich OECD countries, but appears to be highly problematic for the rest of the world.[8] The correlation between unemployment rates from these different data sets is strikingly low within the group of low-income, low-openness countries, which suggests that data quality systematically depends on country characteristics. Such non-random measurement error in our dependent variable (the rate of unemployment) will tend to bias the absolute value of the estimated effect of openness upwards.

Unfortunately, there is very little that one can do about data quality problems except running as many robustness checks as possible or working with the small panel of OECD countries for which data quality is satisfactory.[9] Hence, in a first step, we focus on 20 high-quality OECD countries, for which systematic measurement bias in the rate of unemployment is unlikely (but where the analysis may suffer from non-random sample selection). This choice strongly lim-

[8] In its statistical factbook, the CIA publishes yearly estimates of unemployment rates for a larger sample of countries (as of 2000, there is data for 160 countries). The CIA makes use of all publicly available information plus the insider information of its employees. How exactly the CIA experts obtain these estimates is not made explicit. In the non-OECD sample, average CIA estimates are substantially larger than the information provided by official sources; in the OECD sample there is no such gap.

[9] More details on countries included is provided in the Appendix.

its the cross-sectional scope of our analysis and makes it necessary to use panel data and rely on time-variance for estimation. In addition, we perform purely cross-sectional regressions with larger country samples and also experiment with a short panel for this larger sample. To verify the robustness of our results, we use different data sources for the dependent variable (unemployment rate). Finally, we also report regression results where we use skill-specific unemployment rates.

2.1.2. Openness measures

The summary measure of trade openness nearly always used in empirical work is *nominal* imports plus exports relative to nominal GDP, usually referred to as (trade) openness and denoted by T. For recent examples see Coe and Helpman (1995), Frankel and Romer's (1999), Ades and Glaeser (1999), Alesina, Spolaore and Wacziarg (2000), Dinopoulos and Thompson (2000) or Alcalá and Ciccone (2004). The openness measure has the advantage that it reflects the actual exposure of an economy to international trade and is easily measurable. Trade policy itself is often hard to observe, in particular because of the declining importance of tariffs or quotas and the increasing use of informal trade barriers. Also, membership in regional trade agreements or the WTO does not necessarily provide information about the actual openness of an economy, see Rose (2005).

Alcalá and Ciccone (2004) argue that the Balassa-Samuelson effect distorts nominal price openness measures since countries with low labor productivity and hence a high price of traded relative to non-traded goods have artificially high degrees of openness. They propose to use *real* openness defined as imports plus exports in exchange rate US$ relative to GDP in purchasing-power-parity US$ (PPP GDP). This eliminates cross-country differences in the relative price of non-traded services from the summary measure of trade. They show how the real openness measure can be computed using data provided in the Penn World Tables (PWT). The measure of real openness may be particularly relevant to the extent that the effect of trade openness on aggregate unemployment works through total factor productivity. We use real total trade openness constructed according to Alcalá and Ciccone (2004) as our benchmark measure. Even if accounting for the Balassa-Samuelson effect is not a big issue for countries in our OECD sample, the problem becomes more severe in our large cross sectional regressions. Comparing real and current price openness measures reveals that the effect is smaller for real openness but coefficients are more stable across different models and setups.[10]

As with unemployment rates, the openness measures may be noisy proxies for the actual degree of exposure to international trade. It is less obvious, however, that measurement error

[10] In our robustness checks, we also work with *constant* price openness measures which fix *all* prices at some base year. Moreover, data provided by the World Bank allows to focus on merchandize trade only. This allows to see whether trade in services has a different effect on unemployment compared to trade in goods.

should be *systematically* related to any determinant of the unemployment rate. Random measurement error would bias estimated towards zero, making it harder for us to find significant effects.

2.1.3. Labor market institutions

The OECD has collected data on a wide array of institutional variables that can be expected to affect the equilibrium rate of unemployment. Bassanini and Duval (2006, 2009) discuss the data in detail. These measures include the degree of union density or of union coverage, the extent of employment protection legislation or of active labor market policies, effective average tax rates on wages, the average replacement rate of unemployment insurance, the degree of corporatism and many more. The data also includes a measure of product market regulation which reflects entry barriers. These variables are available for 20 rich OECD countries, and for most of them we have time series ranging from 1980 - 2003.

The data for the wider cross-section of countries is more problematic. By far the most careful data collection has been undertaken by Botero et al. (2004). They provide a data set containing data on various aspects of labor market regulations for 85 countries. Observations range from 1990 - 2000 and were averaged over the whole period. In our study we focus on measures related to the generosity of unemployment benefits, the extent of employment protection (EPL) and the importance of minimum wages. Additionally to those labor market regulations Botero et al. also collected data on the size of the informal economy. Reported unemployment rates and the degree of openness may both be systematically related to the size of the shadow economy so that omitting this variable could easily bias the effect of trade. This is a particularly important issue in the large cross-section, where we cannot control for unobserved heterogeneity and where we have a large number of developing countries.

The Botero et al. data does not contain a time dimension. Therefore, when running panel regressions for the large country sample, we need to rely on data from the Fraser *Freedom of the world* data base, where we have variables on unemployment benefits, labor market institutions and product market regulations. The former variable is an index that collects information on many dimensions of labor market institutions; the latter quantifies the extent of price controls.[11] Observations for 116 countries are available in five year intervals beginning in 1975 and ranging until 2005.

2.2. A first glance at the openness-unemployment nexus

[11] In the original Fraser data higher values indicate more freedom and thus less regulation. To avoid confusion when comparing with the OECD or the Botero et al. data we rescale the Fraser variables by the factor -1.

2.2.1. Time variance in the OECD sample.

The solid line in Figure 1 plots the *unweighted* average unemployment rate of 20 rich OECD countries (see the Appendix for a list of countries). Starting from a low level at about 2 percentage points in 1970, the unemployment rate increased over time to reach a peak of 10 percent in the mid-nineties, but fell back to about 6 percent in 2003. Measured on the right vertical axis of Figure 1, the unweighted average share of trade in total GDP (measured as real openness) also displays a clear upward trend: it increased from about 25 percent in 1970 to about 40 percent in the early years of the new millennium. Because of this common time trend, average unemployment rates and real openness measures appear to be positively correlated.

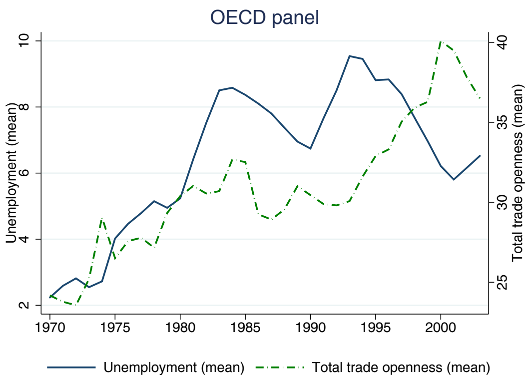

Figure 1: Unemployment and openness

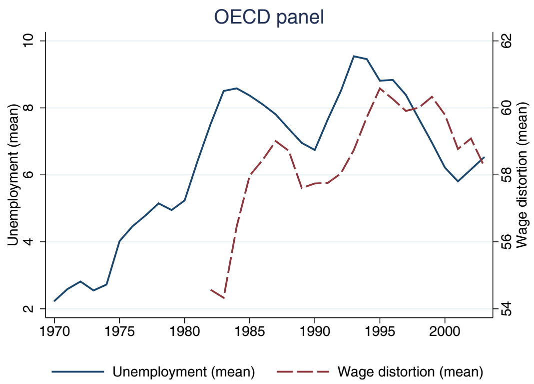

Figure 2: Unemployment and wage distortion

So far, the empirical labor market literature has usually not accounted for any measure of trade openness. Nickell et al. (2005) show that the evolution of labor market institutions has substantial explanatory power for unemployment rates. In particular, tax rates and replacement rates perform well; other institutional variables do not yield robust results. This is not entirely surprising since the theoretical predictions relating to employment protection legislation or union coverage are usually ambiguous. Costain and Reiter (2008) use a theoretical model to argue that tax and replacement rates should have similar qualitative and quantitative effects in a search and matching model of unemployment. They propose to add them. The obtained index consists of the sum of the average wage tax burden and social benefits foregone when a worker switches from unemployment into a job. It therefore measures the *total* fiscal burden imposed on the worker (see also Saez (2002) or Immervoll et al. (2007)) and is sometimes referred to as the participation tax. Figure 2 shows that the average wedge and average unemployment are also positively correlated over time. Hence, the *prima facie* evidence suggests that it is important to control for both variables in any meaningful cross-country unemployment regression that draws on time variance.[12]

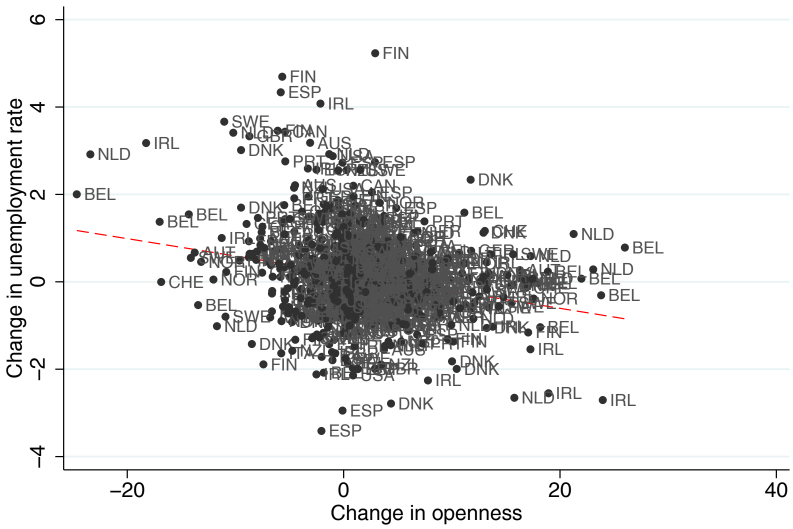

Figure 3: Unemployment and trade openness: first differences of 5-year averages (OECD sample)

Figures 1 and 2 present sample averages over time and fully disregard heterogeneity across countries. In a next step we correlate *first-differences* of the real openness measure against *first-differences* in the unemployment rate. Differencing should eliminate country-specific effects unrelated to openness that may drive the correlation in Figure 1. Figure 3 shows the scatter plot and fits a univariate linear regression. The slope of the line is estimated at -0.04 with a t-value of 5.69. This preliminary evidence points towards a negative effect of trade openness on the rate of unemployment. A one-standard deviation increase (about 10 percentage points) of openness is associated to a decrease in the rate of unemployment of about 0.4 percentage points. Interestingly, our more elaborate multivariate instrumental variable analysis below suggests results of very similar magnitude.

2.2.2. Cross-sectional variance in the large sample

Figure 4 sets the average *level* of unemployment (WDI estimates) against the average *level* of openness (real current price) for the largest cross-section of countries, for which we have data. Averages are based on the period from 1990-2006, but there may be substantial spans of missing values within that period.

The linear regression line fitted to the scatter plot has a slope of about -0.044 with a t-value

[12]In the picture, the unemployment rate leads the measure of wage distortion over time. Costain and Reiter (2008, section 4.3) discuss the endogeneity issues suggested by this fact but conclude that they are unlikely to pose any serious problems.

of 2.20.[13] Hence, also in the large cross-section of countries, the unconditional regression of openness on the rate of unemployment yields a negative correlation. Because the variance of the openness measure is much larger in the large cross-section than in the narrow OECD sample, the point estimate implies that a one-standard deviation increase of openness is associated to a decrease in the rate of unemployment by about 1 percentage point.

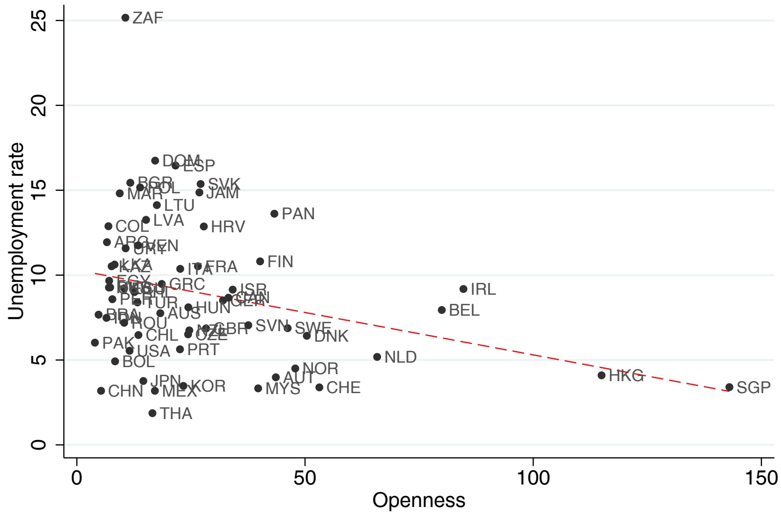

Figure 4: Unemployment and trade openness: averaged levels (large cross-section)

2.3. Implications and challenges

The above figures are suggestive. However, there are several reasons why the correlations in figures 3 and 4 may be spurious. First, while we have used yearly data, there may be business cycle effects: any positive shock on domestic spending is likely to increase domestic as well as import demand, and thus to lower unemployment and increase openness. Second, in periods of reform, countries may simultaneously liberalize their product and labor markets, leading to a simultaneous increase in openness and employment. Third, politicians may react to shocks in the unemployment rate by imposing protectionist measures. More precisely, they may resort to policy measures that discourage imports and encourage exports; since the overt use of tariffs, quotas, or subsidies is strongly restricted by international agreements, governments may use non-tariff measures which are difficult to control for directly. In the case that import-restricting policies dominate, the rise in unemployment would be associated with a reduction in openness.

We deal with the first problem, the *business-cycle effect*, in the following way: In the OECD sample, we take 5-year averages to smooth out business cycle variation. Moreover, in all regres-

[13]The finding of a negative slope is robust to the exclusion of HKG (Hong Kong) and SGP (Singapore); statistical fit is improved by taking logs of both variables.

sions we include a measure of the output gap, based on HP filtering methods, and provided by Bassanini and Duval (2006). In the larger cross-section, we take averages over the entire available period (1990-2006) and also include the output gap.

The second issue relates to an *omitted variables bias*. In the OECD sample, we can draw on high-quality data provided by Bassanini and Duval (2009). For the wider sample, we use the variables provided by Botero et al. (2004). See the Appendix for a detailed description of all our data.

The third and most interesting problem is a classical *simultaneity* problem. We can only address it by instrumenting the openness measures. In the case of the OECD panel, we can exploit the time-variance of the data and use lagged differences and levels as instruments. In the case of the wider cross-section, we draw on the instrument proposed by Frankel and Romer (1999) and used, i.a., by Alcalá and Ciccone (2004). This empirical approach in the cross-section has been criticized in the literature; see Fernandez and Rodrik (2000) or Kraay (2010). The main two issues relate to unresolved omitted variable bias and the validity of the exclusion restriction. We add the variables that have been found in the literature to undo the significance of the growth-openness nexus (e.g., latitude). However, the panel approach is clearly preferable from an econometric point of view.

3. Empirical strategy

We have to adapt our econometric strategy to the nature of the available data. For the OECD sample, where we can draw on meaningful time-variance, we build on the rich tradition of empirical labor market studies surveyed in Bassanini and Duval (2006) and use panel methods. For the wider sample, we use the cross-sectional approach which has been widely employed in the growth-openness literature. While time-variance in the larger cross-section is somewhat problematic, we still check our results by running panel regressions as well.

3.1. OECD sample: GMM panel regressions

We extend Nickell et al. (2005) and estimate variants of a dynamic model

$$u_{i,t} = \sum_{s=1}^{S} \rho_s u_{i,t-s} + \beta \cdot T_{i,t} + \lambda \cdot \mathbf{LMI}_{i,t} + \pi \cdot \mathbf{PMR}_{i,t} + \chi \cdot \ln POP_{i,t} + \gamma \cdot GAP_{i,t} + \nu_i + \nu_t + \varepsilon_{i,t}, \quad (1)$$

where S is the number of lags of the endogenous variables. All variables are five-year averages. The vectors $\mathbf{LMI}_{i,t}$ and $\mathbf{PMR}_{i,t}$ collect variables measuring labor market institutions and product market regulation, respectively. $POP_{i,t}$ refers to population, $GAP_{i,t}$ is the output gap,[14] ν_i is a vector of country-specific effects, ν_t denotes time effects, and $\varepsilon_{i,t}$ is an error term. We are primarily interested in the estimate of β and expect that the effects of LMI and PMR conform with the evidence surveyed in Bassanini and Duval (2009). This evidence is mixed: Baker et al. (2004) show that those panel data estimations lack robustness and that clear results on the role of most labor market institutions hardly exist. There is, however, an emerging consensus that replacement rates and the tax wedge have a robust and theoretically sensible effect; see Costain and Reiter (2008).

The (preferred) equation estimated by Nickell et al. (2005) is similar to (1), but does not include openness or a measure of the country's market size (such as population). They use generalized least squares techniques on this equation and are not particularly worried by the potential endogeneity of labor or product market institutions. Many of the specifications surveyed in Bassanini and Duval (2009) constrain $\rho_s = 0$ and estimate static fixed effects models. Some papers use the log of $u_{i,t}$ as the dependent variable (Nickell, 1997; Costain and Reiter, 2008), but there does not seem any consensus as to which specification is preferred. In our baseline specifications, we use $u_{i,t}$ in levels, but provide robustness checks for the logarithmic case.

We address the potential endogeneity of openness and of the lagged dependent variable by instrumenting with the respective lagged values.[15] In the first-differenced general method of moments (diff-GMM) approach by Arellano and Bond (1991), all variables are differenced and endogenous variables are instrumented by their lags (in differences). The more general approach proposed by Blundell and Bond (1998) adds level equations to the differenced ones. This leads to a system of two different sets of moment conditions (differences and levels). Blundell and Bond use Monte Carlo simulations to show that the sys-GMM approach is more efficient since a larger number of moment conditions is available. All techniques discussed above allow to control for potential endogeneity, even when there is no obvious instrument waiting on the wing. Nevertheless those GMM approaches must be treated cautiously since small degrees of model specification error may induce large effects on results and lagged variables might be weak instruments. There are however, a number of tests that can be used to check whether the conditions of the approach are fulfilled. For both GMM methods, two requirements must hold: i) the instruments must be uncorrelated with the error term and ii) the instruments must be correlated with the instrumented variables. Both types of GMM are valid if we find evidence in favor of first order, but against second order auto correlation in the residuals.[16]

[14] For the OECD output gap is measured as derivation of actual output from potential output (Basanini and Duval (2006). For the large cross section we use a proxy constructed as difference between actual GDP and trend GDP. The latter is obtained by HP-filtering the data, where the smoothing parameter is set to 400.

[15] Additionally, we treat the wage distortion index (sum of average replacement rate and tax wedge) as endogenous.

[16] We have also experimented with the Anderson and Hsiao approach where lagged variables are used as instruments

3.2. Large cross-section of countries: 2SLS regressions

To extend the analysis beyond the 20 rich OECD countries, we focus on a pure cross-section of countries. This approach is strongly related to cross-country income regressions (Frankel and Romer, 1999; Alcalá and Ciccone, 2004), with the most important difference being the change in the dependent variable.

We estimate the following second stage regression

$$u_i = \alpha + \beta \cdot T_i + \lambda \cdot \mathbf{LMI}_i + \pi \cdot \mathbf{PMR}_i + \delta \cdot \mathbf{GEO}_i + \iota \cdot \mathbf{INST}_i + \chi \cdot \ln POP_{i,t} + \gamma \cdot GAP_i + \varepsilon_i, \quad (2)$$

which includes the same type of controls than (1). Given that we have no reliable time-variance available to control for unobserved country-specific fixed effects, we have to add geographical variables to control for the size of the home-market and hence the importance of *within-country trade* as compared to international trade. Frankel and Romer (1999) and much of the following literature use the log of population and the log of land area of country i.[17] Regressions also contain a continuous measure of landlockedness as an additional strictly exogenous control. We proxy for the overall quality of institutions by including distance to the equator and continent dummies.

We follow Frankel and Romer (1999) and instrument T_i by its (exogenous) geographical component; however, our strategy is somewhat more general. It consists in using bilateral trade data (for the year of 2000) and regress total trade (exports plus imports) between country i and j, normalized by country $i's$ GDP, on geographical determinants of trade in an equation of the type

$$T_{ij} = \exp\left[\varphi \mathbf{X}_{ij}\right] \cdot \upsilon_{ij}. \quad (3)$$

The vector $\mathbf{X}$ contains the *log* of bilateral distance between i and j, the *log* of population of i and j as of year 1960, the *log* of land area of i and j, and a continuous measure of landlockedness. It also contains interactions of all those terms with an adjacency dummy. All of the elements in $\mathbf{X}$ are exogenous while υ_{ij} is an error term.

The standard procedure is to take logs of (3) and estimates the vector φ using OLS. Since $T_{ij} = 0$ for many country pairs, we follow Santos and Tenreyro (2006) and estimate (3) using Poisson pseudo maximum-likelihood. Predicting $\hat{T}_{ij}$ and summing over j, we have a measure of the trade share $\hat{T}_i$ that is by construction orthogonal to unemployment and hence a valid instrument.[18] The Poisson approach leads to a stronger instrument since we do not have to

when estimating two stage least square IV regressions. Results are available on request.

[17] While standard in the related literature and crucial for the interpretation of the results, Dutt et al. (2009) do not include these controls.

[18] Note that validity of the instrument does not require that the coefficients associated to $\mathbf{X}$ are *consistently* estimated parameters of a gravity equation. Rather, equation (3) is a constructed exogenous measure of multilateral resistance.

omit the information contained in the zero trade observations and need not resort to out-of-sample predictions to construct the instrument.[19]

3.3. Large sample: Panel regressions

In the setup described in section 3.2., we have averaged yearly available unemployment data for a large set of countries into a cross-section. This seems appropriate to deal with business cycle effects and should also help to reduce (non-systematic) measurement error in both the dependent and the independent variables. It is also possible to generate averages over shorter periods of time (five years), stack data from different periods, and use panel methods. The drawback of this approach is that unemployment data are available only for a very small sample for a long time horizon so that we end up with a strongly unbalanced panel. Nonetheless, applying panel methods still allows us to check the overall robustness of our results in 3.2. to country-specific unobservable effects.

We use the same econometric specification than the one used on OECD data, i.e. equation (1). Since we need time-variant information about labor and product market regulation, we cannot use the Botero et al. (2004) data, but have to work with variables provided by the Fraser Institute (see the Appendix for details on data).

4. The effect of openness on unemployment

In the following section, we present benchmark results for our different samples, empirical strategies and IV strategies. The overall picture is fairly robust and surprisingly clear-cut: regardless of the precise econometric model used, independent from the exact source of data or the definition of the employed openness measure or the nature of controls, we find that higher openness does not increase unemployment. Quite to the contrary, openness *strictly* lowers the equilibrium rate of unemployment in most regressions.

4.1. Benchmark results

4.1.1. OECD sample: panel regressions

Table 1 presents panel regressions for 20 rich OECD countries. The dependent variable is the rate of unemployment in the total working age population (age 15-64). All variables are five-year averages ranging from 1980 - 2003.[20] Robust standard errors are reported. A list of countries

[19] Noguer and Siscart (2005) show that out-of-sample predictions has important adverse implications for the strength of the instrument.

[20] We have also run regressions on yearly data. Results are similar and statistical significance is usually higher. However we prefer to work with averages to better account for business cycle variations.

used in these regressions is provided in the Appendix.

Columns (1) and (2) show standard regressions as carried out by Bassanini and Duval (2009). The first treats country-effects as fixed, the second treats them as random, everything else is equal. We let a Hausman test decide which of the two specifications is preferred. In all cases presented in Table 1 the test recommends the random effects (RE) specification over the fixed effects (FE) model.

The regressions reveal a well-known pattern: only a few labor market controls are statistically significant, and often the sign pattern seems to be counter-intuitive. The stringency of firing restrictions as reflected by our employment protection legislation (EPL) index is negatively associated to the rate of unemployment. Hence, firing restrictions seem to discourage job destruction more than job creation even though the effect is not statistically distinguishable from zero. Similarly, we do not find any robust role for the degree of union density. The degree of wage distortion (the sum of the replacement rate and the average tax rate on wages) is positively related to the equilibrium unemployment rate. Statistically significant at the 1% level, an increase in the wedge by 10 percentage points increases the rate of unemployment by about 1.1 percentage point. Countries with a highly corporatist bargaining culture have an unemployment rate that is by about 2.6 percentage points lower than countries without this tradition. These findings are in line with the literature,[21] and the emerging consensus that the degree of wage distortion is the most important institutional variable in panel regressions.[22] We also add a variable that has received much interest in the last years as a determinant of unemployment, namely the degree of product market regulation (PMR).[23] The effect of PMR on unemployment is positive, but not significant and therefore meaningless.[24]

Although we average our data over five-year intervals to mitigate business cycle concerns, the output gap is strongly significant and has the expected negative sign. This shows that taking averages alone is not sufficient to purge out the business cycle. Also note that country-specific effects are important for the overall explanatory power of the model. A model that explains unemployment only by country-effects yields an R^2 statistic of about 63%; adding year dummies improves the share of left-hand-side variance explained to 75%. In the random effects model shown in column (2), the exact variance decomposition shows that the within component is much larger than the between component.

Columns (3) and (4) include the *real openness* measure proposed by Alcalá and Ciccone (2004) into the fixed- and the random effects models, respectively. Again, the Hausman test recommends the more efficient RE model. Inclusion of the openness measure increases the

[21] As can be seen from the survey by Bassanini and Duval (2009) or the critical discussion in Baker et al. (2002).
[22] See Costain and Reiter (2008).
[23] See Felbermayr and Prat (2009) for theory and evidence on the role of PMR.
[24] Regressions with the logarithm of GDP instead population yield very similar results but raise more serious concerns about regressor endogeneity.

Table 1: **Benchmark regressions: OECD panel**

Dependent variable: Total unemployment (16-64 years old)
Openness measure: Real openness (Alcala & Ciccone, 2004)

	(1) FE	(2) RE	(3) FE	(4) RE	(5) FGLS	(6) Diff-GMM	(7) Sys-GMM
Total trade openness			−0.128***	−0.076***	−0.112***	−0.230***	−0.052***
			(0.035)	(0.021)	(0.021)	(0.062)	(0.019)
Lag dep. var.					0.305***	0.220	0.725***
					(0.047)	(0.174)	(0.089)
Wage distortion (index)	0.114**	0.111***	0.065	0.103***	0.073***	0.016	0.085*
	(0.044)	(0.027)	(0.044)	(0.026)	(0.018)	(0.114)	(0.049)
EPL (index)	−0.444	−1.027	−0.380	−0.969	−0.589	−0.112	−1.188**
	(1.329)	(0.662)	(1.378)	(0.652)	(0.377)	(1.161)	(0.580)
Union density (index)	0.038	0.007	0.025	0.009	0.025*	−0.010	−0.053*
	(0.041)	(0.029)	(0.043)	(0.029)	(0.014)	(0.039)	(0.029)
High corporatism (dummy)	−3.668***	−2.542***	−2.325*	−1.805**	−2.574***	−1.181	−1.572
	(0.822)	(0.735)	(1.203)	(0.744)	(0.467)	(1.399)	(0.981)
PMR (index)	0.745	0.769	0.963	0.835*	0.820***	0.700	0.893*
	(0.553)	(0.478)	(0.591)	(0.462)	(0.230)	(0.669)	(0.476)
Population (ln)	−17.578***	0.739	−19.689**	0.141	−13.402***	−20.200***	−0.610
	(6.007)	(0.540)	(6.994)	(0.605)	(3.391)	(6.832)	(0.704)
Output gap	−0.606***	−0.636***	−0.624***	−0.626***	−0.589***	−0.872***	−0.842***
	(0.082)	(0.114)	(0.089)	(0.114)	(0.047)	(0.168)	(0.125)
Observations	100	100	100	100	100	80	100
R^2 (within)	0.602	0.569	0.648	0.608			
R^2 (between)	0.012	0.353	0.018	0.282			
R^2 (overall)	0.004	0.411	0.008	0.369			
Hausman	0.599		0.188				
Hansen test (OID)						0.407	0.999
AR(1)						0.025	0.017
AR(2)						0.314	0.219

Robust standard errors in parentheses, * significant at 10%, ** significant at 5%, *** significant at 1%. Number of observation N=100 (20 countries observed for 4 5-year periods and 1 4-year period; averages taken; 1980-2003). Hausman test p-values reported (Fixed effects estimator always consistent; random effects estimator efficient under Ho). All models control for unobserved country and period effects. FGLS allows for heteroscedastic errors and country specific first order serial correlation. First lag of dependent variable used for Feasible Least Square and Generalized Methods of Moments regressions. Diff- and Sys-GMM estimators are valid if i) OID test does not reject the H0 (H0: overidentifying restrictions are valid) and ii) if test on AR(1) is positive and negative on AR(2) (H0: no autocorrelation). Openness, output gap and wage distortion treated as endogenous in the GMM regressions. Maximum number of lags used as instruments equals one (21 instruments for diff-GMM, and 36 instruments for sys-GMM). Constant estimated but not reported.

explanatory power (within R^2) of the regression by about 5 percentage points. Focusing on the RE specification and comparing the models with and without the openness measures, we find that the coefficients on the labor market variables change only very slightly so that omitted variable bias from not incorporating openness seems unimportant. This suggests that labor market regulation does not systematically correlate with the degree of openness. Also the output gap does not seem to covary with openness. The effect of openness on the rate of unemployed is estimated to be 0.076. Hence, a 10 percentage point increase lowers the equilibrium rate of unemployment by about 0.76 percentage points.

Given that column (4) reports our preferred estimate, it is worthwhile to note that it implies a rather moderate contribution of trade liberalization for unemployment. Amongst larger countries, such as the US, Japan, or the EU en bloc, pre-crisis openness was at about 30%, 34% and 29%, on average 13% higher than before world war II. The increase in openness was therefore responsible for a decrease in the average unemployment rate of about 1.2 percentage points. Given the standard deviation of unemployment rates in our sample (about 4 percentage points), this seems a sizable effect. Yet, it is clear that other determinants of unemployment rates (such as institutions) play a more important role.

The remaining models presented in Table 1 are dynamic models. Column (5) uses the feasible generalized least square methodology proposed by Nickel et al. (2005) to estimate an autoregressive model.[25] The lagged rate of unemployment has an estimated coefficient of about 0.3, signaling that–over our five-year periods–unemployment rates are only mildly persistent. Again, the effect of openness is precisely estimated and negative. The short-run effect together with the autoregressive coefficient implies that a ten percentage point increase in openness lowers the equilibrium rate of unemployment by roughly 1.1 percentage points in the short-run, and by about 1.6 percentage points[26] in the long-run.[27]

So far we have not dealt with the potential endogeneity of openness. Models (6) and (7) use lagged realizations or lagged differences of openness as instruments. In the first case, GMM estimation is applied to a differenced version of equation (1). In the second case, moment conditions from an additional level equation are used to increase efficiency. In both cases, we find that openness reduces unemployment. In the diff-GMM model (6), the short- and the long-run effects coincide. A ten percentage points increase of openness suggests a reduction in average unemployment rate by about 2.3 points, which seems implausibly large. In the more general sys-GMM model (7), the short-run effect is smaller: a 10 percentage points increase in openness decreases unemployment by about 0.5 percentage points. The long run effect, however, is again comparable: a 10 percent openness increase leads to lower unemployment by 1.9

[25] Their approach includes country effects into the regressions.
[26] $0.112/(1-0.305)$.
[27] Long-run coefficients are found at the fixed-point of the difference equation.

points,[28] which is comparable to the FGLS results. GMM methods are vulnerable to misspecification problems and applicable only under certain conditions. For both models, the OID tests for overidentification yield high p-values so that validity of the instruments cannot be rejected.[29] Furthermore, the AR(1) and AR(2) statistics suggest that the model is not misspecified.

Comparing (long-run) estimates across different columns of Table 1, we find that the point estimates of the openness coefficient are typically larger under the IV strategy. This is consistent with several explanations. First, the non-IV estimates may be biased down (in absolute value) due to endogeneity bias. This would happen if governments respond to adverse unemployment shocks by promoting exports since then total openness, which reflects imports as well, would also go up. Second, the fact that non-IV estimates are biased towards zero may arise when our openness indicator is a noisy proxy of the true relevant degree of openness. Since instrumentation also remedies measurement error, this may explain the observed sign of the bias.

[28] $0.052/(1-0.725)$.

[29] Note that the tests remain stochastic (p-values < 1) and consequently meaningful.

Table 2: **Benchmark regressions: large cross-section**

Dependent variable: Total unemployment (WDI)
Openness measure: Real openness (Alcala & Ciccone, 2004)

	(1) OLS	(2) IV	(3) OLS	(4) IV	(5) OLS	(6) IV	(7) OLS	(8) IV	(9) OLS	(10) IV
Total trade openness	−0.047***	−0.036	−0.093***	−0.102*	−0.087***	−0.099	−0.079***	−0.097**	−0.081***	−0.099**
	(0.011)	(0.025)	(0.019)	(0.055)	(0.023)	(0.066)	(0.022)	(0.04)	(0.028)	(0.048)
Unemployment benefits (index)									1.644	1.659
									(1.638)	(1.405)
EPL (index)									5.330*	5.515**
									(2.830)	(2.553)
Minimum wage (index)									−0.013	−0.413
									(1.546)	(1.608)
Unofficial economy (index)					0.021	0.013	0.016	0.003	0.014	0.006
					(0.041)	(0.057)	(0.041)	(0.046)	(0.037)	(0.039)
Population (ln)			−1.010**	−1.005**	−0.947**	−0.965**	−0.848	−0.916*	−0.610	−0.622
			(0.443)	(0.418)	(0.435)	(0.427)	(0.556)	(0.510)	(0.546)	(0.480)
Land lockedness (index)			−0.778	−0.908	−0.873	−0.974	−1.795	−1.952	−1.991	−2.046
			(1.429)	(1.483)	(1.399)	(1.359)	(1.423)	(1.311)	(1.454)	(1.296)
Latitude			0.018	0.023	0.117	0.086	−0.002	−0.033	−0.058	−0.054
			(0.372)	(0.357)	(0.355)	(0.406)	(0.419)	(0.421)	(0.485)	(0.440)
Area (ln)			−0.355	−0.440	−0.357	−0.445	−0.412	−0.520	−0.743	−0.887*
			(0.358)	(0.523)	(0.362)	(0.505)	(0.451)	(0.444)	(0.466)	(0.507)
Output gap (ln)	−60.837	−63.129	−62.04	−63.33	−61.74	−63.18	−61.86	−64.76	−70.71	−73.70*
	(53.028)	(50.635)	(51.81)	(50.61)	(53.41)	(51.39)	(50.22)	(45.82)	(44.19)	(38.96)
Continent dummies							x	x	x	x
Observations	62	62	62	62	62	62	62	62	62	62
R^2 (adjusted)	0.085	0.08	0.162	0.160	0.150	0.147	0.316	0.311	0.330	0.325
F (1st stage)		27.527		20.456		20.442		31.364		20.649
Partial R^2		0.538		0.269		0.238		0.408		0.392

Robust standard errors in parentheses, * significant at 10%, ** significant at 5%, *** significant at 1%. All variables averaged over the period 1990 - 2006 in order to net out business cycle effects. World development indicators total unemployment rate used as dependent variable and current price real total trade openness captures trade liberalization. Openness instrumented with an improved Frankel & Romer (1999) instrument.

4.1.2. Large sample: cross sections

Next, in Table 2, we study the effect of *real openness* in a cross-section of 62 countries. Unemployment rates are taken from the World Development Indicators data base provided by the World Bank. We average all variables over the window 1990-2006, so that business cycle effects are unlikely to contaminate the results. We nevertheless control for the output gap. We deal with endogeneity as described in section 3.2. by using an improved Frankel and Romer (1999) - type instrumentation strategy.

Column (1) is the most parsimonious model. It uses no additional controls (except the output gap whose inclusion is inconsequential). The OLS regression produces a coefficient of 0.047, estimated with high precision, and implying that a 10 percentage points increase in openness lowers unemployment by about half a percentage point. When openness is instrumented, the point estimate is close to zero and statistical significance is lost. Hence, it appears that, in this very parsimonious model, OLS strongly overestimates the absolute size of the openness effect.

Column (3) and (4) are virtually identical to Table IV in Alcalá and Ciccone (2004) or to Table 3 in Frankel and Romer (1999), with the key differences being the different dependent variable and a slightly more general construction of the instrument. These papers stress the importance of including variables that control for the size of the domestic market (logarithm of population, the logarithm of land area, and a continuous measure of landlockedness). This is crucial since a country's degree of openness is negatively correlated to its own economic size. As suggested by theoretical arguments based on economic geography models, omitting the domestic market size control biases the openness coefficient away from zero if domestic market size is positively correlated to the unemployment rate, and biases it towards zero if it is negatively correlated.[30] The regressions also include a rough proxy for institutional quality–the logarithm of distance to the equator (latitude). The IV estimate is now significant at the 1 percent level. It follows that the failure to produce a significant IV coefficient in column (2) is not due to endogeneity bias, but rather to omitted variable bias.

Models (5) and (6) add a variable provided by Botero et al. (2004), namely the size of the unofficial economy as a share of officially reported GDP. It is plausible to assume that more open economies have smaller unofficial sectors, since exporting or importing requires formal clearing at the borders. It may also be the case that officially reported unemployment rates are inversely proportional to the size of the shadow economy. Indeed, in our data the discrepancy between estimates by the CIA and official data correlates with the size of the unofficial economy. Hence, it seems meaningful to control for the extent of the shadow economy. Compared to the results presented in columns (3) and (4), we find that this additional variable leaves the OLS

[30] Assuming for simplicity that all covariates other than openness and domestic market size are uncorrelated, the bias is $\beta_{size} \times cov\left(open, size\right)/var(open)$.

estimates broadly unchanged but undoes the statistical significance of openness in the instrumental variable regressions. The size and sign of the estimates hardly moves. This is, however, not a robust result. For example, taking out latitude restores significance. More importantly, even with latitude included, we obtain fairly precise and roughly comparable estimates for both the OLS and the IV regressions when the model is augmented by continent dummies. The latter may help to further control for unobserved heterogeneity across countries.

Finally, models (9) and (10) are the most comprehensive in that they include a list of labor market covariates provided by Botero et al. (2004). In particular, we use a measure related to the strictness of employment protection legislation (EPL), an index related to unemployment benefits, a variable indicating the existence of minimum wages and a variable measuring non-wage costs of labor (i.e., taxes). With the exception of EPL, none of those additional controls turns out significant.

Summarizing, we find that across most multivariate cross-sectional regressions, the effect of a 10 percentage points increase in openness lowers unemployment by about 1 percentage point (columns (8) and (10)). As with the high-quality OECD data, and presumably for the same reasons, there is no robust evidence that OLS overestimates the size of the true effect. In particular, in the more complete specification, it is hard to see any difference between IV and OLS results.

4.1.3. Large sample: panel regressions

Table 3 runs panel regression of five-year averages on a larger set of countries. We employ the same econometric specifications and use similar controls as in section 4.1.1.. In particular, we control for the output gap in all specifications. This is important as taking five-year averages does not seem to entirely purge business cycle effects. We control for market size changes by including the logarithm of population. The institutional labor market controls are from the Fraser Institute and measure overall hiring and firing restrictions and the replacement rate.[31] We also use a measure of product market regulation from the same data source. We do not have time-variant information about tax rates. Geographical variables and time-invariant institutional features are accounted for by country effects.

The results confirm the existence of a negative relation between *real openness* and the rate of unemployment. More specifically, columns (1) and (2) show the fixed (FE) and the random effects (RE) model. The Hausman test (p-value of 0.291) prefers random effects. This choice has important quantitative implications in the present setup since the openness coefficient is more than twice as large in the FE model than in the RE specification. The latter indicates that an increase of openness by 10 percentage points lowers unemployment by about 0.78 percentage points. It is striking how close this latter effect comes to our cross-sectional results presented

[31] The benchmark data from Botero et al. (2004) has no time dimension.

Table 3: **Benchmark regressions: large panel**

Dependent variable: Total unemployment (WDI)
Openness measure: Real openness (Alcala & Ciccone, 2004)

	(1) FE	(2) RE	(3) FGLS	(4) Diff-GMM	(5) Sys-GMM
Total trade openness	−0.223***	−0.078***	−0.217***	−0.639**	−0.055*
	(0.063)	(0.020)	(0.023)	(0.288)	(0.031)
Lag. dep. var.			0.106**	−0.410	0.313
			(0.047)	(0.367)	(0.204)
Pop (ln)	−5.337	−0.584*	5.202**	−3.934	−0.663
	(6.987)	(0.306)	(2.119)	(4.093)	(0.870)
LMR (index)	0.638*	0.448*	0.546***	−0.091	1.112**
	(0.372)	(0.248)	(0.101)	(1.104)	(0.544)
Unemployment benefits (index)	0.076	0.128	0.210***	0.407	0.0001
	(0.160)	(0.141)	(0.043)	(0.285)	(0.163)
PMR (index)	−0.227*	−0.126	−0.253***	−0.419**	−0.194
	(0.133)	(0.127)	(0.054)	(0.213)	(0.158)
Output gap (%)	−15.88***	−19.43***	−21.84***	−43.58*	−15.87
	(5.658)	(5.736)	(3.259)	(24.48)	(14.50)
Observations	186	186	164	93	164
R^2 (within)	0.291	0.243			
R^2 (overall)	0.04	0.132			
R^2 (between)	0.063	0.116			
Hausman	0.291				
Hansen (OID)				0.485	0.439
AR(1)				0.598	0.023
AR(2)				0.294	0.645

Robust Standard errors in parentheses, * significant at 10%, ** significant at 5%, *** significant at 1%. All variables averaged over 5 year periods between 1971 - 2005 in order to net out business cycle effects. Number of observations N=186 (77 countries, 5-year periods; data averaged). Panel is strongly unbalanced due to missing observations (186 five year averages available). Dependent variable is World Development Indicators total unemployment rate. Data on labor and product market regulation from Fraser institute. All models control for unobserved country- and period effects. FGLS allows for heteroscedastic errors. First lag of dependent variable used for Feasible Least Square and Generalized Methods of Moments regressions. Diff- and Sys-GMM estimators are valid if i) Sargan test does not reject the H0 (H0: overidentifying restrictions are valid) and ii) if test on AR(1) is positive and negative on AR(2) (H0: no autocorrelation). Openness, output gap and LMR treated as endogenous in the GMM regressions. Maximum number of lags used as instruments equals one (16 instruments for diff-GMM and 28 instruments for sys-GMM). Constant estimated but not reported.

above.

The dynamic models (3) to (5) are problematic because the panel is strongly unbalanced and the number of observations over time is very small for some countries. Interestingly, in all dynamic models, the evidence for persistence in (five-year-averaged) unemployment rates is fairly low and much smaller than in the case of the OECD sample where country coverage is more homogeneous and the panel is longer. The FGLS model signals a short-run openness coefficient close to the one obtained under FE in column (1); the long-run effect is almost identical. Diff-GMM produces similar results. The Sys-GMM model is more efficient, and can make use of more observations. The OID test and the other test statistics are fine, so that we take the Sys-GMM results as the most credible. Here, an increase in openness by 10 percentage points reduces equilibrium unemployment by about 0.55 percentage points in the short-run and by 0.8 points in the long run. Notice the quantitative similarity of these coefficients with those obtained for the smaller OECD sample discussed in section 4.1.1..

5. Robustness checks and additional results

In this section we investigate whether openness affects different skill-classes differently. We also discuss the sensitivity of our main results with respect to alternative openness measures, unemployment data and additional controls. Finally, we show that the effect of openness on unemployment is likely to work through TFP.[32]

Openness and skill-specific unemployment. It is natural to investigate the effects of openness on a more disaggregated level by substituting aggregate with skill-specific unemployment. This allows us to assess whether all skill groups equally benefit from globalization, or whether the beneficial overall effect obscures potential job losses for certain groups of workers. We use data from the World Bank's WDI data set which allows to calculate skill-specific unemployment rates. Unfortunately the data coverage is poor, and observations exist at best from 1994 onwards. Hence, we average the data over time and focus on the cross section. Table 4 reports the results for the key coefficients (full results are in the Appendix). The first four columns refer to standard regressions; columns (5) to (8) include interaction terms with endowment shares. Over all skill classes, openness has a negative effect on the unemployment rate. However, the effect is statistically significant only for high-skilled workers. This pattern suggests that the result found for aggregate unemployment is robust over skill-classes, but the high-skilled labor market segment plays by far the most important role in the aggregate trade-unemployment relationship.

[32] In this section, to save space, we present only the openness coefficients. Full regression output is detailed in a companion paper, which is available on request.

Table 4: **Openness and skill-specific unemployment**

Dependent variable: Skill-specific unemployment
Openness measure: Real openness (Alcala & Ciccone, 2004)

	Skill-specific unemployment				Skill-specific unemployment HO			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	OLS	OLS	IV	IV	OLS	OLS	IV	IV
DEPENDENT VARIABLES ⇒	u (low)	u (high)	u (low)	u (high)	u (low)	u (high)	u (low)	u (high)
Total trade openness *(T)*	-0.015	-0.062^{**}	-0.038	-0.065^{*}	-0.028	-0.089^{*}	-0.099	-0.201^{***}
	(0.039)	(0.027)	(0.041)	(0.037)	(0.053)	(0.050)	(0.061)	(0.070)
Endowment share (L_{low}/L_{high})					0.219	-0.133	0.044	-0.343
					(0.386)	(0.402)	(0.301)	(0.350)
Interaction $(T \times L_{low}/L_{high})$					0.015	-0.002	0.034^{**}	0.050^{**}
					(0.014)	(0.018)	(0.015)	(0.02)

Each row represents one regression. Openness coefficients, endowment share coefficients, and interaction coefficients reported only. Robust standard errors in parentheses, * significant at 10 %, ** significant at 5 %, *** significant at 1 %. We use skill-specific unemployment rates as dependent variable. Data for skill-specific unemployment is available for the period 1994 - 2003 (WDI). We average the data over the whole period to construct a cross section. In row 1 - 4 we regress openness on high and low skill unemployment, in row 5 - 8 we additionally include the interaction between openness and the low to high skill endowment share. We use Barro & Lee data to construct the endowment shares.

Columns (5) to (8) additionally include the endowment ratio and its interaction with openness. We term this set of regression Heckscher-Ohlin (HO) regressions, because in the HO framework, the effect of trade liberalization on skill-specific unemployment rates depends on the relative endowments. Moore and Ranjan (2005) show that lower trade costs reduce the high-skilled unemployment rate in skill abundant countries and increases it elsewhere, while the low-skilled unemployment rate behaves in the opposite way. For low-skilled workers, we find inconclusive results. on the other hand, when looking at the high-skilled segment, the IV regressions show that unemployment falls by less if the country is richly endowed with low-skilled workers, as predicted by HO explanations.[33]

Alternative openness measures. Table 5 presents summary results on alternative openness measures. Each cell reports point estimate and standard error associated to openness. Coefficients pertaining to the dynamic Sys-GMM model refer to the fixed-point of the difference equation. In a first step, we stick with the real openness measure of Alcalá and Ciccone (2004), but use export and import openness rather than the canonical measure (essentially the average

[33] The result implies that there is some threshold value of the endowment share for which the negative effect of openness turns positive. The endowment ratio ranges from 0.18 to 10.47 with an average of 3.16. Computing the threshold for which the marginal effect of openness turns from negative to positive yields 4.00, which is between the minimum and the maximum. For countries with low to high skill endowment ratio greater than 4 openness is positively associated with high-skill unemployment.

of these measures). In all specifications reported in lines **i** and **ii**, we find negative coefficients, except for the system GMM estimator, these are also statically different from zero.

In the main body of this paper, we use the *real openness* measure of Alcalá and Ciccone (2004). This is our preferred indicator, because the effect of openness may affect the tradeable sector differently than the non-tradeable sector. Nonetheless, the growth-openness literature uses an uncorrected measure that we call *current price openness*.[34] Lines **iii**, **iv**, and **v** of Table 5 report results for current price openness. We also try the constant price openness measure reported in the Penn World Tables (line **vi**) and an indicator that draws only on merchandise trade (i.e, excluding services; line **vii**). Across all these specifications, we do not find a single positive coefficient. Coefficient estimates are often algebraically bigger than in our benchmark results, so that the choice of the openness measure does have an influence on the quantitative interpretation of results. Some of the coefficients from the large panel are insignificant statistically, but for reasons detailed above we do not want to over emphasize these findings. Hence, we confirm our general conclusion that openness certainly does not increase unemployment in the long-run.

Log unemployment. There is no apparent consensus in the labor market literature as to whether unemployment regressions have to be run with the dependent variable in logs or in levels. Almost all equations discussed in Bassanini and Duval (2009) are in levels whereas the recent paper by Costain and Reiter (2008) uses logs. In the present setup, results are largely independent of this choice, as can be seen from line **viii** of Table 5, where we keep estimation strategies and samples identical to those used in the upper part but use the log of unemployment as the dependent variable. While significance of the openness coefficient may be lost in some cases, there is no evidence–not in a single regression–that openness increases unemployment in the long run.

Alternative unemployment measures and data sources. Our benchmark regressions use total unemployment rates provided by the OECD, and in the larger samples, data reported by the World Bank in their World Indicator Data base. There are substantial concerns about data quality, in particular in samples that include developing countries. Moreover, even OECD countries have very different approaches to dealing with employment issues for workers at the start or the end of their professional careers. We deal with this problem by running our regressions using alternative unemployment measures.

For the OECD we substitute the total unemployment rate by prime age and youth unemployment but use the Alcalá and Ciccone real openness measure. The first two columns in line

[34] See section 2.1.2. for a more detailed discussion of different openness measures.

Table 5: **Robustness checks**

Dependent variable: Total unemployment (OECD and WDI)						
	OECD panel		Large cross section		Large panel	
Openness measure ⇓	(1) FE/RE	(2) Sys-GMM	(3) OLS	(4) IV	(5) FE/RE	(6) Sys-GMM
Real import and export openness						
i: Import	−0.196***	−0.168**	−0.084***	−0.107**	−0.081***	−0.077
	(0.038)	(0.072)	(0.030)	(0.052)	(0.021)	(0.058)
ii: Export	−0.050***	−0.213***	−0.077***	−0.093**	−0.178***	−0.086**
	(0.019)	(0.065)	(0.026)	(0.045)	(0.064)	(0.039)
Current price openness						
iii: Total trade	−0.057**	−0.214**	−0.026	−0.123*	−0.032**	−0.061
	(0.027)	(0.105)	(0.017)	(0.066)	(0.014)	(0.039)
iv: Import	−0.081***	−0.257**	−0.023	−0.140*	−0.029**	−0.041
	(0.031)	(0.115)	(0.019)	(0.077)	(0.014)	(0.049)
v: Export	−0.036	−0.160*	−0.028*	−0.110*	−0.032**	−0.079**
	(0.024)	(0.091)	(0.016)	(0.057)	(0.013)	(0.037)
Constant price total trade openness						
vi: Total trade	−0.075***	−0.171**	−0.027	−0.130*	−0.042***	−0.039
	(0.021)	(0.073)	(0.018)	(0.072)	(0.015)	(0.037)
Merchandize trade openness						
vii: Total trade	−0.035	−0.154*	−0.013	−0.073*	−0.029**	−0.07*
	(0.032)	(0.082)	(0.010)	(0.040)	(0.014)	(0.04)
Log total unemployment and real total trade openness						
viii: Total trade	−0.006*	−0.018***	−0.009**	−0.009*	−0.009***	−0.008
	(0.003)	(0.005)	(0.003)	(0.005)	(0.003)	(0.008)
Unemployment rate	Sys-GMM Prime	Sys-GMM Youth	IV CIA	IV IFS	Sys-GMM ILO	Sys-GMM IFS
ix: Total trade	−0.196**	−0.112	−0.166**	−0.083*	−0.103*	−0.091*
	(0.083)	(0.190)	(0.067)	(0.045)	(0.054)	(0.047)

In row i - ix, each cell represents one regression. Openness coefficients reported only. Robust standard errors in parentheses, *significant at 10%, ** significant at 5%, *** significant at 1%. All variables averaged over 5-year periods between 1980 - 2003 (OECD panel), 1971 - 2005 (large panel) and over the whole period 1990 - 2006 (large cross section) to net out business cycles. Long-run effects reported for sys-GMM regressions. Total unemployment rate (OECD and WDI) used as dependent variable in row i - viii. Real import export openness measures used in row i and ii, Current price openness measures used in row iii - v, constant price openness in row vi, merchandize in row vii. In row viii we use the respective ln unemployment variable. In row ix we use prime age, youth, CIA, IFS, and ILO data instead of total unemployment. An improved Frankel & Romer (1999) instrument used for the IV regressions. FE/RE: fixed or random effects model selected according to Hausman test (RE is preferred for all regressions). For further details see Tables 1,2, and 3.

ix of 5 show sys-GMM estimates. For prime age unemployment, openness has a stronger effect than for youth unemployment and is not statistically significant in the latter case. This is not overly surprising because youth unemployment is probably much more strongly related to institutional features of labor markets rather than to the extent of trade openness.

The remaining columns in line **ix** of Table 5 report results for the larger cross-section and then for the larger panel, but use unemployment data from alternative data sources. Most importantly, data from the CIA leads to a much stronger effect of openness on the structural rate of unemployment. This is a robust finding, for which we present more evidence in the supplement paper. The other data sources also yield negative coefficients that are of similar size to those obtained with our preferred data base, the WDI.

TFP and trade openness. Next, we present evidence consistent with the view that the effect of openness on unemployment works via TFP. Our results are tentative, because the construction of a TFP measure from observable data requires critical assumptions so that the measure is very imperfect. [35] Also, TFP is likely not exogenous. For these reasons, we do not want to overemphasize our results but rather view them as a first piece of evidence.

Column (1) in Table 6 shows that countries with higher TFP have lower unemployment rates. Note that the relationship cannot be driven by business cycle variation since we work with averages over 5-years, and have included year dummies as well as a measure of the output gap into the regressions. The effect is fairly strong in the OECD panel: a one percent increase in TFP lowers the equilibrium rate of unemployment by about 0.3 percentage points. Going from the sample mean of TFP to the highest realization, the decrease in unemployment is about 6 percentage points. The other cells in the first and second panel show that the relationship continues to hold when using more elaborate regression methods. If anything, controlling for endogeneity biases strengthens the size of the correlation. The third and last panel reports results for the large cross-section where TFP is important, too. Then a one percent increase in TFP lowers unemployment by about 0.04 percentage points. Due to greater variance of TFP measures in the large cross-section, moving from the sample mean to the highest realization of TFP yields an unemployment reduction of about 2.8 percentage points.

These findings are not necessarily contradictory with the concurrent increases in productivity and unemployment observed in Europe over the post-war period because the structure of the regressions is such that TFP levels are not relevant *per se*.[36] Identification relies on time variation and demeaned cross-country variance so that lower unemployment will arise for two reasons. First, countries that had higher TFP *growth* should exhibit lower unemployment, as

[35] We construct our measure of TFP by following the procedure in Benhabib and Spiegel (2005). We apply the perpetual inventory method to back out estimates for capital and then compute TFP as the Solow residual. We use the original estimates published in Benhabib and Spiegel (2005) for the large cross-section.

[36] We thank an anonymous referee for raising this point.

Table 6: **Channels of interaction**

Dependent variable: total Unemployment (OECD and WDI), or "channel variables"
Channel variable: TFP
Openness measure: Real openness (Alcala & Ciccone, 2004)

	(1)	(2)	(3)	(1)	(2)	(3)
I Dep. var. ⇒	u	log TFP	u	u	log TFP	u
	OECD panel			**OECD panel**		
	FE/RE	FE/RE	FE/RE	FGLS	FGLS	FGLS
log TFP	−0.312***		−0.295***	−0.491***		−0.364***
	(0.080)		(0.095)	(0.079)		(0.087)
Total trade openness *(real)*		0.264**	−0.014		0.390***	−0.066**
		(0.119)	(0.030)		(0.07)	(0.031)
	OECD panel			**OECD panel**		
	Diff-GMM	Diff-GMM	Diff-GMM	Sys-GMM	Sys-GMM	Sys-GMM
log TFP	−0.789*		−0.670	−0.477*		−0.516*
	(0.479)		(0.521)	(0.284)		(0.289)
Total trade openness *(real)*		0.635*	0.002		2.476**	−0.017
		(0.341)	(0.141)		(0.976)	(0.082)
	Large cross section			**Large cross section**		
	OLS	OLS	OLS	IV	IV	IV
log TFP	−4.231**		−2.949	−4.231***		−2.244
	(1.783)		(2.376)	(1.471)		(3.599)
Total trade openness *(real)*		0.008***	−0.027		0.008***	−0.042
		(0.001)	(0.034)		(0.002)	(0.067)

Each column in each cell represents one regression. Openness and channel variable coefficients reported only. As channel variables we use Total Factor Productivity. In (1) we regress the channel variable on unemployment, in (2) we regress the channel variable on openness, and in (3) we regress openness and the channel variable on unemployment. Robust standard errors in brackets, * significant at 10%, ** significant at 5% and *** significant at 1%. For the OECD panel we run benchmark type fixed and random effects regressions in the upper left panel (Hausman test indicates that RE is efficient in (1) and (3)) and FGLS regressions in the upper right panel. Openness, output gap and wage distortion treated as endogenous when preforming diff- and sys-GMM in the middle left and right panel (OECD). For the large cross section we run benchmark type OLS and IV regressions. An improved Frankel & Romer (1999) instrument used as instrument for the IV regressions.

extensively documented by Pissarides and Vallanti (2004). Second, countries with higher TFP than the cross country average are also likely to have smaller unemployment rates, as implied by the theoretical model in Felbermayr, Prat and Schmerer (2008).

Column (2) in the table shows that openness and TFP are positively related. We treat openness as endogenous using the same empirical strategy than in the benchmark regressions. The results are broadly in line with Alcalá and Ciccone, who use a somewhat different definition of TFP for the year of 1985 in their cross-sectional analysis. Doubling real openness from the sample mean (about 35 for the OECD panel and 30 in the large cross section) leads to an increase in TFP by about 10 percent in the FE/RE effects benchmark OECD regressions and by about 24 percent in the large cross-section for both OLS and IV. The additional FGLS and GMM regressions in the upper right and middle panel reveal the same significant relationship and thus support the benchmark results.

Let us now turn our attention to our main interest, that is the interaction between TFP and trade openness. The third columns of each cell use both real openness and the log of TFP in the same unemployment regressions. Interestingly enough, adding TFP leads to drastic losses in statistical significance for trade openness. Among all specifications, only the FGLS regression in the OECD sample yields a statistically significant negative coefficient for our preferred measure of openness, a finding that stands in sharp contrast to the robustness exhibited in previous regressions. However, out of the five non significant coefficients, four are negative.

These results suggest that that the impact of openness mostly goes through TFP. This is an intriguing implication because it echoes recent theoretical research on the interactions between trade, firm selection and unemployment. In search-theoretic explanations of equilibrium unemployment, firms with higher productivity find it more attractive to post vacancies; see Epifania and Gancia (2005) or Felbermayr, Prat and Schmerer (2008). In the latter work, more openness forces inefficient firms to quit and allows more productive ones to expand. The average firm's productivity increases, its revenue per match relative to the costs of vacancy creation goes up, and so do its incentives to create jobs. Hence, increased openness leads to lower equilibrium unemployment in the long-run through higher productivity. Establishing the existence of causal links from trade to TFP and then from TFP to unemployment would obviously require more detailed data on industry structure with potentially exogenous episodes of trade liberalization. Our findings can nonetheless be interpreted as encouraging piece of evidence for further research in that direction.

6. Conclusion

This paper establishes an empirical regularity: trade openness does not increase structural unemployment in the long run. Quite to the contrary, in most of our regressions, we find overwhelming evidence for a beneficial effect. This finding is robust to the choice of sample, estimation strategy, and does not hinge on our particular choice of openness measure or the definition of the unemployment rate.

Our analysis draws on two long-standing research traditions: panel unemployment regressions for OECD countries, recently summarized by Nickel et al. (2005), and cross-sectional analysis of the effect of trade liberalization pioneered by Frankel and Romer (1999). In all cases, we average our data and use information on the output gap in order to control for business cycle effects. We include a large host of institutional variables and of geographical controls related to the importance of domestic as compared to international trade. Whenever possible, we include country and year effects. We deal with the possible endogeneity of openness either by exploiting the time dimension of the data or by using the geography-based instrumentation strategy developed by Frankel and Romer (1999). All of our different approaches have advantages and drawbacks. However, the picture across all models is fairly stable and robust: There is no evidence for an unemployment-increasing effect of openness.

Our results are therefore in line with theoretical work that points towards a negative effect of trade liberalization on the structural rate of unemployment. Models of this type are presented in Dutt et al. (2009) or in Felbermayr, Prat, and Schmerer (2008). The recent work by Helpman, Itshoki, and Redding (2008) is also compatible with the evidence.

Finally, it is worth noting that the present paper has a focus on long-run effects. We pay special attention to netting out business cycle disturbances. In this sense, our work is complementary to a growing number of empirical papers on the short-run implications of trade liberalization for labor markets.

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A Data description and summary statistics

A1. Unemployment rates

Table 7: **Unemployment rates according to different sources**

Year	Sample	Unemployment rate (average)			ratio CIA / ILO	
		WDI	ILO	CIA	Avg.	Median
1990	Full (N=48)	7.74	7.79	9.69	1.29	1.16
	OECD 20	6.90	6.88	7.02	1.07	1.00
	RoW	8.16	8.24	11.03	1.40	1.18
1995	Full (N=68)	8.69	9.00	9.64	1.16	1.10
	OECD 20	8.74	8.75	10.39	1.22	1.17
	RoW	8.68	9.10	9.34	1.13	1.08
2000	Full (N=77)	9.06	9.43	10.88	1.39	1.02
	OECD 20	6.15	6.13	6.73	1.09	1.03
	RoW	10.09	10.59	12.34	1.50	1.02
2005	Full (N=69)	8.94	8.94	9.89	1.15	1.07
	OECD 20	6.39	6.34	6.63	1.04	1.03
	RoW	9.98	9.99	11.23	1.20	1.08

Data sources: CIA (Central Intelligence Agency); ILO (International Labor Organization), WDI (World Development Indicators, World Bank).
OECD20 sample includes the 20 OECD countries used in Bassanini & Duval (2009) and in our panel regressions.

Countries included: Albania^{C}, Argentina^{BCD}, Australia^{ABCD}, Austria^{ABCD}, Belgium^{ABCD}, Bolivia^{BCD}, Brazil^{BCD}, Bulgaria^{BCD}, Canada^{ABCD}, Chile^{BC}, China^{BC}, Colombia^{BC}, Costa Rica^{C}, Croatia^{BCD}, Czech Republic^{BCD}, Denmark^{ABCD}, Dominican Rep.^{BC}, Ecuador^{BC}, Egypt^{BC}, El Salvador^{C}, Estonia^{C}, Finland^{ABCD}, France^{ABCD}, Georgia^{CD}, Germany^{ABCD}, Greece^{BCD}, Guatemala^{C}, Honduras^{C}, Hong Kong^{BCD}, Hungary^{BCD}, Iceland^{C}, Indonesia^{BCD}, Ireland^{ABCD}, Israel^{BCD}, Italy^{ABCD}, Jamaica^{BC}, Japan^{ABCD}, Jordan^{CD}, Kazakstan^{BD}, Korea^{BCD}, Kuwait^{C}, Kyrgyz Republic^{D}, Latvia^{BCD}, Lithuania^{BCD}, Malaysia^{BC}, Mauritius^{C}, Mexico^{BCD}, Moldova^{C}, Morocco^{BCD}, Netherlands^{ABCD}, New Zealand^{ABCD}, Nicaragua^{C}, Norway^{ABCD}, Pakistan^{BCD}, Panama^{BCD}, Paraguay^{C}, Peru^{BC}, Philippines^{BCD}, Poland^{BCD}, Portugal^{ABCD}, Romania^{BCD}, Russian Federation^{BCD}, Singapore^{BCD}, Slovak Republic^{BCD}, Slovenia^{BCD}, South Africa^{BCD}, Spain^{ABCD}, Sri Lanka^{BC}, Sweden^{ABCD}, Syria^{C}, Switzerland^{ABCD}, Thailand^{BC}, Tunisia^{C}, Turkey^{BCD}, Ukraine^{BCD}, United Kingdom^{ABCD} , United States^{ABCD}, Uruguay^{BCD}, Venezuela^{BC}.

A: included in the OECD sample, B included in the large cross section, C: included in the large panel, D included in the skill specific unemployment regressions, large cross section.

A2. OECD sample

Unemployment rates For our OECD benchmark regressions we use total unemployment, measuring the percentage share of unemployed workers in total labor force (15 - 66 years old individuals). Data taken from Basanini and Duval. Original Source: OECD, Database on Labour Force Statistics; OECD, Annual Labour Force Statistics.

Openness measures Total trade openness is defined as imports plus exports divided by two times GDP in current prices. Real openness measures are constructed as respective current price openness measure times price level (taken from the Penn World Table 6.2) in order to account for the Balassa Samuelson effect by using real purchasing power GDP as denominator. Merchandise openness excludes services. The variable is taken from the WDI data base. Constant price total trade openness comes from the Penn World Table 6.2.

Wage distortion Wage distortion lumps replacement rate and tax wedge together. Both variables affect unemployment through the same channel, namely wages. Therefore lumping both variables together further reduces the number of instruments when estimating GMM regressions.

Replacement rate Average unemployment benefits taken from the Basanini and Duval data set. Original source: OECD Benefits and Wages Database. According to Basanini and Duval data is available for odd years only, so that they had to fill the gaps by linear interpolation.

Tax wedge This variable measures taxation on wages by computing the difference between wages paid by employers and wages earned by employees. The variable on tax wedge is constructed using the OECD taxing wages data. Some observations were adjusted by B&D in order to fill the gaps in the data, thus providing a complete sample for the period 1982 - 2003.

Union density Union density measures the percentage share of workers associated to unions. According to B&D the data was taken from the OECD Employment Outlook 2004 and inter / extrapolated in order to maximize the sample.

High corporatism Dummy variable that takes the value one if wage bargaining is highly centralized. Source: Basanini and Duval.

Table 8: **Summary statistics**

OECD panel			Large cross section			Large panel		
Variable	Mean	Std. Dev.	**Variable**	Mean	Std. Dev.	**Variable**	Mean	Std. Dev.
Unemployment (total)	7.532	3.890	Unemployment (WDI)	8.964	4.343	Unemployment (WDI)	8.343	4.372
Unemployment (prime)	6.631	3.323	Unemployment (CIA)	11.269	7.812	Unemployment (ILO)	8.378	4.290
Unemployment (youth)	15.122	8.029	Unemployment (IFS)	8.591	3.961	Unemployment (IFS)	8.225	4.041
Total trade(real)	34.466	18.598	Total trade(real)	26.679	25.947	Total trade(real)	27.092	22.315
Import (real)	33.533	17.193	Import (real)	26.332	24.526	Import (real)	26.809	21.049
Export (real)	35.398	20.162	Export (real)	27.026	27.466	Export (real)	27.375	23.749
Total trade(cur. P.)	32.879	16.036	Total trade(cur. P.)	41.508	31.111	Total trade(cur. P.)	38.387	24.607
Import (cur. P.)	32.206	15.142	Import (cur. P.)	41.345	29.742	Import (cur. P.)	38.612	24.027
Export (cur. P.)	33.552	17.157	Export (cur. P.)	41.671	32.698	Export (cur. P.)	38.162	25.723
Total trade (con. P.)	30.390	16.527	Total trade (con. P.)	39.924	28.231	Total trade (con. P.)	38.317	25.589
Total trade (merch.)	26.938	14.142	Total trade (merch.)	65.009	48.742	Total trade (merch.)	30.899	21.281
Wage distortion	58.142	17.825	EPL	0.487	0.187	LMR	−5.235	1.347
Replacement rate	29.430	12.572	Unemployment benefits	0.599	0.344	Unemployment benefits	−4.740	2.134
Tax wedge	28.712	8.928	Minimum wage	0.758	0.432	PMR	−5.833	2.261
Union density	40.263	20.768						
High corporatism	0.554	0.486						
EPL	2.086	1.092						
PMR	3.848	1.293						
Population	16.689	1.255	Population	9.863	1.377	Population	9.788	1.463
Output gap	−0.819	1.736	Output gap	−0.004	0.011	Output gap	0.000	0.033
			Land lockedness	0.486	0.348	Land lockedness	0.457	0.338
			Area	12.517	2.011	Area	12.475	1.892
			Latitude	3.314	0.997	Latitude	3.409	0.860
			Unofficial economy	28.619	13.851			
			F & R trade share	37.189	19.265			
TFP	266.838	6.902	TFP	1.199	0.432			
Skill specific unemployment rates								
			Low skill unemployment	11.379	6.499			
			High skill unemployment	6.842	5.868			
			Low/High skill endowment ratio	3.159	2.521			

EPL Measures the stringency of employment protection legislation, taken from Basanini and Duval. Original source: OECD, Employment Outlook 2004.

PMR Measures the regulation on product markets and competition, taken from Basanini and Duval. Original source: Conway et al. (2006).

Output gap Output gap measures the difference between actual and potential GDP as percentage of potential output. As source B&D cite the OECD Economic outlook and IMF International finance statistics.

A3. Large global cross country sample

Unemployment rate We use three different sources for total unemployment: The World Developing Indicators mainly provide official estimates on unemployment and are used as benchmark. Average unemployment rates constructed with less than 10 observations dropped. For additional robustness checks we include unemployment rates taken from the CIA factbook and IFS data base.

For our skill specific unemployment regressions we use data from the World Developing Indicators. We have percentage information on the fraction of total unemployment with primary, secondary, and tertiary skilled labor force. In order to derive specific skill-group unemployment rates, we construct skill specific total unemployment rates, multiply them with a measure on the total labor force in order to drive the number of skill specific unemployed workers, and divide by the number of workers belonging to the respective skill group (available in the WDI data base).

Openness measures See OECD sample data description for further details.

Frankel and Romer instrument (F&R) Our improved Frankel and Romer instrument bilateral trade data was used to regress total trade (exports plus imports) between country i and j, normalized by country $i's$ GDP, on geographical determinants of trade. The standard procedure is to take logs and estimate using OLS. Since $T_{ij} = 0$ for many country pairs, we follow Santos and Tenreyro (2006) and estimate (3) using Poisson pseudo maximum-likelihood. Predicting $\hat{T}_{ij}$ and summing over j, we have a measure of the trade share $\hat{T}_i$ that is by construction orthogonal to unemployment and hence a valid instrument.

EPL Employment laws index measuring the protection of labor and employment (EPL). The index variable includes: 1) Alternative employment contracts, 2) cost of increasing hours worked, 3) cost of firing workers and 4) dismissal procedures. Source: Botero et al. (2004).

Unemployment benefits Unemployment benefits is an index variable taken from Botero et al. (2004), including: 1) time of employment needed to qualify for unemployment benefits, 2) percentage of workers monthly income, paid to finance unemployment benefits, 3) waiting time on unemployment benefits, 4) percentage of income covered by unemployment benefits in case of a one year unemployment spell.

Minimum wage Dummy variable which takes the value one if there are binding minimum wages in the respective economy, taken from Botero et al. (2004).

Latitude Measures the distance between a country's capital and the equator. Data taken from the CIA factbook.

Area We control for the size of the economy in terms of its log area.

Land lockedness Land lockedness is constructed as index, measuring the length of neighboring borders relative to total length of borders.

Population We use Penn World Table 6.2 data on the size of population and take logs.

Unofficial economy This variable measures the size of the shadow economy, taken from Botero et al. (2004).

Output gap We construct output gap as difference between ln GDP and ln trend GDP, where the latter one is constructed by HP filtering the GDP data with smoothing factor 400. GDP is constructed as real GDP per capita (chain) times population taken from the Penn World Table 6.2.

A4. Large panel

Unemployment (u) See large cross section for further details. We also use unemployment rates from the ILO Laborsta database for robustness checks.

Openness measures See OECD data description for further details.

Labor market regulations (LMR) An index variable capturing labor market regulations. This index contains information on minimum wages, mandated hiring costs, unemployment benefits and other variables. Notice that higher index values indicate more freedom and thus lower labor market regulations. Higher values indicate more freedom in terms of less regulation. Between 1975 and 2000 data was estimated in 5-year intervals. From 2000 till 2006 yearly data is available. Source: Fraser Freedom of the World data set, 2008. Recoded by multiplying with -1.

Unemployment benefits Higher values indicate more freedom in terms of less regulation. Source: Fraser Freedom of the World Data set, 2008. Recoded by multiplying with -1.

Product market regulations (PMR) Taken from the Fraser freedom of the world database. We use price control as proxy for product market regulations. Higher values indicate more freedom in terms of less regulation. Source: Fraser Freedom of the World data set, 2008. Recoded by multiplying with -1.

Output gap See large cross section data description for more details.

Population See large cross section data description for more details.