Markups and the Euro

By David Cook[1]
HKUST Department of Economics
Clearwater Bay, Kowloon HKSAR

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Abstract: Recent theoretical work indicates that when firms set nominal prices in advance they are exposed to business cycle risk. Risk averse firms with market power may set relatively high markups over marginal cost to offset this risk. An appropriate monetary policy could reduce the risk of pre-set nominal prices. This paper uses industry-level panel data to estimate how the exchange rate regime affects the average markup in developed economies. This paper finds that the adoption of fixed exchange rates in the lead-up to European currency union is significantly associated with declines in labor share of income which is indicative of increased markups.

Keywords: Sticky Prices, Markups, Euro, fixed exchange rates
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I. Introduction

In this paper, I study the impact of exchange rate policy on industry-level labor share of income in developed economies. I find robust evidence that the adoption of fixed exchange rates by Western European economies during the process of developing a single European currency is associated with a drop in the share of income paid to workers. Under theories of monopolistic competition that form the basis for current models of nominal price rigidities, these declines in labor share might be associated with an increase in markups over marginal cost. In modern sticky-price models, price-setting firms choose their price level in order to maximize the expected utility entailed to their owners from the profits. Menu costs, however, prevent firms from adjusting their prices in response to the changes in the economic environment implying unpredictable profits. Obstfeld and Rogoff (1998) introduce the idea that profit-maximizing firm should take this risk into account in setting their level of markups. As Obstfeld (2000) notes, monetary policy can have first order effects on markups (and thereby first order effects on consumption, the terms of trade, and ultimately welfare) when macroeconomic risk affects the price setting decisions of firms or workers with market power.

A number of theoretical papers have focused on the effects of exchange rate policy in environments when the effect of risk on markups has first order macroeconomic effects. Devereux and Engel (2000) show that fixed exchange rates can implement an optimal monetary policy when exporters set their prices in foreign currency. Similarly, Corsetti and Pesenti (2001) show that the first order effects of exchange rate risk on the markups of producers implies that monetary policy-makers should put some weight on exchange rate stability. Bacchetta and van Wincoop (2000) find ambiguous effects of exchange rate stability on trade flows as fixed exchange rates may increase average price levels of monopolistic firms. Devereux, Engel, and Tille (2001) explicitly model the creation of a currency union in a sticky price open economy model and find that the adoption of a common currency in one region (similar to Europe) brings markups down to levels observed in a previously unified region (similar to the United States).

Kollman (2002) and Bergin and Tchakarov (2004) calibrate numerical dynamic general equilibrium models with sticky prices in which stochastic shocks lead to an increase in average markups. In a model with producer currency pricing (as in Obstfeld and Rogoff, 1995), Kollman finds that switching from a monetary target to a fixed exchange rate will increase average markups. By contrast, in a model with local currency pricing (as in Betts and Devereux, 1996) Bergin and Tchakarov (2004) find that average markups are lower under fixed exchange rates.

In a panel of industry-level annual data from 21 OECD countries over the period 1980-2002, the adoption of fixed exchange rates is associated with a decline in labor shares. I find that this effect holds true for most 1-digit sectors with the significant exception of commodity sectors in which monopolistic competition might be relatively weak. The effect holds true for most of the countries in the sample with the exceptions of Germany and Spain; the effect is especially strong and significant in Finland, France, Italy, and Ireland.

There is a natural tension when estimating the effects of fixed exchange rates on macroeconomic risk, since the nature of macroeconomic risk will likely affect the choice of exchange rate regime. This study controls for this possibility in two ways. First, the negative effect of fixed exchanges rates on labor share are derived with a within-effect estimator that focuses attention strictly on time series variation. In the panel data setting, we are able to control for fixed structural differences in the markups that may apply for reasons specific to each industry in each country. Second, all of the OECD economies that adopted fixed exchange rates in the sample period did so as part of the process of adopting the Euro. The adoption of a single currency by western European countries as part of a greater program of economic integration could be considered as being driven by political reasons more than reasons of business cycle stabilization. To the extent that this is true, the creation of the Euro may be as close to a natural experiment as is available for a broad set of countries.

In related empirical work, Broda (2004) finds that fixed exchange rates are associated with relatively high price levels in a broad set of countries. By contrast, this paper concentrates on evaluating the time series effects of the adoption of the Euro on markups in goods markets in developed economies. By focusing on markups, we are able to abstract from technological differences in marginal cost which may affect price levels for Belassa-Samuelson reasons. Conversely, with this approach, I am not able to assess whether exchange rate volatility affects markups in imperfectly competitive labor markets (as in Obstfeld and Rogoff, 2001)

II. The Model and the Data

A. The Model

Monopolistic firms in industry j in country i face an isoelastic demand curve and must set prices in advance.

$$y_{i,j,t} = X_{i,j,t} \cdot p_{i,j,t}^{-\lambda}$$

where $y_{i,j,t}$ and $p_{i,j,t}$ is the output of the representative firm in that sector and $X_{i,j,t}$ is random demand term which could be affected by macroeconomic variables. Firms face a constant cost of production $MC_{i,j,t}$ as they are price takers in input markets and produce goods with a constant returns to scale, Cobb-Douglas value added function with fully flexible factors of production. Firms pre-set prices and maximize:

$$E_{t-1}\left[\Omega_{i,t}\cdot\left\{p_{i,j,t}\cdot y_{i,j,t}-MC_{i,j,t}\cdot y_{i,j,t}\right\}\right]=E_{t-1}\left[\Omega_{i,t}X_{i,j,t}\cdot\left\{p_{i,j,t}^{1-\lambda}-MC_{i,j,t}\cdot p_{i,j,t}^{-\lambda}\right\}\right]$$

where $\Omega_{i,t}$ is the discount factor for profits. Define $Z_{i,j,t}=\Omega_{i,t}X_{i,j,t}$. Profit maximization is characterized by the first order condition:

$$\frac{1-\lambda}{\lambda}\cdot E_{t-1}\left[Z_{i,j,t}\right]=E_{t-1}\left[Z_{i,j,t}\frac{MC_{i,j,t}}{p_{i,j,t}}\right]\rightarrow$$

$$E_{t-1}\left[\frac{MC_{i,j,t}}{p_{i,j,t}}\right]=\frac{\lambda}{1-\lambda}-\frac{\lambda}{1-\lambda}\frac{Cov_{t-1}\left(Z_{i,j,t},\frac{MC_{i,j,t}}{p_{i,j,t}}\right)}{E_{t-1}\left[Z_{i,j,t}\right]}$$

Based on the Cobb-Douglas value added function, labor's share of value added, $lshare_{i,j,t}$ is proportional to real marginal cost (as in Gali and Gertler, 1999, and Sbordone, 2002).

$$lshare_{i,j}=\cdot\alpha_{i,j}\frac{MC_{i,j,t}}{p_{i,j,t}}\rightarrow$$

$$E_{t-1}\left[lshare_{i,j,t}\right]=\frac{\alpha_{i,j}\cdot\lambda}{(1-\lambda)}-\frac{\alpha_{i,j}\cdot\lambda}{(1-\lambda)}\frac{Cov_{t-1}\left(Z_{i,j,t},\frac{MC_{i,j,t}}{p_{i,j,t}}\right)}{E_{t-1}\left[Z_{i,j,t}\right]}\qquad(0.1)$$

Under certainty, average markups would be constant, $\frac{1-\lambda}{\lambda}$. Under uncertainty, expected labor share is a function of the conditional covariance between the markup and Z which includes both aggregate demand and the stochastic discount factor. Monetary policy affects average markups through its affects on the stochastic environment. The precise theoretical effects of any given monetary policy can only be determined in a full equilibrium model. The literature described in the introduction shows that the results can be sensitive to the specific setup of the model. Further, empirically measuring monetary policy is complicated. I focus on the average effect of the most observable monetary policy choice: the adoption of fixed exchange rates. I define $fix_{i,t}=1$ for countries that have permanently adopted a fixed exchange rate and estimate the panel data equation

$$lshare_{i,j,t} = \delta_{i,j} + \tau_t + \beta \cdot fix_{i,t} + \varepsilon_{i,j,t} \tag{0.2}$$

including time dummies and country-sector specific intercepts.

B. The Data

1. <u>Exchange Rate Regimes</u>

Reinhart and Rogoff (2004) classify monetary policies for a large number of countries including all OECD economies every year through year 2001 into five categories in terms of exchange rate stability. The categories include fixed exchange rates, crawling pegs, managed floats, free floats and freely falling rates. Define

- *fix*: Indicator variable equal to 1, if Reinhart and Rogoff (2004) classify the annual monetary policies of country i as a Fixed Exchange Rates in time t *and* every subsequent period through 2001. For 2002, we set the variable equal to 1 if the exchange rate is classified as fixed in 2001.

Japan and the United States have floating currencies for the entire period between 1980 and 2001 as does Germany prior to 1999 and Australia after 1983. Belgium operates a fixed exchange rate over the entire period. The remaining economies operate crawling pegs or managed floats in 1980.

In 1999, twelve European economies adopted a common currency. At various points prior to that date, some of these economies adopt exchange rate regimes which can be classified as fixed and maintain that regime until the end of the sample period. Table 1, Column A shows the dates at which the various economies in our sample[2] adopt the fixed exchange rate that culminates in the Euro currency. Denmark is classified as adopting a fixed exchange rate but to date has not joined the common currency. I classify *fix* for Denmark as equal to 1, because of their continuing fix to the Euro. In one case, the United Kingdom in 1991-2, a fixed exchange rate was adopted then abandoned. To focus attention on the Euro and a way from non-permanent regimes, we set *fix=0* for the UK for all periods.

2. <u>Labor Shares</u>

Data on markups is measured using labor shares of value added.

- *lshare*: Labors share of value added are obtained from the OECD's STAN Indicators database for 21 OECD economies for the aggregate economy, nine one-digit industries[3].

[2] These include Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Iceland, ** Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, United Kingdom and the United States. The average aggregate labor share for three other OECD countries, Greece, Mexico, and South Korea, are much lower and may be due to some measurement error, see Gollin, 2002.

[3] These include Agriculture, Fishing & Forestry, Mining, Manufacturing, Trade, Transport, Energy and Utilities,

Table 1, Column A shows aggregate labor shares for available years over the period 1980-2002 for 21 OECD countries. Though workers in most countries receive close to 60% of GDP on average, there is substantial variation ranging from about 47% in Italy to 63% in Sweden. Though not as large, there is also substantial variation in labor shares across time with the time average falling through the 1980's and 1990's as reported in Column B. I test the hypothesis that, in a panel of one digit industries for the 21 countries, *lshare* is driven by individual unit roots using the Im, Pesaren, and Shin (2003) W Statistic. The hypothesis is rejected at a 1% critical value. Column C reports industry-level data for labor shares at the one digit level. Here, the variations in labor shares are the largest with agriculture reporting a 20% labor share while personal services has an 80% market share.

Table 1, Column A also reports, for those countries that adopted the Euro, average labor shares after they adopted fixed exchange rates. Two countries labor shares displayed small increases (Spain and Portugal) while labor shares were less than the overall average for most economies. For a number of economies, including Finland, France, Ireland, and Italy, labor shares were substantially lower after they adopted fixed exchange rates. However, overall labor shares are largest in the earliest years, so this may indicate a time effect.

III. Empirical Results

A. Benchmark Results

I estimate equation *(1.1)* using an annual data pooled specification with 9 one digit sectors.

$$lshare_{i,j,t} = \delta_{i,j} + \tau_t + \beta \cdot fix_{i,t} + \varepsilon_{i,j,t} \qquad (1.1)$$

There are advantages gained by estimating the model with industry-level data beyond providing additional observations to improve the precision of the estimates. As different sectors have substantially different labor shares, aggregate labor shares are determined by the sector composition of production. By examining labor shares at the country level, we can avoid conflating the effects of fixed exchange rates on markups with changes in the sectoral composition of output.

The coefficient, β, is reported in Table 2, Panel A. The effect of the fixed exchange rate on labor share is negative and approximately equal to -2.4%. I compare standard errors that correct for heteroskedasticity with errors that allow for heteroskedasticity *plus* cross-sectional correlation of an unknown form. In each case, the coefficient β is significant at the .1% level. The standard errors are nearly identical which might indicate that controlling for

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Business Services, Personal and Social Services.

cross-sectional correlation is not crucial. I also calculate standard errors that are corrected to allow for time series dependence of unknown form. Following Wooldridge (2002), a panel test for residual auto-correlation can be obtained by regressing the residuals on their lag.

$$\varepsilon_{i,j,t} = e_{i,j} + e_t + \underset{(.049)}{.761}\varepsilon_{i,j,t-1} + \eta_{i,j,t}$$

It is possible to reject the hypothesis of zero-autocorrelation at any reasonable critical value. Therefore, it appears important to correct for auto-correlation. Table 2 reports heteroskedasticity, auto-correlation consistent estimates of the standard errors. The coefficient is significant at the 1% critical value with a p-value of about .0003.

The model is estimated with fixed cross-section and time effects because of the large observed differences in labor shares across countries, sectors, and time. I formally test, jointly and separately, the hypotheses that time fixed effects are equal to zero and cross-sectional fixed effects are zero. Each hypothesis is rejected at the 1% critical value. I also perform a Hausman test of the more efficient random effects estimator vs. fixed effects. The test does not reject the hypothesis that *fix* is exogenous to a random intercept term. In fact, the estimates of β are nearly identical when either fixed or random effects are assumed. Despite this I choose to use the fixed effects estimator which is more logically consistent with auto-correlated error terms. I also estimate a model in which there are separate time effects for European and non-European countries to control for the possibility that there is some Europe specific effect occurring over time. The coefficient estimate is $\beta = -\underset{(.006)}{.024}$ and is significant at the 1% critical value.

I consider some other specifications. Panel B reports β estimated with aggregate labor shares (using fixed time and cross-sectional effects and heteroskedasticity, auto-correlation consistent standard errors). The estimate is negative but slightly smaller (in absolute terms) than the disaggregated estimate. The much smaller sample size results in a more imprecise estimate; in this specification the estimate is significantly different from zero only at a 15% critical value. One reason for the lack of significant results in the aggregate data is that negative effects of fixed exchange rates on labor shares are concentrated in industries of a relatively small size. Panel C shows the results of a weighted OLS regression of *(1.1)* using the average share of a sector in generating a country's value added over 1980-2002 as weights. The coefficient estimate is $\beta = -\underset{(.009)}{.019}$ and is significant at the 5% critical value. The weighted averages indicate that though the negative effects of fixed exchange rates might be concentrated in relatively smaller sectors, the aggregate effects are still substantial and statistically significant.

With a mean labor share of .514, the average percentage increase in the markup suggested by a decrease in labor share of -.024 is about 4.8%. Panel D reports an estimate using the log of *lshare* in the possibility that a constant percentage effect is a better specification. The coefficient on *fix* is significant at the 1% critical value; the estimate is -.062 which is approximately consistent with the benchmark result.

B. Variation in Effects by Sector

There are theoretical reasons to think that the effect of fixed exchange rates on markups may differ by sector. Devereux and Engel (2000) show that a fixed exchange rate may allow perfect risk sharing when traded goods are priced in the currency of the customer. In general, the risk associated with the pricing of goods subject to international trade might be different than non-traded goods. In addition, commodity goods which are sold in perfectly competitive markets may not be subject to the pricing risk faced by monopolistically competitive firms that set their prices in advance.

There are nine one-digit sectors. No particular sector is generating the leverage which produces the negative relationship between fixed exchange rates and labor shares in developed economies. I estimate the regression in *(1.1)* dropping each of the 9 sectors in turn. In each of the nine regressions, the estimate of β is below -.015 and significant at the 1% critical value.

To assess whether the effects of fixed exchange rates differ across sectors, I examine the specification:

$$lshare_{i,j,t} = \delta_{i,j} + \tau_t + \beta_T \left(1_j^T \cdot fix_{i,t} \right) + \beta_{NT} \left(1_j^{NT} \cdot fix_{i,t} \right) + \varepsilon_{i,j,t}$$

defining 1_j^T as equal to one when the sector is manufacturing, mining, or agriculture while 1_j^{NT} is one otherwise. Table 2, Panel E reports the estimates of $\beta_T = -.014$ and $\beta_{NT} = -.029$. Both coefficients are negative. Indeed, the effect of *fix* is only significant (and at the 1% critical value) in the non-traded sector. However, a Wald test is unable to reject the hypothesis that the two coefficients are the same at even the 10% critical value.

Estimates for the effect of fixed effects on labor share at higher levels of disaggregation are reported in Table 2, Panel F. The parameter β is allowed to vary across 1 digit sectors. The effects of fixed exchange rates are positive in two traded goods sectors characterized by standardized commodities, agriculture and mining. There is also a slightly positive effect in the Trade sector which represents Wholesale and Retail Trade, Restaurants, and Hotels. None of the coefficients are significantly positive. In other sectors, there are negative coefficients which are either comparable in size or larger than in the Benchmark regression. Coefficients that are negative at the 10% critical value or lower include Manufacturing, Utilities, Construction, Transport, and Personal Services.

A Wald test of the hypothesis that the coefficient on each of the nontraded goods is equal is rejected at a p-value of .13. A test of the hypothesis that the coefficients on each of the traded goods sectors are equal is rejected at a p-value of .03. We cannot reject the hypothesis that the coefficients on *fix* are the same or the two non-commodity sectors but we can still reject the hypothesis that the 7 non-commodity sectors have the same coefficient β.

Panel G reports the estimates from a specification of the form:

$$lshare_{i,j,t} = \delta_{i,j} + \tau_t + \beta_{CY}\left(1_j^{CY} \cdot fix_{i,t}\right) + \beta_{NCY}^{l}\left(1_j^{NCY} \cdot fix_{i,t}\right) + \varepsilon_{i,j,t}$$

Where 1_j^{CY} is defined as an indicator variable that is 1 for the commodity sectors of Agriculture and Mining, and 1_j^{NCY} is one for the noncommodity sectors. The effect of fixed exchange rates on the commodity sector is positive, $\beta = .003$ but insignificant; the effect of fixed exchange rates on the non-commodity sector is negative and significant at the 1% critical value. A Wald test can reject the hypothesis that the two coefficients are equal at the 10% critical value.

C. Credibility

Though many European countries fixed their currencies at various times over the sample period, the Euro was not permanently adopted until late in the 1990's. Fully credible fixed exchange rates like the Euro may have less negative effects on markups than the less credible exchange rate arrangements that preceded it. Table 3, Column A reports the estimates of:

$$lshare_{i,j,t} = \delta_{i,j} + \tau_t + \beta_{80's}\left(1_t^{80's} \cdot fix_{i,t}\right) + \beta_{90's}\left(1_t^{90's} \cdot fix_{i,t}\right) + \varepsilon_{i,j,t}$$

where $1_t^{80's}$ is an indicator dummy when *t = 1980-1990* and $1_t^{90's}$ is an indicator dummy when *t = 1991-2002*. Each of the parameters is significant at the 1% critical value. The best estimate is that the effect of fixed exchange rates on markups is stronger in the 1980's than the 1990's: $\beta_{80's} \approx -0.030, \quad \beta_{90's} \approx -0.023$. However, a $\chi^2(1)$ Wald statistic of the hypothesis that $\beta_{80's} = \beta_{90's}$ equals .656 which fails to reject the hypothesis at the 40% critical value.

I also directly test whether the adoption of the Euro currency union has different effects on labor share than does some other form of fixed exchange rates. Table 3, Column B reports the estimates of

$$lshare_{i,j,t} = \delta_{i,j} + \tau_t + \beta_{EZ}\left(1_{i,t}^{EZ} \cdot fix_{i,t}\right) + \beta_{NotEZ}\left(1_{i,t}^{NotEZ} \cdot fix_{i,t}\right) + \varepsilon_{i,j,t}$$

where $1_{i,t}^{EZ}$ is an indicator dummy which equals one for euro zone countries (which does not include Denmark at any period) in the period 1999-2002 and zero otherwise, while $1_{i,t}^{NotEZ} = 1 - 1_{i,t}^{EZ}$. Both coefficients are similar in size to the estimates for the whole sample reported in Table 2, Panel A. The estimated effect is actually stronger in the actual Euro-zone period. **Devereux shows that a bilateral exchange rate peg may reduce pricing flexibility relative to a one-sided peg.** However, a $\chi^2(1)$ Wald statistic of the hypothesis that $\beta_{EZ}^1 = \beta_{NotEZ}^1$ equals .412, which fails to reject the hypothesis at the 50% critical value.

Credibility of the fixed exchange rate might also be achieved through durability. Another specification allows the coefficient to vary across time as a function of the number of years since the adoption of a fixed exchange rate, $\beta_{i,t}^1 = \gamma^0 + \gamma^1 \cdot length_{i,t}$. Define $length_{i,t}$ as the number of years that the exchange rate has been in place, with $length_{i,t}$ equal to 1 in the first period of the fixed exchange rate. Table 3, Column C reports the estimates of the regression

$$lshare_{i,j,t} = \delta_{i,j} + \tau_t + \gamma^0 \cdot fix_{i,t} + \gamma^1 \cdot length_{i,t} \times fix_{i,t} + \varepsilon_{i,j,t}$$

There is little evidence that γ^1 is significantly different than 0. Moreover, the inclusion of the interaction term between *fix* and *length* has little effect on the estimate of the direct effect *fix*.

I also estimate equation *(1.1)* using 5 year averages at non-overlapping 5 year intervals (including 1985, 1990, 1995, 2000). To clarify, the left-hand variable is the average over the previous 5 years of *lshare* and the right hand variable is the number of years during that period that the economy had a fixed exchange rate divided by five. The estimate of $\beta^2 \approx -.036$ is slightly larger than in the yearly static equation and is significant at the 1% critical value (using HAC corrected standard errors) despite the much smaller sample size.

D. Dynamics and Business Cycles

It is possible that there are business cycle effects on markups. We want to control for this possibility since business cycle effects may generate cross-sectional dependence in the residuals which could complicate inference. To control for business cycles, we include a term *Output Gap*$_{i,t}$ which is the percentage deviation at year t of constant price GDP per capita in country i from the Hodrick-Prescott trend which is calculated over the period 1960-2002.

$$lshare_{i,j,t} = \delta_{i,j} + \tau_t + \beta^1 fix_{i,t} + \beta^2 \cdot Output \quad Gap + \varepsilon_{i,j,t} \qquad (1.2)$$

I estimate equation (1.2) with pooled OLS with time and cross-section fixed effects in order to best control for the joint shocks. The results are reported in Table 4, Column A. The coefficient estimate of β^1 is little changed from the benchmark result in Table 2, Column A.

The coefficient estimate of β^2 is not statistically significant.

It might be argued that the output gap is not strictly exogenous to unobserved time series shocks to markups. Therefore, I account for fixed effects with Arellano and Bover's (1995) orthogonal deviations method and estimate equation (1.2) with 2SLS with 3 lags of the $Output \;\; Gap$ as instruments.[4] The estimates are in Table 4, Column B. Standard errors are again corrected for auto-correlation and heteroskedasticity. Again I find that β^2 is not statistically significant. Even after controlling for business cycles, the coefficient β^1 is still negative and statistically significant at the 1% critical value.

The theory suggests that the reason average markups vary across exchange rate regimes is that fixed exchange rates impact the macroeconomic risk faced by firms in setting their optimal markup. I am also interested in determining whether the effects of business cycles on markups are actually different under fixed exchange rates. The results of the estimation of

$$lshare_{i,j,t} = \delta_{i,j} + \tau_t + \beta^1 fix_{i,t} + \beta^2 \cdot Output \;\; Gap_{i,t} + \beta^3 \cdot fix_{i,t} \cdot Output \;\; Gap_{i,t} + \varepsilon_{i,j,t}$$

where the direct impact of business cycles on markups differs across exchange rate regimes are reported in Table 4, Column C. In this regression, I find that after controlling for regime specific business cycle effects, the average impact of fixed exchange rates on labor shares, β^1, is still significant at the 1% critical value. Interestingly, we find that the difference between fixed exchange rate regimes and floating regimes in terms of the business cycle impact on labor share is significant; the parameter β^3 is negative and significant at the 1% critical value. This is consistent with the idea that exchange rate regimes are an important determinant of macroeconomic risk for markups. However, we should be careful not to over-interpret the sign of β^3 because our simple model may not have fully accounted for all of the dynamic interactions between business cycles and markups.

I estimate a dynamic panel version of *(1.1)* in which *lshare* follows an AR(1) process.

$$lshare_{i,j,t} = \delta_{i,j} + \tau_t + \rho^1 fix_{i,t} + \rho^2 \cdot lshare_{i,j,t-1} + \varepsilon_{i,j,t} \qquad (1.3)$$

The estimates of the parameters in equation (1.3) (again derived with 2SLS using the Arellano-Bover orthogonal deviations method) are reported in Column D. The instruments are three lags of $lshare_{i,j,t-2} .. lshare_{i,j,t-4}$. The fixed exchange rate indicator, *fix*, is treated as exogenous. The standard errors are again corrected for auto-correlation and heteroskedasticity. Movements in the labor share are strongly auto-correlated. The short-term effect of fixed

[4] Efficient GMM coefficient estimates are similar in size and have smaller standard errors.

exchange rates is smaller than those estimated in the static equation from *(1.1)* though they are significant at the 5% level (with HAC corrected standard errors). However, the estimated steady state effect, $\rho^1 \big/ 1-\rho^2$, implied from the dynamic equation is very similar to the estimate of β^2 in Table 1, Column A. In Column E, I report the estimates of a dynamic version of the model controlling for business cycle effects. The qualitative results hold even after explicitly after controlling for the business cycle.

E. Cross-Country Differences

No single country is responsible for the leverage that generates the negative and statistically significant estimate of β . I re-estimate equation *(1.1)* twenty-one times, successively dropping from the sample each of the twenty-one countries. In each case, the estimate of β is negative, near the full sample estimate and significant at the 1% critical value.

I estimate a specification in which the effect of fixed exchange rates on labor share is country specific for the nine European countries for which exchange rates were fixed after 1981. The findings are in Table 5, Column A. The coefficients are negative in seven of the nine countries and of a size comparable to the benchmark results. The coefficients are significant at the 10% level or lower in Finland, France, Italy, and Ireland. The results are positive) only in Germany and Spain and these are not statistically significant. A Wald test of the hypothesis that the coefficients are the same in all countries is rejected at the 5% critical value.

Spain and Germany are relatively large economies. A large economy may be less subject to macroeconomic risk for a number of reasons. Greater size may allow for greater diversification of macroeconomic risk. Unlike a small open economy, the Euro-zone as a whole does have an independent monetary policy. This monetary policy may be more responsive to economic conditions in the larger economies. Also, large economies might also rely less on external trade and may be less exposed to the impact of a fixed exchange rate. Another factor that may cause the effects of fixed exchange rates to differ across countries is fiscal policy. In lieu of monetary policy, changes in taxes or government spending might cushion an economy from macroeconomic shocks. The 1997 Growth and Stability Pact limit the ability of economies to run budget deficits. Countries with budget flexibility may face less macroeconomic risk.

Table 5, Column B reports estimates from a specification in which the effect of the fixed exchange rate is allowed to vary across countries:

$$share_{i,j,t} = \delta_{i,j} + \tau_t + \beta_1^i fix_{i,t} + \varepsilon_{i,j,t}$$
$$\beta_1^i = \alpha_0 + \alpha_1 \cdot GDP_i^{80} + \alpha_2 \cdot budgetdeficit_i + \alpha_3 exshare_i$$

where GDP^{80} is 1980 Gross Domestic Product in US dollars relative to the United States in the same year; *budgetdeficit* is the average of budget deficit as a share of GDP in the years before the Stability Pact.(1980-1996); *exshare* is exports as a share of GDP in the years (1980-2001).

The results suggest that size per se is not related to the effect of fixed exchange rates on labor share; the coefficient on GDP^{80} is insignificant. However, the coefficient on *budgetdeficit* is negative and significant at the 5% critical value. Countries that had run large budget deficits before the Growth and Stability Pact faced a more negative effect of fixed exchange rates on labor share. The coefficient on exports as a share of GDP is also insignificant. Honohan and Lane (2003) note differentials in inflation that depend on the degree that country has *outside* the EU. I split *exshare* into two parts, exports to the European Monetary Union, $exshare^{EU}$, as a share of GDP and exports to the rest of the world, $exshare^{NOT\_EU}$.. The resulting coefficient estimates are in Table 5, Column C. In this regression, high levels of trade outside the EU results in a much sharper effect of fixed exchange rates on labor share.

The effects of fixed exchange rates on markups may also depend on how much monetary policy changes when the economy joins the fixed exchange rate regime. If the optimal policy for a given economy is similar to aggregate ECB policy, the effect of joining the Euro might be most positive. First, I measure, Corr(mm,mm^{DM}), the quarterly correlation of money market interest rates in a given country with the money market interest rates in Germany measured over the period 1975 until the year in which the economy fixed its exchange rate. If the interest rate moved closely with the Mark before the adoption of the Euro, fixing the exchange rate may have implied little change. I also measure, Corr(GDP^{i},GDP^{EMU}) where GDP is annual HP filtered constant price local currency GDP and GDP is annual HP filtered constant price GDP in the EMU. A country that shares the business cycles of the aggregate area, may face little cost in fixing its currency to the Euro. I estimate a specification with

$$\beta_1^i = \alpha_0 + \alpha_1 \cdot GDP_i^{80} + \alpha_4 \cdot Corr(mm^i, mm^{DM}) + \alpha_5 \cdot Corr(GDP^i, GDP^{DM})$$

The results are reported in Table 5, Column 4. None of the coefficients (with the exception of α_0) are significant at the 10% critical value.

IV. Conclusion

I find evidence that the adoption of a fixed exchange rate by European economies leading up to the adoption of the Euro is strongly and robustly associated with a decline in labor shares of value added at the one-digit industry level. One explanation is that a monetary policy that does not adequately adjust to macroeconomic shocks leads to risky profits for monopolistic firms setting sticky prices. Firms may respond to increased risk by increasing their average markups and reducing aggregate economic efficiency. The effect on labor share does not appear in commodity based sectors which may not be well described by monopolistic competition. The effect of the fixed exchange rate on labor share seems to be positive for Germany, but negative for all other countries but Spain.

The effects of the adoption of the fixed exchange rates on markups are substantial in the range of a 5-6% increase. Quantitatively, this rise is much larger in scale than the effects of changes in monetary policy on steady state markups in calibrated models.

This study can be read in light of a larger body of literature which studies the macroeconomic effects of goods market competition in the EU. Blanchard (2004) argues that economic reform in European goods markets has increased competition and reduced monopoly power. Bayoumi, Paxton and Pesenti (2004) calibrate a model of the European economy to demonstrate that an increase in goods and labor market competitiveness, represented by a substantial decline in markups would substantially narrow the gap between the EU and the United States.

V. Data Appendix

Data on labor shares and industry shares of value added are from the OECD STAN Indicators database. Data on fixed exchange rate regimes are from Reinhart and Rogoff (2004). Data on (1) constant (1995) prices GDP per capita used to construct the business cycle indicator; (2) 1980 GDP in PPP adjusted international dollars; (3) exports as a share of GDP are all from World Development Indicators. Data on exports to the EU and exports to the World in US dollars are from IMF Direction of Trade Statistics. Data on money market interest rates are from IMF's International Financial Statistics.

VI. Bibliography

Arellano, M. and Bover, O. (1995), "Another Look at the Instrumental Variables Estimation of Error Components Models," Journal of Econometrics 68, 29-51.

Bacchetta, P and E. van Wincoop, 2000, "Does Exchange-Rate Stability Increase Trade and Welfare?," American Economic Review 90,1093-1109

Bayoumi, T, D. Laxton, and P. Pesenti, 2004, "Benefits and Spillovers of Greater Competition in Europe: A Macroeconomic Assessment." NBER Working Paper 10416.

Bergin, P.R. and Tchakarov, I., 2004, "Does Exchange Rate Risk Matter for Welfare," NBER Working

Broda, Christian, 2004, "Exchange Rate Regimes and National Price Levels," Mimeo. Federal Reserve Bank of New York.

Betts, C. and M.B. Devereux, 1996, "The Exchange Rate in a model of Price to Market," European Economic Review 40, 1007-1021.

Blanchard, O., 2004, The Economic Future of Europe, NBER Working Paper 10310.

Corsetti, Giancarlo and Paolo Pesenti, 2001, "International Dimensions of Monetary Policy," NBER Working Paper 8230

Broda, C., 2004, "National Price Levels and Fixed Exchange Rates," Mimeo. Federal Reserve Bank of New York.

Devereux, M. B. and C.Engel, 2000, "Monetary Policy in the Open Economy Revisited: Price Setting and Exchange Rate Flexibility," Review of Economic Studies 70, 765-83.

Devereux, M.B., C.Engel, and C. Tille, "Exchange Rate Pass-Through and the Welfare Effects of the Euro," International Economic Review 44, 223-44.

Gali, J. and M. Gertler, 1999, Inflation dynamics: A structural econometric analysis" Journal of Monetary Economics 44, 195-222.

Gollin, D., 2002, "Getting Income Shares Right," Journal of Political Economy 110, 458-74.

Honohan, P. and P. Lane, 2003, Divergent Inflation Rates in EMU, Economic Policy 37, 357-387.

Im, K.S., M.H. Pesaran, and Y. Shin, "Testing for Unit Roots in Heterogenous Panels," Journal of Econometrics 115, 53-74.

Kollmann, R., 2002, "Monetary policy rules in the open economy: effects on welfare and business cycles," Journal of Monetary Economics 49, 989-1015.

Obstfeld, M., 2001, "International Macroeconomics: Beyond the Mundell-Fleming Model," NBER Working Paper 8369.

Obstfeld, M. and K. Rogoff, 2000, New Directions for Stochastic Open Economy Models, Journal of International Economics 50, 117-53.

Obstfeld, M. and K. Rogoff, 1998, Risk and Exchange Rates, NBER Working Paper 6648.

Obstfeld, M. and K. Rogoff, 1995, “Exchange Rate Dynamics Redux", Journal of Political Economy

Reinhart, C. and K. Rogoff, 2004, “The Modern History of Exchange Rate Arrangements: A Reinterpretation.” Quarterly Journal of Economics 119, 1-48.

Sbordone, A.M., 2002, “Prices and unit labor costs: a new test of price stickiness,” Journal of Monetary Economics 49, 265-292.

Wooldridge, J.M., 2002. “Econometric Analysis of Cross Section and Panel Data,” MIT Press. Cambridge.

Table 1. Column A reports the countries included in the sample, the years in which various economies adopted fixed exchange rates, the average aggregate labor share, and the aggregate labor share during the fixed exchange rate period. Column B reports the years included and the average labor share in those time periods. Column C reports the sectors included and the average labor share of those sectors.

Labor Shares					
A. By Country		B. By Period		C. By Sector	
USA	0.577	1980	0.541	Manufacturing	0.625
Australia	0.532	1981	0.544	Agriculture	0.223
Austria (1981)	0.573 (0.572)	1982	0.534	Mining	0.361
Belgium (1980)	0.564 (0.564)	1983	0.522	Utilities	0.327
Canada	0.595	1984	0.513	Construction	0.639
Denmark (1999)	0.612 (0.608)	1985	0.509	Trade	0.578
Finland (1995)	0.585 (0.550)	1986	0.514	Transport	0.584
France (1997)	0.577 (0.564)	1987	0.515	Personal Services	0.801
Germany (1999)	0.590 (0.580)	1988	0.511		
Iceland	0.574	1989	0.506	Business Services	0.457
Ireland (1997)	0.508 (0.462)	1990	0.510		
Italy (1997)	0.472 (0.442)	1991	0.517		
Japan	0.515	1992	0.519		
Netherlands (1983)	0.560 (0.554)	1993	0.515		
New Zealand	0.478	1994	0.506		
Norway	0.530	1995	0.495		
Portugal (1994)	0.526 (0.534)	1996	0.497		
Spain (1994)	0.515 (0.526)	1997	0.498		
Sweden	0.626	1998	0.500		
Switzerland	0.577	1999	0.506		
UK	0.601	2000	0.507		
		2001	0.505		

Table 2. Panel A reports benchmark results from an unbalanced pool regression of labor shares of income, *lshare*, from nine 1-digit sectors in 21 OECD countries over the period 1980-2002 on a indicator variable for fixed exchange rates leading to the adoption of the Euro, *fix*. Panel B reports the result when aggregate labor shares are used. Panel C reports the result when the sector's value added share in aggregate value added is used in a weighted OLS regression. Panel D reports the result when the natural log of *lshare* is used as a left hand side variable Panel E reports the result when the effect of fixed exchange rates is allowed to be different in traded and nontraded goods sectors. Panel F reports the result when the effect of fixed exchange rates is allowed to vary between commodity sectors and non-commodity sectors. Panel G shows the results when the effect of a fixed exchange rate is sector-specific. All regressions include 189 country-sector dummies as well as time dummies. All standard errors and Wald statistics are corrected for heteroskedasticity and sector-country level auto-correlation of unknown kind. Coefficients denoted ♠ are significant at the 1 % critical value; ♥ at the 5% critical value; and ♦ at the 10% critical value.

Dependent Variable: *lshare*

(A) Panel	
Fix	-0.024♠
	(.006)
	N=4041
(B) Aggregate	
Fix	-0.016
	(0.010)
	N =459
(C) Weighted OLS	
Fix	-0.019♥
	(.009)
	N = 4041
(D) log *lshare*	
Fix	-0.062♠
	(.019)
	N = 4041

Coefficient β^1 varies by:

(E) Sector Category	
Traded Goods×fix	-0.014
	(.013)
NontradedGoods×fix	-0.029♠
	(.008)
	N =4041
Wald	.882
	(.345)

(F) Sector Category	
Commodities×fix	0.003
	(.017)
Non-commodities×fix	-0.031♠
	(.007)
	N = 4041
Wald	2.915
χ^2 (1) p-value	(.088)

(G) 1 Digit Sector	
Manufacturing×fix	-0.043♥
	(.019)
Agriculture×fix	0.003
	(.013)
Mining×fix	0.002
	(.039)
Utilities×fix	-0.045♦
	(.025)
Construction×fix	-0.041♥
	(.017)
Trade×fix	0.002
	(.015)
Transport×fix	-0.061♠
	(.014)
Personal Services×fix	-0.025♠
	(.009)
Business Services×fix	-0.002
	(.015)
	N = 4041
Wald Test	22.256♠
χ^2 (1) p-value	(.008)

Table 3 Panel A shows the results of a panel regression of labor share on a fixed exchange rate indictor when the effects are allowed to be different in the period *1980s* (1980-1990) from the effects in the *1990s* (1991-2002). Panel B shows the result when the effect of fixed exchange rates is allowed to differ between actual adoption of the Euro with *€zone* as an indicator for Euro zone countries between 1999 and 2002 and fixed exchange rates leading to the adoption of the Euro before 1999 and Denmark afterward. Column C shows the result when the effect of fixed exchange rates is allowed to vary with the number of periods that the fixed exchange rate regime has been in place, *length*. Column D are the results from a regression of non-overlapping 5 year averages of labor shares and the fixed exchange rate indicator.
All regressions include country-sector dummies and time dummies. All standard errors and Wald statistics are corrected for heteroskedasticity and sector-country level auto-correlation of unknown kind. Coefficients denoted ♠ are significant at the 1 % critical value; ♥ at the 5% critical value; and ♦ at the 10% critical value.

	Dependent Variable: *lshare*			
	Variations In Credibility			5 Year Averages
	(A)	(B)	(C)	(D)
fix			-0.026♠ (.007)	-0.036♠ (.010)
1980's×*fix*	-0.030♠ (.009)			
1990's×*fix*	-0.023♠ (.007)			
€zone×*fix*		-.028♠ (.011)		
Not€zone×*fix*		-.023♠ (.006)		
length×*fix*			0.001 (.001)	
Wald χ^2 (1) p-value	.656 (.441)	.412 (.521)		
N	4041	4041	4041	712

Table 4 Column A reports the result when the business cycle, measured by log Hodrick-Prescott filtered real GDP, can directly affect labor share estimated with Fixed Effects OLS. Column B reports the results of the same model estimated with 2SLS-Orthoganol deviations using 3 lags of the output gap. Column C reports the results of a model in which the effect of the business cycle on labor shares differs across exchange rate. Column D reports the results of a dynamic panel model. Column E reports the business cycle model adjusted for dynamics. The models in Column (B) through (E) are estimated with 2SLS-Ortoganol deviations fixed effects. All regressions include time dummies. All standard errors are corrected for heteroskedasticity and sector-country level auto-correlation of unknown kind. Coefficients denoted ♠ are significant at the 1 % critical value; ♥ at the 5% critical value; and ♦ at the 10% critical value.

	Dependent Variable: *lshare*				
	Controlling For Business Cycle Dynamics (Estimation Methods)				
	(OLS, FE)	(2SLS, Orthoganol Deviations)			
	(A)	(B)	(C)	(D)	(E)
fix	-0.024♠ (.006)	-0.023♠ (.007)	-0.025♠ (.007)	-0.007♥ (.003)	-0.007♠ (.003)
Outputgap	0.006 (.056)	.108♦ .060	0.146♥ (.063)		0.180♠ (.029)
Outputgap×*fix*			-0.551♠ (.162)		-0.219♥ (.079)
$lshare_{-1}$				0.734♠ (.070)	0.726♠ (.041)
N	4041	3336	3336	3669	3323

Table 5 In Column A, the coefficient varies by country. In Column (B) & (C), the coefficient varies according to country specific macroeconomic variables including GDP in 1980 relative to the US, GDP^{80}; average budget deficit as a share of GDP, *budget deficit;* exports as a share of GDP, *exshare*; exports to EU countries as a share of GDP, $exshare^{EU}$; and exports to non-EU countries as a share of GDP, $exshare^{NOT\_EU}$. In Column D, the coefficient varies by country specific macroeconomic moments including correlation of money market rates with the Deutsche Mark rate, Corr(mm,mm^{DM}); and correlation of detrended GDP with EMU GDP, $Corr(GDP,GDP^{EMU})$.

Dependent Variable: *lshare*					
Country Specific		Country Macro Characteristics			
	(A)		(B)	(C)	(D)
Denmark × *fix*	-0.020	*fix*	0.016	0.069	-0.034
	(.019)		(.026)	(.112)	(.011)
Finland × *fix*	-0.044♠	GDP^{80} × *fix*	-0.071	0.076	0.055
	(.01)		(.116)	(.109)	(.13)
France × *fix*	-0.025♦	*budgetdeficit* × *fix*	-0.628♥	-0.669♥	
	(.015)		(.293)	(.294)	
Germany × *fix*	0.033	*exshare* × *fix*	-0.044		
	(.038)		(.054)		
Ireland × *fix*	-0.056♥	$exshare^{EU}$ × *fix*		0.066	
	(.023)			(.062)	
Italy × *fix*	-0.059♠	$exshare^{NOT\_EUS}$ × *fix*		-0.470♥	
	(.022)			(.214)	
Netherlands × *fix*	-0.017	$Corr(GDP^{i},GDP^{EMU})$ × *fix*			0.003
	(.015)				(.024)
Portugal × *fix*	-0.026	$Corr(mm,mm^{DM})$ × *fix*			0.014
	(.02)				(.019)
Spain × *fix*	0.009				
	(.009)				