Economic Insecurity and the Globalization of Production*

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A common claim in debates about globalization is that economic integration increases worker insecurity. Although this idea is central to both political and academic debates about international economic integration, the theoretical basis of the claim is often not clear. There is also no empirical research that has directly tested the relationship. In this paper, we argue that economic insecurity among workers may be related to riskier employment and/or wage outcomes, and that foreign direct investment may be a key factor contributing to this increased risk by making labor demands more elastic. We present new empirical evidence, based on the analysis of panel data from Great Britain collected from 1991-1999, that FDI activity in the industries in which individuals work is positively correlated with individual perceptions of economic insecurity. This relationship holds in yearly cross-sections, in a panel accounting for individual-specific effects, and in a dynamic panel model also accounting for individual-specific effects.

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1. Introduction

Across the world, there appears to be significant opposition to policies aimed at further liberalization of international trade, immigration, and foreign direct investment (FDI). A large number of political events in recent years suggest a marked turn away from liberalization, and many prominent observers have raised alarms about this "globalization backlash."[1]

There is a growing body of research examining what political-economy forces underlie this backlash. In Scheve and Slaughter (2001 a, b), we documented a strong cleavage between labor-market skills and U.S. public preferences over trade and immigration policy. Less-skilled individuals, measured by educational attainment or wages earned, are much more likely to oppose freer trade and immigration than their more-skilled counterparts. This finding is consistent with the distributive consequences of liberalization predicted by the factor endowment model. We also found that many other possible cleavages, surprisingly, do not materialize. Across both trade and immigration preferences, no other cleavage is as consistently important as the skills divide. Subsequent research has documented this division in preferences about international economic liberalization across a wide number of countries, where the magnitude of the cleavage across countries varies in predictable ways according to national endowments and labor-market institutions (O'Rourke and Sinnott 2001; Mayda and Rodrik 2001; Hayes, Ehrlich, and Peinhardt 2002; Dancygier 2002; Mayda 2002; Gabel 1998a, b; Scheve 2000). Thus, the distributive consequences in the labor market for individuals with varying endowments of human capital seem to be one factor contributing to the backlash against globalization.

[1] Examples of political events include public protests at virtually every meeting of international economic institutions beginning with the 1999 WTO meetings in Seattle. In August, 2000, Federal Reserve Chairman Alan Greenspan acknowledged that liberalization efforts have stalled out, with outbreaks of protectionism a distinct possibility: "'Despite extraordinary prosperity, the ability to move forward on various trade initiatives has clearly come to a remarkable stall … there remains considerable unease among some segments [of society] about the way markets distribute wealth and about the effects of raw competition on society … it is quite imaginable that support for market-oriented resource allocation will wane and the latent forces of

These distributive arguments, however, largely focus on the relationship between international economic integration and the level and distribution of wages. Another body of research has focused on whether globalization has increased individual *economic insecurity*. This line of research recognizes that risk-averse workers are concerned not only about the level but also the volatility of their earnings—in particular, volatility from the risk of unemployment.

Economic insecurity may contribute to the globalization backlash in at least two ways. First, there is the potential for a direct effect, very similar to that documented for the skill divide in opinion formation about international economic policies. Individuals that perceive globalization contributing to their own economic insecurity are much more likely to develop policy attitudes hostile towards economic integration. Second, a number of scholars have argued that increases in economic insecurity from globalization may generate demands for more generous social insurance that compensates workers for a riskier environment (e.g. Rodrik 1997; Garrett 1998; Burgoon 2001; Hayes, Ehrlich, and Peinhardt 2002; Boix 2002). However, many scholars have also suggested that globalization limits the capacities of governments to provide such compensation (e.g. Rodrik 1997; Desai 1999; Besley, Griffith, and Klemm 2001). Thus, individuals may develop concerns about globalization because they believe it reduces the insurance provided by the state for all labor market risks, including those heightened by global integration.

The claim that international economic integration increases economic insecurity and thus contributes to political conflict over globalization is, however, highly contested. Rodrik (1997) presents evidence that exposure to external risk, measured by the interaction between trade openness and the standard deviation of a country's terms of trade, is positively correlated with

protectionism and state intervention will begin to reassert themselves in many countries, including the United States.'" (Stevenson 2000).

both growth volatility and government expenditures. From this Rodrik concludes that globalization increases economic insecurity and thus demands for a robust welfare state.

Iversen and Cusack (2000) argue that it is not sufficient to show that international-price volatility is correlated with growth volatility and government spending. Rather, they claim it is necessary either that price volatility in international markets be greater than in domestic markets or that trade concentrates more than it diversifies economic risks. Iversen and Cusack then present evidence that, at least for advanced economies, there is no correlation between trade- or capital-market openness and volatility in output, earnings, or employment. They therefore dismiss both the argument that globalization increases economic insecurity and the claim that this connection leads to demands for welfare-state growth. Similar differences in methodology and conclusions to those between Rodrik and Iversen and Cusack are found in many of the contributors to this debate (e.g. Garrett 1998a, b; Burgoon 2001; Garrett and Mitchell 2001; Hayes, Ehrlich, and Peinhardt 2002; Swank 2002; McLaren and Newman 2002; Cameron and Kim 2001).

While this line of research has investigated reasonable hypotheses about how globalization may increase economic insecurity and thus demands for a robust welfare state, we have two broad concerns about this approach. The first is theoretical. It is actually *not* necessary for globalization to increase the magnitude of price and/or technology shocks for integration to increase individual economic insecurity in terms of riskier employment and/or wage outcomes. Thus, a lack of correlation between volatility in terms-of-trade and volatility in employment, wages, and output does not necessarily imply that globalization has not contributed to increased economic insecurity. In this paper, we present a simple model of a competitive labor market that clarifies this point by illustrating the key mechanism—labor-demand elasticities—through which we believe globalization may increase economic insecurity.

Our second concern is methodological. The theoretical connection between globalization and economic insecurity is an *individual-level* phenomenon, as is the subsequent connection to increased demands for a robust welfare state. All the empirical work we know of, however, employs aggregate country-level data to indirectly test this individual-level phenomenon. We know of no individual-level empirical study of whether exposure to the world economy increases worker insecurity.[2] Thus, a theoretically informed *individual-level* test of the prediction that globalization generates insecurity is lacking both in the political-economy literature on the forces underlying the backlash against globalization and in the literature on the connection between globalization and the welfare state.

In this paper, we present a theoretical model clarifying a critical mechanism through which globalization can increase individual economic insecurity. Drawing on standard frameworks of labor economics, we argue that FDI by multinational enterprises (MNEs) may be an important factor generating worker insecurity. FDI by MNEs may increase firms' elasticity of demand for labor. More-elastic labor demands, in turn, raise the volatility of wages and employment, all of which tends to make workers feel more insecure. It is important to note that this link from higher labor-demand elasticities to greater labor-market volatility does *not* require any change in aggregate shocks to the labor market: it holds even for some fixed amount of aggregate volatility.

Our emphasis on FDI as the key mechanism by which globalization generates economic insecurity is rare in the literature. The relative lack of attention on FDI is unfortunate because in

[2] This does not imply that the aggregate-level analyses and methodological debates are without merit. We only show that an increase in price volatility is not essential in order to argue that globalization increases worker insecurity. Nevertheless, it is not clear that there is a enough information in the aggregate data to make conclusive inferences about the correlations of interest. See, for example, Hayes, Ehrlich, and Peinhardt's (2002) demonstration of the sensitivity of estimates to changes in model specification to deal with selection problems in the aggregate analyses.

recent decades, cross-border flows of FDI have grown at much faster rates than have flows of goods and services or people.[3] Moreover, it is the multinationalization of production which a number of scholars have pointed to as the distinguishing feature of the current phase of globalization compared to previous episodes (Bordo, Eichengreen, and Irwin 1999). Finally, the lack of focus on FDI is surprising because, as we demonstrate, there are strong theoretical reasons to believe that, through its effect on labor-demand elasticities, it substantially influences economic insecurity among workers by increasing employment and earnings risks.

This theoretical framework then motivates our empirical test of the relationship between the multinationalization of production and the economic insecurity of workers. We present new evidence, based on the analysis of panel data from Great Britain collected from 1991-1999, that FDI activity in the industries in which individuals work is positively correlated with individual perceptions of economic insecurity. This relationship holds in yearly cross-sections, in a panel accounting for individual-specific effects, and in a dynamic panel model also accounting for individual-specific effects. Consequently, it is not only true that individuals more exposed to FDI activity are more likely to report greater insecurity. It is also the case that changes in exposure for a single individual, controlling for previous levels of insecurity, are correlated with changes in worker insecurity. We regard the individual-level panel results as the first valid evidence consistent with a causal relationship between globalization and worker insecurity.

There are four remaining sections to the paper. The next section provides a theoretical framework for the economics of worker insecurity. Section 3 describes the data to be used in the study and the econometric models to be estimated. Section 4 reports the empirical results and the final section concludes.

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[3] UNCTAD (2001) reports that from 1986 through 2000, worldwide cross-border outflows of FDI rose at an annualized rate of 26.2%, versus a rate of just 15.4% for worldwide exports of goods and services. In the second half of the 1990s this difference

2. *Theoretical Framework for Worker Insecurity, Labor-Demand Elasticities, and FDI*

2.1 Defining Worker Insecurity

Although there are a number of alternative definitions of economic insecurity, most often it is understood to be an individual's perception of the risk of economic misfortune (Dominitz and Manski 1997). Consequently, researchers have focused on the risk of events such as the loss of health insurance, being a victim of a burglary, losing a job, and significant decreases in wages. Some analysts have distinguished between *perceptions* of the risk of such events and the actual *anxiety and stress* caused by the risk (Anderson and Pontusson 2001, Gardner and Oswald 2001, Mughan and Lacy 2002).

For our research this distinction is very important because, as will be discussed below, our key measure of economic insecurity addresses most directly the anxiety/stress dimension. Consistent with many researchers in this area, we will assume that perceptions of risk do generate anxiety, and thus that our stress/anxiety measure is linked to the perceptions of economic misfortune. There are likely individual characteristics and environmental factors that influence this link (OECD 1997, Anderson and Pontusson 2001, Mughan and Lacy 2002), and one important task for our empirical analysis will be to address this.

It is likely that most people's perceptions of economic insecurity depend heavily on their purchasing power, which in turn depends on both their asset ownership and their labor-market status—both employment and income earned therefrom. In reality, the large majority of people rely much more on labor income than capital income for purchases; accordingly, we think labor-market status is the main determinant of perceptions of economic insecurity.

widened to 37.0% versus just 1.9%.

In light of this labor-market focus, we conjecture that the economic misfortunes underlying people's economic insecurity stem mainly from more-volatile employment and/or wage interactions with their employers. That is, risk-averse workers are not indifferent between employment options that yield the same amount of expected earnings but with differing degrees of certainty. More-certain earnings outcomes—due to more-certain wage and/or employment realizations—are preferred to less-certain ones, and insecurity rises with this uncertainty.[4]

2.2 Worker Insecurity in Labor-Market Equilibrium

Figure 1 visualizes equilibrium in a standard competitive labor market. The vertical axis plots wages, and the horizontal axis employment (measured in people or, if issues like overtime are thought to be important, hours).[5]

The labor-supply curve, LS, is aggregated across individuals, and is typically assumed to be upward sloping. At each point along the supply schedule, the elasticity of labor supply, η^S, is defined as the percentage change in the quantity of labor supplied by workers in response to a one-percent increase in the price of labor. An increase in wages is typically thought to generate both a substitution effect and an income effect among persons who work. Higher wages raise the opportunity cost of choosing leisure rather than work, and thus induce people to substitute towards more work. But higher wages also raise total income from the initial amount of work, and thus induce people to work less and choose more leisure. The substitution effect typically

[4] It is important to note that there is now a large body of evidence that labor-market volatility has been rising in many countries, especially in the 1990s, in terms of greater earnings volatility, declining job tenure, and self reports. Gottschalk and Moffitt (1994) report substantial increases in year-to-year earnings volatility for the United States over the 1970s and 1980s. Looking at the 1990s as well, a symposium issue of the *Journal of Labor Economics* (1999) documented declines in U.S job stability, especially in the 1990s for large groups of workers such as those with more tenure. Within that symposium issue, Schmidt's (1999) analysis of individual surveys finds that U.S. workers in the 1990s were more pessimistic about losing their jobs than they were during the 1980s. A wide range of surveys have found evidence of rising U.S. job insecurity over the 1990s relative to earlier decades, despite the ongoing economic expansion (e.g., Bronfenbrenner 2000).

[5] For a formal derivation of key labor-market concepts such as elasticities, see Hamermesh (1993). For discussion of labor-demand elasticities in general equilibrium trade models, see Reddy (2000).

dominates, and thus higher wages induce more work. Aggregated across workers, LS thus slopes up.

The labor-demand curve, LD, is aggregated across firms, and is typically assumed to be downward sloping. At each point along the demand schedule, the elasticity of labor demand, η^D, is defined as the percentage decline (in absolute value) in the quantity of labor demanded by that firm in response to a one-percent increase in the price of labor. This elasticity consists of two parts. One is the substitution effect. It tells, for a given level of output, how much the firm substitutes away from labor towards other factors of production when wages rise. The second is the scale effect. It tells how much labor demand changes after a wage change thanks to the change in the firm's output. Higher wages imply higher costs and thus, moving along the product-market demand schedule, lower firm output. When wages rise, both the substitution and scale effects reduce the quantity of labor demanded. The firm substitutes away from labor towards other factors, and with higher costs the firm produces less such that it demands less of all factors, including labor.

Labor-market equilibrium prevails at the intersection of LD and LS at point E, where the quantity of labor supplied equals the quantity of labor demanded. At that point, suppose that η^S and η^D are the relevant elasticities. We introduce volatility into the labor market by assuming that the position of the LD schedule is stochastic. This accords with a wide range of empirical evidence that labor-market volatility stems mainly from movements in LD rather than LS.

To see what forces drive volatility in LD, note that the labor demand schedule for each firm traces out the *marginal revenue product* of its workers as the wage rate varies. Each profit-maximizing firm hires workers until the wage paid to the last worker hired just equals the value of output—i.e., revenue—generated by that last worker. Vary the wage facing the firm, and the optimal number of workers to hire by this maximization rule varies. For each firm, the product prices and technology it faces are two key determinants of marginal revenue products.

Aggregated across all firms, then, the position of the LD schedule depends crucially on all relevant product prices and production technologies. Movements in prices and technologies trigger movements in LD and thus in equilibrium wages and/or employment. Define $\hat{mrp}$ as the percentage shift in the LD schedule due to shocks to prices and/or technologies. It is then straightforward to show that the resulting percentage change in equilibrium wages and employment are respectively given by $\hat{w} = \left(\frac{\eta^D}{\eta^S + \eta^D}\right)\hat{mrp}$ and $\hat{e} = \left(\frac{\eta^D \eta^S}{\eta^S + \eta^D}\right)\hat{mrp}$. Treating $\hat{mrp}$ as a random variable and the elasticities as parameters, then $Var(\hat{w}) = \left(\frac{\eta^D}{\eta^S + \eta^D}\right)^2 Var(\hat{mrp})$ and $Var(\hat{e}) = \left(\frac{\eta^D \eta^S}{\eta^S + \eta^D}\right)^2 Var(\hat{mrp})$.

For workers, the critical issue to note in the above expressions is that volatility in labor-market outcomes depends not just on the volatility of LD shifters such as product prices and production technology. It also depends on the magnitudes of the elasticities of labor supply and demand. If elasticities are assumed to be fixed, then greater labor-market volatility arises if and only if there is greater aggregate volatility in prices or technology. But this is not the only way to generate greater labor-market volatility. It can also be generated from increasing the elasticity of demand for labor, holding fixed the amount of aggregate risk. For some given values of $Var(\hat{mrp})$ and η^S, as η^D rises so too does $Var(\hat{w})$ and $Var(\hat{e})$. Higher labor-demand elasticities trigger more-volatile labor-market responses to price or technology shocks to labor demand.

This can be envisioned graphically by comparing the consequences for wages and employment of equal shifts in the labor demand curves LD and LD′ in Figure 1. In this figure, an increase in labor demand elasticity—perhaps induced by globalization as will be discussed below—would appear as a flattening of the labor demand schedule around the point E. LD′ is one such labor demand schedule. For an equal shift in LD and LD′ from a shock to prices and/or technology, there is a greater adjustment in wages and employment along the LD′ schedule characterized by more elastic demand.

For risk-averse workers, all this means that more-elastic labor demands should tend to raise economic insecurity. In the introduction, we highlighted the ongoing question about whether globalization increases aggregate risk—shocks to prices and/or technologies. Our simple framework here clarifies that for issues of worker insecurity, the answer need not be "yes." But in that case, the important question to ask becomes whether globalization makes labor demands more elastic.

2.3 How FDI by MNEs Can Make Labor Demands More Elastic: Theory and Evidence

Standard models in international trade predict that greater FDI by MNEs should make labor demands more elastic through both the scale and substitution effects. This should boost insecurity via the greater labor-market volatility just described. Consider each effect in turn.

Many models predict that FDI and its related international trade make product markets more competitive. Through the scale effect, this should make labor demands more elastic. For example, liberalization of FDI policies can force domestic firms to face heightened foreign competition. Or developments abroad related to MNEs (e.g., capital accumulation via FDI) can be communicated to domestic producers as more-intense foreign competition. In these cases more competitive product markets mean that a given increase in wages and thus costs translate into larger declines in output and thus demand for all factors. Different models predict different magnitudes of FDI and/or trade's impact on product-market demand.[6]

The second way through which FDI can increase labor-demand elasticities is through the substitution effect. Suppose that a firm is vertically integrated with a number of production stages. Stages can move abroad either within firms as multinationals establish foreign affiliates (e.g., Helpman 1984) or arm's length by importing the output of those stages from other firms (e.g., Feenstra and Hanson 1997). Globalization of production thus gives firms access to foreign factors of production as well as domestic ones, either directly through foreign affiliates or indirectly through intermediate inputs. This expands the set of factors firms can substitute towards in response to higher domestic wages beyond just domestic non-labor factors to include foreign factors as well. Thus, greater FDI raises labor-demand elasticities.

In the literature on globalization and labor markets, there are several recent studies indicating that MNEs and FDI influence labor-demand elasticities in the ways just discussed. Using industry-level data for U.S. manufacturing, Slaughter (2001) estimates that demand for production labor became more elastic from 1960 to the early 1990s, and that these increases were correlated with FDI outflows by U.S.-headquartered MNEs. Using plant-level data for all U.K. manufacturing from 1973 to 1992, Fabbri, Haskel, and Slaughter (2002) estimate that both U.K.-multinational plants and foreign-owned plants as well each had larger increases than did U.K. domestic plants in the elasticity of demand for production labor. They also estimate that increases were driven largely by greater substitutability between production labor and materials.

[6] One example is a monopolistically-competitive industry producing for Dixit-Stiglitz consumers who value product variety (e.g., Helpman and Krugman, 1985). Here the representative firm is usually assumed to face a demand elasticity (greater than one) that equals the elasticity of substitution (EOS) among product varieties in consumers' utility function. But the actual demand elasticity is only approximately equal to the EOS. It equals EOS plus a second term, $\frac{(1\text{-EOS})}{N}$, where N is the number of firms in the industry. As N rises—thanks, for example, to FDI by foreign MNEs—so, too, does this elasticity.

One important margin on which MNEs may affect elasticities is on the extensive margin of plant shutdowns. In response to wage increases, MNEs may be more likely than domestic firms to respond by closing entire plants. Evidence that multinational plants are more likely to close than are domestically owned plants has now been documented for the manufacturing sectors in at least three countries. For the United Kingdom, Fabbri, et al (2002) estimate that multinational plants—again, both U.K.- and foreign-owned—are more likely to shut down than domestic plants are (conditional on a set of operational advantages enjoyed by multinationals that make them less likely to shut down, like being older and larger). Gorg and Strobl (2002) find that foreign-owned plants in Irish manufacturing are more likely to exit. And for the United States, Bernard and Jensen (2002) report higher death probabilities for plants owned by firms that hold at least 10% of their assets outside the United States.

2.4 Summary of Theory Framework

To summarize, standard economic models of labor markets suggest that the globalization of production via MNEs may increase labor-demand elasticities. This, in turn, will tend to make labor-market outcomes more volatile and thus workers more insecure. This analysis suggests an empirical test of whether individual self-assessments of economic insecurity are related to FDI exposure in the labor market. The remainder of this paper turns to this empirical test.

Before doing that, we want to note one other important aspect of MNEs and labor markets. Several studies—of both developed and developing countries—have documented that establishments owned by MNEs pay *higher* wages than do domestically owned establishments. This is true even controlling for a wide range of observable worker and/or plant characteristics such as industry, region, and overall size. The magnitudes involved are big. Doms and Jensen (1998) document that for U.S. manufacturing plants in 1987, worker multinational wages exceeded domestically owned wages by a range of 5 to 15 percent, with larger differentials being enjoyed by production workers rather than non-production workers.[7]

What accounts for this "multinational wage premium" remains unknown, largely because this cross-sectional evidence is consistent with several alternative explanations, about which very little is currently known. The premium could be accounted for by higher worker productivity due to superior technology and/or capital at multinationals; or by higher worker productivity due to unobservable worker qualities; or by multinationals being more profitable and therefore able to share more rents with workers. Our framework in this section suggests another possibility: that MNEs pay more to compensate workers for the greater labor-market volatility associated with MNEs—e.g., for the greater risk of plant shut-downs. If workers for MNEs face a greater risk of job separation because MNEs have more elastic labor demands than purely domestic firms do, then to compensate they may receive higher wages.

Regardless of the cause(s) of the multinational wage premium, its existence is very important for considering how the globalization of production affects economic insecurity. All else equal, this premium is very likely to make multinational employees feel *more* economically secure. Our focus on elasticities and labor-market volatility highlights MNE influences on different dimensions of the overall worker-firm relationship. These contrasting issues of labor-demand elasticities and wage premia suggest that the net impact of MNEs on worker insecurity is *ex ante* unclear. Whether the wage premiums are sufficient to compensate workers for increases in risks from higher elasticities is an empirical question.

[7] Production workers receive an average of 6.9 percent less at comparable domestic plants employing more than 500 employees and 15.2 percent less at comparable domestic plants employing fewer than 500 employees. Non-production workers receive an average of 5.0 percent less at comparable domestic plants employing more than 500 employees and 9.5 percent less at comparable domestic plants employing fewer than 500 employees. For additional U.S. evidence see Howenstine and Zeile

3. Data Description and Empirical Specification

3.1 Data Description

The objective of our empirical work is to examine the impact of international capital mobility on economic insecurity. Specifically, we evaluate how individual self-assessments of economic insecurity correlate with the presence of highly mobile capital in the form of FDI in the industries in which individuals work. Our data cover Great Britain, which we think is an excellent case to examine both because inward and outward FDI have long figured prominently in the overall U.K. economy and because of the high quality of data available.

The individual data are from the *British Household Panel Survey* (BHPS) (2001). This study is a nationally-representative sample of more than 5,000 U.K. households and over 9,000 individuals surveyed annually from 1991 to 1999.[8] It records detailed information about each respondent's perceptions of economic insecurity, employment, wages, and many other characteristics. The most important pieces of survey information required for the analysis in this section is a measure of economic insecurity, identification of the respondents' industry of employment, and repeated measurement of the same individual over time.

We measure economic insecurity by responses to the following question asked in each of the nine years of the panel.

> "I'm going to read out a list of various aspects of jobs, and after each one I'd like you to tell me from this card which number best describes how satisfied or dissatisfied you are with that particular aspect of your own present job—job security."

The ordered responses are on a seven-point scale ranging from "not satisfied at all" to "completely satisfied." As discussed in the previous section, economic insecurity is most often

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(1994). Griffith (1999) presents similar evidence for the United Kingdom; Globerman, et al (1994) for Canada; Aitken et al (1996) for Mexico and Venezuela; and Te Velde and Morrissey (2001) for five African countries.

[8] The BHPS is ongoing but our data are through 1999 only.

in the literature understood either to be an individual's perception of the risk of economic misfortune and/or to be the anxiety or stress caused by the risk. We interpret our BHPS question as measuring the latter concept, as it has individuals report on their (dis)satisfaction with job security rather than on job security per se. We follow previous studies in assuming that perceptions of economic insecurity generate anxiety or lack of satisfaction, and thus that our BHPS question correlates with individual economic insecurity—albeit mediated by individual characteristics and environmental factors.

Using our BHPS question, we constructed the variable *Insecurity* by coding responses in the reverse order from the original question with a range from 1 for individuals who give the response "completely satisfied" to a 7 for those individuals giving the response "not satisfied at all." Higher values of *Insecurity* thus indicate less satisfaction with job security.

Our theoretical framework hypothesizes that high FDI activity in industries may generate economic insecurity among workers in those industries by increasing labor-demand elasticities. To test this hypothesis, we constructed the variable *FDI* to measure FDI exposure. We obtained data about inward and outward FDI investment positions in all 2-digit 1992 Standard Industry Classification (SIC92) industries for the U.K. from 1991 to 1999.[9] Since the BHPS records the industry the respondent is employed in according to the 1980 Standard Industry Classification (SIC80), we concorded the FDI data to 2-digit SIC80 industries.[10] We then merged the industry-level FDI data with the BHPS survey.

The variable *FDI* is a dichotomous industry-level variable that we set equal to one if two conditions were met: if the industry had any positive FDI investment, inward or outward, and if

[9] This data was obtained directly from the Office of National Statistics. We thank Simon Harrington for his assistance in generating this data.

[10] The BHPS records industry of employment according to the SIC80 classification scheme in all years but does report this information according to the SIC92 system in two years.

the industry's activities do not require producers and consumers to be in the same geographic location. If either of these conditions were not met, we coded *FDI* equal to zero.

Our logic in defining *FDI* with these two conditions runs as follows. The first condition for an individual's industry of employment to have positive FDI investment is straightforward. Any inward or outward FDI activity satisfies this.[11] The second condition recognizes that FDI activity is less likely to alter labor-demand elasticities if business activities cannot be outsourced across countries because the consumer and producer must be in the same geographic location. Consider the examples of wholesale trade, retail trade, and personal services (e.g., haircuts). The large majority of business activities in these industries require the co-location of producers and consumers: customers interacting with sales clerks, or sitting in the barber's chair. The notions of economic insecurity related to FDI that we discussed in Section 2 focus on the ability of MNEs to shift business activities across countries (i.e., on the substitution effect). In reality, in many industries, FDI does not have this characteristic; indeed, this FDI arises precisely because foreign customers cannot be served at a distance via international trade. Accordingly, our *FDI* variable identifies not all industries with FDI, but instead only those industries with FDI in which business activities can be outsourced across countries. So for industries such as wholesale trade, retail trade, and personal services we coded *FDI* as zero regardless of the data on actual FDI.

It is important to recognize the level of aggregation for the *FDI* regressor. The 2-digit SIC80 level we use is dictated by the FDI data available from the U.K. Office of National Statistics. Theoretically, we could imagine a more specific FDI exposure regressor that indicated FDI activity at the level of the respondent's company, rather than at the more-aggregated industry

[11] It is theoretically ambiguous if, in addition to the existence of FDI activity, the actual amount matters. Moreover, it is not clear, given that both inward and outward FDI must be measured, how to normalize the amount of activity across sectors. Thus, for both theoretical and empirical considerations, we employ a dichotomous measure. We discuss alternative measures of FDI exposure in Section 4.

level.[12] Our specification implicitly mixes the FDI activity of firms within each industry, and thereby assumes that within each industry the individual perceives any threat from FDI equally regardless of whether s/he works for a firm with foreign investment activity. This assumption seems reasonable. We are simply assuming that the labor-demand schedule faced by workers is set in the industry of employment rather than the firm or, for that matter, the entire economy.[13]

Given that our dependent variable measures the anxiety generated by economic insecurity, rather than that economic insecurity per se, it is critical that we control for differences among individuals in the link between the risk of economic misfortune and the stress caused by such risk. Previous research has suggested that there is systematic variation in perceptions of economic insecurity across demographic groups.

For our baseline cross-sectional analysis, we constructed four standard controls. The variable *Income* is equal to annual household income in thousands of U.K. pounds.[14] *Education* is a categorical variable ranging from one to four, with higher values indicating increasing educational attainment.[15] The variable *Gender* is equal to one for female respondents and zero for males. Finally, the variable *Age* equals the respondent's age in years at the time of survey.

Each of these control variables is likely to account for some of the differences among individuals in perceptions of economic insecurity in general, and in the link between the risk of job misfortune and the resulting anxiety generated in particular. However, it must be acknowledged that other unmeasured or unobservable differences among individuals may matter.

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[12] Of course, this is only a theoretical possibility. Even if we had firm-level FDI data, it would not be usable because the BHPS does not report the respondent's firm.

[13] Our focus on industries as the relevant aggregate for labor market effects is also consistent with many empirical findings in the labor-economics literature. For example, a common finding in studies of profit-sharing is that wage-bargaining keys off of industry profits above and beyond firm considerations. Of course, over longer time horizons than we consider in this paper, workers should be assumed to be facing an economy-wide labor demand curve.

[14] Annual household income is a variable calculated by the BHPS to include income from all sources in the twelve months prior to the September of the survey year as virtually all of the fieldwork for each survey year is done from September to December. The BHPS does impute some data in constructing this variable.

For example, individuals almost surely vary in their degree of risk aversion. In addition, individuals probably vary in their interpretation of the BHPS question. One individual may think about job security in compensated terms and assess satisfaction with job security conditional on their wages and any perceived compensating wage differential. Another individual in otherwise similar circumstances may respond without conditioning in this manner.

Unmeasured or unobservable individual heterogeneity is, of course, a problem that faces all survey research but it seems particularly acute in this analysis where our key variable to be explained is answers to a question that permits variation in interpretation. To address this heterogeneity, above and beyond our demographic controls we will exploit the panel structure of our data by including individual-specific effects for each respondent.

For each year of our panel, Table 1 reports summary statistics of our measure of economic insecurity and explanatory variables. The summary statistics and the analyses reported below are based on the BHPS sub-sample of private sector, full-time workers who are not self-employed. It is for this group of workers that our theoretical framework most directly applies; at the end of Section 4, we discuss the robustness of the results for larger samples. The average score on the insecurity scale is just below 3 in most years, suggesting that the average respondent was fairly satisfied with his or her job security.

FDI, our key explanatory variable, averages slightly over half in most years—i.e., slightly over half of respondents worked in FDI-exposed industries in most years. Industries that satisfied the two conditions for an FDI-exposed sector include metal manufacturing, mechanical engineering, and banking and finance. Among FDI-exposed industries in 1991, the sector with the most respondents was mechanical engineering. The industries meeting the two conditions for

[15] For example, category one indicates no qualifications or still in school and no qualifications, and category four includes teaching qualifications, first degree, or higher degree.

being FDI-exposed vary across each of the nine years of the panel, with sectors such as instrument engineering and business services being added to the list.[16]

3.2 Econometric Models

By matching each BHPS observation with the relevant industry FDI information, we examine how self-assessments of economic insecurity relate to FDI activity. Our starting point is to examine cross-sectional variation in economic insecurity for each year of the panel. This cross-section analysis can be formalized as follows,

$$Insecurity_i = \alpha + \beta FDI_i + \gamma Z_i + \varepsilon_i \quad (1)$$

where the subscript i indexes individuals; $Insecurity_i$ is our measure of economic insecurity; FDI_i is our measure of FDI presence; the vector Z_i is our set of control regressors discussed above; α, β, and γ are parameters to be estimated; and ε_i is an additive error term. We treat our measure of individual economic insecurity as a normally distributed random variable, and estimate the parameters in the equation using ordinary least squares regression with heteroskedastic consistent standard errors.[17]

[16] The two-digit 1980 SIC industries designated as FDI-exposed sectors vary for each of the nine years of our data. To get a sense of the variation over time, we list in this footnote the FDI sectors with BHPS respondents for the first and last years of our data. In 1991, the FDI-exposed sectors are agriculture & horticulture; coal extraction & manufacture of solid fuels; extraction of mineral oil & natural gas; metal manufacturing; chemical industry; production of man-made fibers; manufacture of metal goods not elsewhere specified; mechanical engineering; electrical & electronic engineering; manufacture of motor vehicles & parts thereof; manufacture of other transport equipment; food, drink & tobacco manufacturing industries; textile industry; manufacture of paper & paper products, printing and publishing; processing of rubber & plastics; postal service & telecommunications; banking & finance; insurance, except for compulsory social security. In 1999, the FDI-exposed sectors are agriculture & horticulture; extraction of mineral oil & natural gas; mineral oil processing; nuclear fuel production; production & distribution of electricity, gas, & other forms of energy; water supply industry; metal manufacturing; extraction of minerals not elsewhere specified; manufacture of non-metallic mineral products; chemical industry; manufacture of metal goods not elsewhere specified; mechanical engineering; manufacture of office machinery & data processing equipment; electrical & electronic engineering; manufacture of motor vehicles & parts thereof; manufacture of other transport equipment; instrument engineering; food, drink & tobacco manufacturing industries; textile industry; manufacture of leather & leather goods; footwear & clothing industries; timber & wooden furniture industries; manufacture of paper & paper products, printing and publishing; processing of rubber & plastics; other manufacturing industries; postal service & telecommunications; banking & finance; insurance, except for compulsory social security; business services; renting of movables; owning & dealing in real estate; education; research & development.

[17] All the cross-sectional results reported below are robust to alternative specifications in which *Insecurity* is treated as an ordinal categorical variable and ordered probits are estimated.

The coefficient estimates of β indicate whether and to what extent individual perceptions of economic insecurity are correlated with FDI activity. Exposure to FDI activity is increasing in the variable *FDI,* and we expect this may be positively correlated with the dependent variable *Insecurity*. This is the central hypothesis of our empirical work. Thus, our null hypothesis is that $\beta = 0$, with the alternative $\beta > 0$.

Estimating the effect of FDI on insecurity using period-by-period cross-sectional regressions is, however, inefficient because it fails to take advantage of all the information available in panel data sets (Wawro 2002, Yoon 2000). Pooling the data across the years of the panel has obvious advantages but generates a number of estimation issues regarding individual heterogeneity. Since the same respondents are surveyed over time, it is likely that observations for the same individual will be more similar than observations across different individuals. This might be in part because there is persistence in an individual's perceptions of economic insecurity, or because there are unmodeled characteristics about the individual that cause them to have similar perceptions in each period. This is particularly pertinent to our analysis because, as discussed above, there are good reasons to think that there are unobserved factors that may affect perceptions of economic insecurity. We can model this heterogeneity by revising the cross-sectional equation for the pooled data,

$$Insecurity_{it} = \alpha_i + \beta FDI_{it} + \gamma Z_{it} + \varepsilon_{it} \quad (2)$$

where the variables and parameters are the same as in Equation 1 but now each observation is indexed by *i* and *t,* for individuals and years. Further, α is allowed to vary across individuals to model unmeasured or unobserved heterogeneity across respondents, and *Z* now includes dichotomous indicator variables for each year of the survey.

Equation 2 can be estimated via random- or fixed-effects estimators. The random-effects estimator generates consistent parameter estimates if the individual effects are uncorrelated with the explanatory variables. The fixed-effects estimator is also consistent under this assumption, but is less efficient. Under the alternative hypothesis that the individual effects are correlated with the explanatory variables, only the fixed-effects estimator is consistent. We use both methods to estimate Equation 2, and report diagnostics to evaluate the estimators.

Although modeling individual-specific effects is one way of accounting for persistence in panel data, it does not allow us to differentiate between the idea that persistence in observations of insecurity are accounted for by the influence of past experiences of insecurity on present perceptions and the alternative that some types of individuals just have unobserved characteristics that lead them to have certain types of perceptions (Green and Yoon 2002, Wawro 2002). To make this assessment and to verify the robustness of our estimates of β, it is necessary to add a lag of the dependent variable to Equation 2. The final model we estimate is

$$Insecurity_{it} = \rho Insecurity_{it-1} + \alpha_i + \beta FDI_{it} + \gamma Z_{it} + \varepsilon_{it} \quad (3)$$

where ρ is the coefficient on the lag of the dependent variable.

This specification is a dynamic panel model. Introducing a lag of the dependent variable to the equation generates correlation between the individual-specific effects and the lag of the dependent variable. Consequently, this equation clearly cannot be estimated using random effects. Moreover, when the number of periods is small, as in our data, the fixed-effects estimator is also biased and inconsistent in the presence of a lagged dependent variable. Wawro (2002) reviews a number of alternative estimators for this situation, some of which first-difference Equation 3 to deal with individual-specific effects and then use instrumental variables to address the correlation between the error term and lagged dependent variable generated by the initial transformation of Equation 3. We use the Arellano-Bond generalized method-of-moments estimator, and report diagnostics to evaluate the assumptions required for its valid application.

4. Empirical Results

4.1 Baseline Specifications

Table 2 reports the results of our cross-sectional analysis. These results are ordinary least squares coefficient estimates for the parameters in Equation 1, with heteroskedastic consistent standard errors. The key finding is that FDI activity is positively correlated with individual economic insecurity. Holding other factors constant, individuals employed in FDI sectors systematically report less satisfaction with their job security. The coefficient estimate for the variable *FDI* ranges between 0.274 (with a standard error of 0.070) in 1994 to 0.397 (with a standard error of 0.071) in 1993. In every year, the estimated parameter is significantly different from zero at at least the 99% level. Although there is some variation across years in the size of the estimate, in most years it is very close to 0.30 and no trend is evident. Substantively, it generally has the largest effect of any of the regressors. We regard the cross-sectional estimates in Table 2 to be strongly consistent with the hypothesis that FDI activity generates economic insecurity among workers.

The results in Table 2 for the demographic control variables are also of interest. Older and more educated respondents are generally less satisfied with their job security than those who are younger and less educated. The education effect may be related to the “aspiration effect” documented in previous studies of general job satisfaction: more educated workers are thought to expect more from all aspects of their jobs, perhaps including job security. The results also indicate that women are more satisfied with their job security than men. This difference, while statistically significant in all years, declines in magnitude over time. Finally, the estimates in

Table 2 indicate an unstable relationship between household income and economic insecurity. The only statistically significant estimates are negative, consistent with wealthier households being able to self-insure against the risks of job separation and thus more satisfied with their job security. This result, however, holds in only half the years in the panel.

Despite the robustness of the correlation between *FDI* and our measure of economic insecurity, there are number of reasons to be concerned about the validity of these inferences. The period-by-period cross-sectional analysis is inefficient. Further, and more importantly, unmeasured and perhaps unobservable differences among individuals—such as variation in risk aversion—are likely correlated with both perceptions of economic insecurity and the propensity to be employed in a FDI exposed sector—correlations which would bias cross-sectional parameter estimates. These issues were discussed in detail in Section 3.

To address these concerns, we pooled the panel data sets and explicitly modeled individual-specific effects as in Equation 2. Table 3 reports the results of the random-effects and fixed-effects estimators of this equation.[18] For both estimators, the main substantive finding is, as in Table 2, a continued positive correlation between *FDI* and the dependent variable *Insecurity*. The magnitude of the estimated effect is over twice as large in the random effects specification. Both specifications include a full set of year indicator variables; the coefficients of which indicate whether mean levels of insecurity deviated in each year from the base year 1991. The parameter estimates are negative for every year except 1992, and turn significantly negative in both specifications after 1995. This indicates lower average levels of insecurity in later years. It

[18] In results not reported, we included the demographic control variables *Gender, Education*, *Age,* and *Income* in the random effects specification and *Education*, *Age,* and *Income* in the fixed effects model (*Gender* is time invariant so cannot be included in the fixed effects model). All the results for the *FDI* parameter are robust to retaining these variables. They were dropped because the parameters for these regressors are all not significantly different from zero in our preferred specification in the dynamic panel reported below.

is broadly consistent with the pattern of U.K. macroeconomic performance over the 1990s: initial recession followed by increasingly strong economic growth.

Although the main substantive story is the same across the two specifications in Table 3, it is still necessary to determine our relative confidence in the two estimators. We employed the Hausman specification test: if the random-effects assumption that the individual-specific effects are uncorrelated with the explanatory variables is true, then coefficient estimates from the two models should not be statistically different. The test statistic is χ^2 distributed with degrees of freedom equal to the number of coefficients (9 in our application) and is equal to 54.03. This rejects the null hypothesis that the coefficients do not differ statistically, and suggests violation of the key random-effects assumption. Consequently, the fixed-effects specification is preferred.

It is important to contrast the sources of variation in Tables 2 and 3 that are generating our main finding of a positive correlation between FDI presence and economic insecurity. The cross-section estimates of Table 2 exploit variation across individuals in their industry of employment and economic insecurity at a single point in time. In contrast, the panel estimates of Table 3 identify off of changes in FDI exposure over time. Individuals for whom there is no change in the FDI activity in their industry and who also do not change their industry of employment have their FDI-presence measure fully absorbed by their individual fixed effects. Variation across these individuals was used in Table 2 but is not in Table 3. Instead, identification in Table 3 comes from changes over time in individuals' self-assessments of economic insecurity that occur either with changes over time in FDI activity in individuals' industry of employment and/or with changes over time in individuals' industry of employment.

Table 4 reports the results of Equation 3, our application of the Arellano-Bond estimator that adds a lag of the dependent variable to our econometric model of economic insecurity. This addition is substantively of interest because it allows us to differentiate two ideas: the idea that persistence in observations of insecurity for any given respondent are accounted for by the influence of past experiences of insecurity on present perceptions, and the alternative that some types of individuals just have unobserved characteristics that lead them to have certain types of perceptions. Recall that the Arellano-Bond estimator purges individual-specific effects by first-differencing the data. It then uses instrumental variables to address the correlation between the error term and lagged dependent variable generated by the first-differencing.

In comparing the results in Table 4 with those in earlier tables, it is important to note that the number of individuals and total observations has significantly declined. First differencing and the use of lagged instruments results in the loss of the 1991 and 1992 data altogether. It also means that individuals must be retained in the panel for three years to be included in the analysis.

The estimate for the coefficient on the lagged dependent variable, ρ, is equal to 0.198 with a standard error of 0.021. This suggests that past shocks to individual perceptions of economic insecurity do affect current perceptions though the magnitude of this effect is not large. In this sample, persistence in individual economic insecurity depends both on individual-specific characteristics that make some individuals more likely to have particular perceptions and also on the effect of past perceptions of insecurity on those in the present.

The estimate of the coefficient β is 0.110 with a standard error of 0.049. To interpret, the long-run effect of FDI exposure on economic insecurity, it is necessary to divide this estimate by $1-\rho$ (i.e. 1-0.198). Consequently, the estimated impact of FDI exposure on economic insecurity is 0.137. The magnitude of this estimate is approximately the same as for the pooled fixed-effects estimator reported in Table 3, and it is statistically significant at the 0.05 level. We regard this to be a quite rigorous test of our central hypothesis. A significant correlation between exposure to

FDI and perceptions of economic insecurity remains conditional on our controls for individual heterogeneity, for the persistence of perceptions of economic insecurity, and for year-to-year shocks in insecurity.

To assess the validity of the results reported in Table 4, we conducted three diagnostic tests recommended by Arellano and Bond (1991). These results are reported in Table 5.[19] The consistency of their estimator requires that the errors, ε_{it}, in Equation 3 are serially uncorrelated. Arellano and Bond point out that if this is the case, then the first differenced residuals should display negative first-order serial correlation but not second-order serial correlation. Table 5 reports that we can reject the null hypothesis of no first-order serial correlation but cannot reject the null of no second-order serial correlation. Arellano and Bond also develop a Sargan test that helps further assess whether the assumptions about serial correlation hold. The null hypothesis of this test is that the model's overidentifying restrictions are valid; rejection of the null suggests the need to respecify the model (see Arellano and Bond, 1991, and Wawro, 2002). Table 5 reports that we do not have evidence to reject the null hypothesis that the overidentifying restrictions are valid. Overall, the three diagnostic tests reported in Table 5 do not raise significant concerns about the basic assumptions required for valid implementation of the Arellano-Bond estimation results reported in Table 4.

4.2 Robustness Checks

To verify our main findings in Tables 2 through 5, we conducted a number of robustness checks. One important issue is the possibility of estimation bias due to endogeneity and/or measurement error. On endogeneity, it might be argued that individual FDI exposure is not strictly exogenous because individuals may choose their industry of employment based (at least

[19] Following Arellano and Bond (1991), we report coefficient estimates based on their one-step estimator with robust standard errors and diagnostics based on their two-step estimator.

partly) on their perceptions of economic insecurity. Less-secure risk-averse workers might choose not to work in FDI-exposed sectors, while risk-loving workers might choose the opposite. On measurement error, it might be argued that our key *FDI* regressor is an imperfect measure of the underlying economic concept of interest, the labor-market riskiness linked to multinationals.

The endogeneity of worker industry choice is certainly a possibility. That said, in the panel a substantial proportion of changes over time in individual FDI exposure arise from changes in industry FDI status, which are clearly exogenous relative to individual perceptions, rather than from changes in individual industry of employment. Moreover, to the extent that people do switch industries endogenously, this should bias the coefficient on FDI exposure in the negative direction—away from the hypothesized positive effect. These considerations mean that our results reported in Tables 2 through 5 may underestimate the effect of FDI exposure on perceptions of economic insecurity.

Nonetheless, both to relax the strict exogeneity assumption and to address measurement concerns, we used the panel structure of the data by allowing previous errors—i.e., unforecasted realizations of *Insecurity*—to influence future changes in *FDI*. The model estimated is the same dynamic panel reported in Table 4, but now our FDI regressor is instrumented for using its lagged levels and changes in the same way that the Arellano-Bond estimator uses instruments for the lagged dependent variable. This approach accounts for potential endogeneity, and it also generates consistent estimates on *FDI* in the presence of random measurement error.

The results of this analysis are reported in the first column of Table 6. As expected, the coefficient estimate on *FDI* is substantially larger than before: the implied long-run effect is now 0.34, well over twice as large as the estimate in Table 4. Although the standard error is also now relatively larger, the coefficient estimate is still significant at the 0.15 level. One method for evaluating whether relaxing the strict exogeneity assumption is warranted is comparing the p-value of Sargan test in the two specifications. Since the p-values are essentially identical, there is little evidence that the specification in Table 6 is preferred. Based on our results in this table and our related discussions, we conclude that endogeneity and measurement error are not serious problems for the key result that FDI exposure influences perceptions of economic insecurity.[20]

Another important issue we considered was specification choice and potential omitted-variable bias. For example, all the empirical evidence in Section 2.3 on how multinationals influence labor-demand elasticities involves the manufacturing sector only. That suggests that the *FDI* regressor might actually be capturing something about manufacturing, not FDI. To test for this possibility we constructed and included in regressions the dichotomous variable *Manufacturing*, equal to one if an individual worked in a manufacturing industry and zero otherwise.[21] In both cross-section and pooled analyses, our FDI estimates were not substantially affected by including this variable (which usually had a positive and significant coefficient only in the cross sections). The second column of Table 6 reports one such pooled estimate.

A second specification check we tried was to add to the regressions reported in Tables 3, 4, and 6 our various demographic controls from Table 2. These had no effect on our key *FDI* regressor, and in our preferred dynamic panel specification, none of the estimated coefficients for the demographic variables are significantly different from zero. One final check we mention is sensitivity to estimation sample. Our core results are for a sample of private-sector, full-time, not-self-employed workers—the labor market participants for which the theoretical framework

[20] For measurement error, we also investigated whether small changes in the industry-by-industry coding of the *FDI* variable had any effect on its coefficient estimates. For example, we re-estimated all tables when agriculture and mining industries were not included as possible FDI-exposed industries. All changes we investigated had minimal impact on our core results.

[21] The one-digit SIC industries that constitute manufacturing are manufacture of metals, mineral products, and chemicals; metal goods, engineering, and vehicle industries; and other manufacturing industries.

most directly applies. In estimates of key specifications using broader BHPS samples, our FDI-insecurity correlation of interest maintained.

5. Conclusions

A central question in political and academic debates about international economic integration is whether globalization increases economic insecurity. In this paper, we present a theoretical framework clarifying a critical mechanism through which globalization can increase individual economic insecurity. Drawing on standard frameworks of labor economics, we argue that FDI by multinational enterprises (MNEs) may be an important factor generating worker insecurity. FDI by MNEs may increase firms' elasticity of demand for labor. More-elastic labor demands, in turn, raise the volatility of wages and employment, all of which tends to make workers feel more insecure. It is important to note that this link from higher labor-demand elasticities to greater labor-market volatility does *not* require any change in aggregate shocks to the labor market: it holds even for some fixed amount of aggregate volatility.

We then provide the first empirical test at the individual level of the relationship between the multinationalization of production and the economic insecurity of workers. Our analysis of panel data from Great Britain over the 1990s finds that FDI activity in the industries in which individuals work is positively correlated with individual perceptions of economic insecurity. This relationship holds in yearly cross-sections, in a panel accounting for individual-specific effects, and in a dynamic panel model also accounting for individual-specific effects. Consequently, it is not only true that individuals more exposed to FDI activity are more likely to report greater insecurity but also the case that changes in exposure for a single individual, controlling for previous levels of insecurity, are correlated with changes in worker insecurity.

We regard the individual-level panel results as the first valid evidence consistent with a causal relationship between globalization and worker insecurity.

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Figure 1: Labor Market Equilibrium

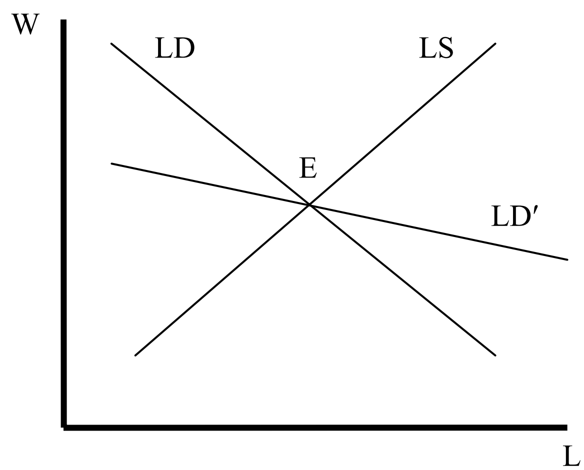

Table 1: Summary Statistics

Variable	1991	1992	1993	1994	1995	1996	1997	1998	1999
	Year								
Insecurity	2.978	3.026	2.902	2.917	2.881	2.789	2.682	2.663	2.726
	(1.982)	(1.747)	(1.663)	(1.701)	(1.642)	(1.563)	(1.532)	(1.465)	(1.579)
FDI	0.425	0.425	0.612	0.551	0.573	0.625	0.584	0.599	0.582
	(0.494)	(0.494)	(0.487)	(0.497)	(0.495)	(0.484)	(0.493)	(0.490)	(0.493)
Gender	0.349	0.353	0.363	0.369	0.352	0.356	0.349	0.352	0.346
	(0.477)	(0.478)	(0.481)	(0.483)	(0.478)	(0.479)	(0.477)	(0.478)	(0.476)
Education	2.262	2.325	2.399	2.437	2.468	2.511	2.540	2.558	2.539
	(0.897)	(0.893)	(0.898)	(0.910)	(0.905)	(0.901)	(0.884)	(0.870)	(0.877)
Age	35.459	35.563	35.425	35.447	35.644	35.550	35.508	35.809	36.111
	(12.017)	(11.719)	(11.572)	(11.574)	(11.566)	(11.527)	(11.723)	(11.885)	(11.718)
Income	23.778	25.278	25.902	26.486	27.804	29.319	29.666	30.572	30.721
	(13.564)	(14.126)	(13.596)	(14.564)	(15.789)	(16.417)	(17.260)	(20.565)	(22.782)
Observations	2,649	2,385	2,280	2,410	2,377	2,525	2,695	3,060	4,059

Notes: The BHPS sample in each year is private-sector, full-time workers who are not self-employed. Each cell reports the variable mean and, in parentheses, its standard deviation. *Insecurity* takes values from 1 to 7, with higher values indicating greater job insecurity. *FDI* is a dichotomous variable equal to one in industries with FDI presence as defined in the text. *Gender* is a dichotomous variable equal to one for females. *Education* takes values from 1 to 4, with higher values indicating more education. *Age* is age in years. *Income* is household income in thousands of pounds.

Table 2: Cross-Sectional Analysis of Economic Insecurity

Regressor	Year 1991	1992	1993	1994	1995	1996	1997	1998	1999
FDI	0.311	0.322	0.397	0.274	0.315	0.278	0.296	0.371	0.300
	(0.079)	(0.073)	(0.071)	(0.070)	(0.069)	(0.063)	(0.060)	(0.053)	(0.050)
Gender	-0.289	-0.334	-0.285	-0.336	-0.164	-0.158	-0.109	-0.176	-0.106
	(0.081)	(0.074)	(0.071)	(0.070)	(0.071)	(0.064)	(0.063)	(0.054)	(0.052)
Education	0.062	0.113	0.135	0.078	0.189	0.128	0.011	0.047	0.000
	(0.045)	(0.042)	(0.041)	(0.042)	(0.039)	(0.036)	(0.036)	(0.032)	(0.030)
Age	0.009	0.007	0.011	0.012	0.011	0.009	0.011	0.011	0.010
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.002)	(0.002)
Income	-0.001	0.000	-0.005	0.001	-0.005	0.000	-0.003	-0.002	-0.003
	(0.003)	(0.002)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)
Constant	2.519	2.497	2.174	2.230	2.031	2.027	2.232	2.059	2.318
	(0.186)	(0.175)	(0.168)	(0.165)	(0.160)	(0.152)	(0.150)	(0.135)	(0.127)
S.E.R.	1.967	1.726	1.636	1.679	1.619	1.548	1.519	1.444	1.566
Observations	2,649	2,385	2,280	2,410	2,377	2,525	2,695	3,060	4,059

Notes: These results are ordinary least squares regression coefficient estimates for each year for equation (1). Each cell reports the coefficient estimate and, in parentheses, its heteroskedastic-consistent standard error. For variable definitions, see the notes to Table 1.

Table 3: Panel Analysis of Economic Insecurity, 1991-1999

Regressor	Random Effects	Fixed Effects
FDI	0.234	0.105
	(0.024)	(0.032)
Year 1992	0.067	0.099
	(0.038)	(0.039)
Year 1993	-0.092	-0.027
	(0.039)	(0.041)
Year 1994	-0.069	-0.014
	(0.038)	(0.041)
Year 1995	-0.090	-0.020
	(0.039)	(0.042)
Year 1996	-0.196	-0.127
	(0.038)	(0.042)
Year 1997	-0.280	-0.205
	(0.038)	(0.041)
Year 1998	-0.295	-0.203
	(0.037)	(0.042)
Year 1999	-0.243	-0.174
	(0.036)	(0.042)
Constant	2.832	2.855
	(0.031)	(0.032)
Observations	24,636	24,636
Individuals	7,320	7,320
T	$1 \leq T \leq 9$	$1 \leq T \leq 9$

Notes: Each cell reports the coefficient estimate and, in parentheses, its standard error for equation (2). For variable definitions, see the notes to Table 1.

Table 4: Dynamic Panel Analysis of Economic Insecurity, 1993-1999

Regressor	Arellano-Bond
Δ *Insecurity*$_{(t-1)}$	0.198
	(0.021)
Δ *FDI*	0.110
	(0.049)
Δ *Year 1993*	-0.091
	(0.041)
Δ *Year 1994*	-0.043
	(0.042)
Δ *Year 1995*	-0.002
	(0.041)
Δ *Year 1996*	-0.088
	(0.039)
Δ *Year 1997*	-0.171
	(0.036)
Δ *Year 1998*	-0.093
	(0.033)
Constant	-0.025
	(0.009)
Observations	13,377
Individuals	3,781
T	$1 \leq T \leq 7$

Notes: Each cell reports the coefficient estimate and, in parentheses, its heteroskedastic-consistent standard error for equation (3). For variable definitions, see the notes to Table 1. The Arellano-Bond estimator is a first-difference estimator so the dependent variable is actually the difference between the *Insecurity* measure in period *t* and period *t-1*.The sample estimated in this table is two years shorter than in Table 3 because two lags are required to estimate the model.

Table 5: Specification Tests for Dynamic Panel Analysis of Economic Insecurity: 1993-1999

Test	Results
1st-order serial correlation test:	
Z	-20.43
Probability value under null of no autocorrelation	0.000
2nd-order serial correlation test:	
Z	1.15
Probability value under null of no autocorrelation	0.250
Sargan test of over-identifying restrictions:	
$\chi^2(27)$	29.45
Probability value under null that overidentifying restrictions are valid	0.339

Table 6: Robustness Checks for Pooled Analyses

Regressor	Arellano-Bond with Instruments for *FDI*	Arellano-Bond with Instruments for *FDI*
Δ *Insecurity*$_{(t-1)}$	0.230	0.230
	(0.016)	(0.016)
Δ *FDI*	0.260	0.289
	(0.179)	(0.198)
Δ *Manufacturing*		-0.078
		(0.146)
Δ *Year 1993*	-0.115	-0.120
	(0.051)	(0.053)
Δ *Year 1994*	-0.056	-0.059
	(0.046)	(0.046)
Δ *Year 1995*	-0.015	-0.016
	(0.043)	(0.043)
Δ *Year 1996*	-0.109	-0.111
	(0.043)	(0.043)
Δ *Year 1997*	-0.182	-0.181
	(0.037)	(0.037)
Δ *Year 1998*	-0.100	-0.100
	(0.034)	(0.034)
Constant	-0.033	-0.034
	(0.011)	(0.011)
Observations	13,377	13,377
Individuals	3,781	3,781
T	$1 \leq T \leq 7$	$1 \leq T \leq 7$

Notes: Each cell reports the coefficient estimate and, in parentheses, its heteroskedastic consistent standard error for equation (3) relaxing the assumption that the FDI exposure variable, *FDI,* is completely exogenous. In these specifications, unforecasted changes in *Insecurity* are allowed to affect future changes in FDI exposure. Instrumental variables from lagged levels and lagged changes in *FDI* are used in the same way that the Arellano-Bond estimator instruments for the lagged dependent variable. For variable definitions, see the notes to Table 1.